FINANCIAL LIBERALISATION AND CONSUMPTION BEHAVIOUR

Adrian Blundell-Wignall*
Frank Browne**
Stefano Cavaglia**
and
Alison Tarditi*

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*Economic Research Department
Reserve Bank of Australia

**Organisation For Economic Co-operation and Development

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ABSTRACT

The paper addresses the question of whether financial liberalisation and innovation have significantly altered consumption behaviour by reducing liquidity constraints as capital markets have become more flexible. A consumption model in which the permanent income hypothesis and extreme Keynesian consumption functions are nested as special cases is the starting point for this analysis. Estimated values for the sensitivity of consumption to current income for different time periods and for several OECD countries are assessed and compared in the light of various econometric properties, country-specific liberalisation measures and a variety of proxies reflecting changing liquidity constraints.

TABLE OF CONTENTS

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1. INTRODUCTION

The last ten to fifteen years have seen substantial deregulation in the financial sectors of most OECD economies. Early innovation in financial markets was motivated in part by the large incentives to circumvent official regulations. Subsequently, financial change has been facilitated by the progressive dismantling of these same regulations. While liberalisation policies have been motivated mainly by the desire to improve efficiency within the financial system, important implications for the functioning of the macro economy also arise.

One of these is the presumption that financial market liberalisation has eased liquidity constraints facing households and has therefore allowed consumption to be smoothed over time. If households have rational expectations and consume from permanent income (the RE-PIH hypothesis) then consumption can be shown to follow a random walk (see Hall (1978)). This model of consumption behaviour has been widely rejected in empirical tests, however. These rejections, based on the finding that consumption behaviour is excessively sensitive to current disposable income, are frequently attributed to the failure of one particular maintained hypothesis of the PIH, namely that individuals with access to perfect capital markets can borrow or lend at the same interest rate to smooth consumption over their lifespans. If financial liberalisation has rendered capital markets somewhat less imperfect, and if the above interpretation for the failure of the RE-PIH to be supported by the data is correct, we should observe that the sensitivity of consumption to current income has fallen over time and is lower in countries that deregulated earlier and more thoroughly.

In attempting to test this proposition, the interpretation of estimated sensitivity parameters is subject to some ambiguity. Several hypotheses, some maintained, underlie the random walk model of consumption, notably that households are not liquidity constrained and that they do not behave myopically. We attempt to determine which of these hypotheses has failed by examining the random walk model both across time and between countries.

The plan of the paper is as follows. Section 2 sets out the Hall model of consumption and discusses the reasons that have been offered for its rejection in empirical tests. Section 3 sets out modifications to the Hall model to reflect the fact that a certain proportion of households experience difficulties in borrowing against the collateral of future labour income, and hence may be unable to attain the optimal profile of consumption over time implied by the PIH. This section also reviews previous empirical work which used this model to interpret the importance of liquidity constraints. Difficulties experienced with previous econometric studies guide the approach adopted in Section 4, where a version of the standard model is used to test whether the sensitivity of consumption to income has declined in successive decades in countries where financial liberalisation has eased liquidity constraints over time. In Section 5, the importance of pooling to allow for cross-correlation of residuals between countries (thus permitting more efficient parameter estimates) is demonstrated. In Section 6, time-varying parameter estimates of countries that liberalised during the 1970s and 1980s are presented. Finally, some concluding remarks are made in Section 7.

2. THE HALL MODEL AND MODIFICATIONS

Over the last decade or so most research on aggregate consumption has taken the Hall (1978) model as a starting point of analysis. Hall argued that the PIH implies that consumption behaviour should obey the first-order conditions for life-time utility maximisation of a representative individual. He begins with the conventional consumer model of life-cycle consumption under uncertainty:

Maximise: $$E_t \sum_{\tau=0}^{T-t}(1+\delta)^{-\tau}U(c_{t+\tau}) \qquad (1)$$

Subject to a lifetime wealth constraint: $$A_t = \sum_{\tau=0}^{T-t}(1+r)^{-\tau}(c_{t+\tau} - \omega_{t+\tau}) \qquad (2)$$

where:

E_t is the expectations operator, conditional on all information available in time t;
δ is the rate of time preference;
$U(c_t)$ is the one-period utility function;
r is the real rate of return on assets, assumed to be constant over time;
A_t is the consumer's assets, excluding human capital;
T is the length of economic life;
c_t is consumption;
ω_t is labour income, assumed to be stochastic, which is the model's only source of uncertainty.

Intuitively, an individual consumer, seeking to maximise his utility, is faced with the decision of whether to consume today or at some time in the future. This decision will depend upon his rate of time preference, the opportunity cost of interest foregone on income consumed today, and his expectation of the utility he would derive from consuming this income in the future. This may be written algebraically as:

$$E_t[U(c_{t+1})] = [(1+\delta)/(1+r)]U(c_t) \qquad (3)$$

If one is prepared to maintain the following somewhat strong assumptions regarding individual consumers:

(i) they have identical, time-separable preferences with a quadratic representation for instantaneous utility:

$$U(c_t) = -(\alpha - c_t)^2 \tag{4}$$

(ii) they cannot die in debt;

(iii) they have access to perfect capital markets in which the constant real rate of interest is equal to the subjective rate of time discount; and

(iv) they form expectations of future income rationally,

the first-order condition for an optimum can be shown to be:

$$c_t = \{1-(1+\delta)/(1+r)\}\alpha + \{(1+\delta)/(1+r)\}c_{t-1} + e_{1t} \tag{5}$$

where:

e_{1t} is the error term and is uncorrelated with all variables known to the consumer at time t-1.

Under these conditions, consumption follows a random walk. The present level of consumption is the optimal forecast of its future level or, alternatively, changes in consumption are unforecastable.

Alternatively, if the utility function is a power function of the form:

$$\mu(c_t) = c_t^{1-\alpha} \tag{6}$$

the behaviour of consumers can be approximated by:

$$c_t/c_{t-1} = \theta + \left[\frac{1+r}{1+\delta}\right]^{1/\alpha} + e_t \tag{7}$$

where a drift term θ is included to represent the long-run rise in aggregate consumption.

If disposable income is assumed to be generated by a process of the form:

$$y^d_t = X_{t-1}\beta + u_t \tag{8}$$

where:

X_{t-1} is a set of variables known to the consumer at time t-1;
u_t is a white-noise expectational error;

then the coefficient on any variable belonging to X_{t-1} should not be significantly different from zero in a regression of consumption on a constant, its own first lag and X_{t-1} (equation (5)), or in a regression of the rate of growth of consumption against a constant term and X_{t-1} (equation (7)).

Using variants of the Hall model, several researchers have found that current aggregate consumption is significantly more sensitive to changes in current disposable income than the PIH predicts. This excess sensitivity is frequently rationalised as arising from the presence of liquidity constraints. In terms of the life-cycle model, individuals make labour supply and consumption decisions over a known lifetime. Income will typically fall short of desired consumption in youth, exceed it in middle age, and again fall short of it in retirement. With perfect capital markets, individuals should be able to smooth consumption relative to income by borrowing when they are young and lending in middle age. In the presence of liquidity constraints, however, consumption cannot be fully smoothed because, for example, households cannot borrow when they are young against their future labour income.

Clearly, a breakdown of one or other of the abovementioned maintained hypotheses underlying the derivation of the random walk model, such as perfect capital markets; rational expectations of future labour income; additive time-separable preferences; separability between consumption, leisure and other goods (Mankiw, Rotemberg and Summers (1985)); or constant real rates of interest and discount rates (Mankiw (1981)), could cause current income to be sensitive to current consumption. Furthermore, the pattern of consumption for non-durables will be affected if durables and non-durables are non-separable in consumption, and if durables are subject to gradual adjustment to optimal levels. Bernanke (1985) has suggested that the illusion of excess sensitivity could consequently be created by the failure to account properly for durables expenditures. Another factor which could potentially explain the observed "excess" sensitivity has been emphasised by Zeldes (1989a and b). If there is uncertainty about future labour income, then consumers will self-insure by engaging in precautionary savings. An increase in such uncertainty will increase savings and reduce consumption relative to income. In other words, relative to a world of certainty, individuals' current consumption is "too" low and expected consumption growth "too" high, again creating an impression of "excess" sensitivity.

3. CROSS-COUNTRY λ COMPARISONS AND THE DEGREE OF FINANCIAL LIBERALISATION

One procedure for evaluating the PIH is to postulate a general model within which both the PIH and the Keynesian model (emphasising current income) are nested as special cases (as in Flavin (1985), Delong and Summers (1986), Jappelli and Pagano (1989), Campbell and Mankiw (1989) and Bayoumi and Koujianou (1989)). The significance of transitory income in explaining consumption is assumed to be evidence against one or all of the hypotheses listed in Section 2. This paper focuses on the last two of these, namely, no liquidity constraints and rational expectations, as the assumptions most likely to be responsible for the empirical rejection of the random walk consumption model.[1]

[1] Augmenting the nested specification by including a variable which causes consumers to experience liquidity constraints may permit the data to reject one or

Following Hayashi (1982) and Delong and Summers (1986), Jappelli and Pagano (1989) modify Hall's model of consumption to allow for a proportion of the population to be liquidity constrained. A first group of consumers represent that proportion of the total population who behave according to the RE-PIH and exhibit little if any sensitivity of consumption to transitory income. Jappelli and Pagano (J and P) assume quadratic utility, that is, they adopt equation (5) as their starting point, such that:

$$c_{1t} = a_0 + a_1 c_{1t-1} + e_{1t} \tag{9}$$

where:

$a_0 = \{1-(1+\delta)/(1+r)\}\alpha$
$a_1 = (1+\delta)/(1+r)$

A second group of consumers represent that proportion, λ, of the total population who, perhaps because of liquidity constraints, consume all of their disposable income instead of obeying the RE-PIH:

$$c_{2t} = \lambda y^d{}_t \tag{10}$$

Total per capita consumption is then:

$$C_t = c_{1t} + c_{2t} \tag{11}$$

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other of these competing explanations. Flavin suggests that the unemployment rate is such a variable, since it can be interpreted as a proxy for the proportion of the population subject to liquidity constraints. Although a much greater fraction of the population than those who are unemployed probably experience liquidity constraints, nevertheless an increase in unemployment will increase the fraction of the population who are constrained to consume from current disposable resources. The full gamut of tests presented in this paper for a simpler model (focusing on current income) was also applied to the Flavin model. Since the overall conclusions based on the two models are essentially no different, it was decided not to work with the Flavin model, which proved somewhat unwieldly for examining some of the issues raised.

Simple algebraic manipulation of equations (9), (10) and (11) yields an expression incorporating a **non-linear** constraint on the coefficient a_1:

$$C_t = a_0 + a_1C_{t-1} + \lambda(y^d_t - a_1y^d_{t-1}) + e_{1t} \tag{12}$$

The coefficient λ measures the degree of excess sensitivity of consumption to income. The model implies that this coefficient can also be interpreted as the share of income accruing to consumers who do not behave according to the RE-PIH.

Equation (12) is estimated using the technique of non-linear instrumental variables (NLIV). This circumvents the inconsistency problems associated with using Ordinary Least Squares to estimate a system of equations exhibiting correlated errors (the transitory consumption error, e_t in equation (12), is likely to be correlated with u_t in the disposable income equation (8)). The variables in X_{t-1} of equation (8) are used as instruments. J and P use a constant term, a linear trend and the first lag of consumption, disposable income, government expenditure and exports. Estimated λ values may then be used to rank countries according to the degree of financial liberalisation, provided the estimates are based on the same definitions of regressors and instruments. Such comparisons also assume the degree of financial liberalisation has not been subject to change over time.

J and P estimate equation (12) on annual data for seven OECD countries. Their results are interesting in terms of the questions raised earlier. They find that their estimates of λ for each country have roughly the reverse ordering to various measures of consumer debt scaled by total consumption. This suggests that those countries with greater access to credit markets or, alternatively, those with a greater desire to use them, show a smaller proportion of the population that is liquidity constrained. Separate examination of a variety of demand side factors suggests that these do not explain the differences in λ values, leading the authors to conclude that capital market imperfections were the main factor.

These results are intuitively appealing. However, the approach appears to ignore the difficulties associated with estimating equation (12) in levels form; such estimation requires the assumption of trend stationarity of the regressors. This reservation with the model is echoed in Campbell and Mankiw (1991). The weight of evidence from other studies conducted over a reasonable time period is that at least some of J and P's regressors are non-stationary.[2] West (1988) argues that, under certain circumstances, models containing one non-stationary regressor may still produce consistent and asymptotically normal estimators, provided that the regressor also exhibits non-zero drift. A selection of tests implemented on extended sample periods for two countries studied by J and P could not reject zero drift in a number of the non-stationary variables (see Appendix). It is also worth noting that J and P's use of annual data implies a very small number of observations (12 in the case of Japan and around 20 for most of their countries). This is not an adequate sample period to estimate their model. Nor do annual observations, which generate artificially smooth time averages of quarterly data, capture the notion of transitory income necessary to test the PIH.

Attempts to replicate and update J and P's non-linear instrumental variables estimation for a selection of three of their countries (the United States, Japan and Italy) were not satisfactory. For the United States and Japan, inconsistencies in the data were identified as the chief source of difference between the findings reported here and those of J and P (see Appendix for detailed estimation results and comparisons):

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[2] Existing tests for non-stationarity cannot be applied to J and P's data as the tests' powers are particularly low over small samples containing less than 100 observations. Nevertheless, augmented Dickey-Fuller and Phillips-Perron tests are applied to extended sample periods for two of the seven countries investigated by J and P. The results are presented in the Appendix. The small sample periods over which the non-linear estimation is performed (e.g. 12 observations for Japan, and an average of about 20 for other countries) poses problems for these tests of the individual variables, and certainly for the adequate estimation of J and P's model.

(i) J and P define their consumption variable as total private consumption "...excluding expenditure on durables, defined as the sum of appliances, furniture and means of transportation" (p1102). This definition is applied to each of the countries in their sample with the exception of the United States. Their US data appear to exclude consumption expenditure on non-durable services. Correcting for this, the value of λ rises from 0.21 to 0.42, altering the ranking of the US from the second least liquidity-constrained country behind Sweden to the fourth (behind Sweden, Japan and the United Kingdom). This result throws the US out of line with their ranking in terms of debt to consumption. Furthermore, updating the sample period for the US from J and P's 1961-1984 to 1961-1990 leads to a rise in the λ parameter from 0.42 to 0.51. Since the period 1985 to 1990 encompassed increasing liberalisation and household borrowing, one might have expected the values of λ to fall.

(ii) In the case of Japan, J and P's results could only be replicated when national rather than personal disposable income, and total government outlays rather than the national accounts measure of government expenditure were used. Consequently, estimates for Japan using definitions consistent with those for the United States increased J and P's reported λ value of 0.34 to 0.51. This affects their ranking of Japan (making it the fourth rather than the third least liquidity-constrained country).

It did not prove possible to replicate J and P's results for Italy using the data sources referenced.

The sensitivity of the estimates of λ to precise definitions of the variables used for consumption, income and instruments, together with the above-mentioned econometric problems associated with using non-stationary variables and a small number of observations, led us to downplay the findings of J and P. This point can be illustrated intuitively with the specific country example of Sweden. Both J and P and Campbell and Mankiw (C and M) (1991) rank Sweden as the least liquidity-constrained country amongst those studied. This does not sit well with the fact that the radical liberalisation of Sweden's financial markets was delayed until the second half of the 1980s (certainly outside of J and P's sample period), suggesting that other country-specific factors might have influenced the estimate of λ (for example, unique social security arrangements).

Similarly, important changes in the degree of financial liberalisation occurred in a number of the countries studied by J and P and C and M in the second half of the 1970s and throughout the 1980s. This is ignored by J and P and is touched on relatively briefly by C and M. The latter test whether λ has been a linear function of time, or was subject to a structural break in 1980, and find little (though not zero) support for either proposition. This dismissal of time variation in the responsiveness of consumption to income could, however, be due to the very specific nature of the tests carried out by C and M. Thus λ is unlikely to have been a linear function of time prior to the major deregulatory moves. With regard to their dummy variable test, C and M do not permit the drift parameter to change, and they impose the value of the co-efficient used to weight current and lagged changes in income. By forcing the intercept coefficient to be the same in both sub-periods, their estimates of λ may be biased.

For these reasons, and given the difficulties with cross-country comparisons, further exploration of the extent to which λ may have varied over time in response to financial liberalisation measures seems warranted.

4. ESTIMATES OF λ BY DECADE AND THE LIQUIDITY CONSTRAINTS INTERPRETATION

The following steps were taken to overcome some of the concerns that arose in applying the J and P approach. First, utility was assumed to be represented as in equation (6), so that non-liquidity constrained consumers behave according to equation (7):

$$c_{1t}/c_{1t-1} = b_0 + e_{2t} \tag{13}$$

where $b_0 = \theta + \left[\frac{1+r}{1+\delta}\right]^{1/\alpha}$

This avoids the non-stationarity problems discussed in the previous section. Second, it was assumed that current consumption of liquidity-constrained consumers is a constant fraction of current income -- so the expected rate of growth of consumption equals that of disposable income. Even if liquidity constrained, it seems unreasonable to assume that Keynesian consumers always spend exactly all of their income. This implies:

$$c_{2t}/c_{2t-1} = y^d{}_t/y^d{}_{t-1} + e_{3t} \tag{14}$$

Equations (13) and (14) are weighted according to the proportion of the population λ that are liquidity constrained:

$$c_t/c_{t-1} = \lambda\ (y^d{}_t/y^d{}_{t-1}) + (1\text{-}\lambda)\ b_0 + e_{4t} \tag{15}$$

where $e_{4t} = \lambda e_{3t} + (1\text{-}\lambda)\ e_2 t$

Third, quarterly time series of all variables are employed. Fourth, common consumption definitions are utilised to facilitate international comparisons. Equation (15) was estimated for total consumption for eight OECD countries.[3] It was also estimated for total consumption less the purchases of durables for that subset of countries for which such data are available.[4]

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[3] The countries in question are the United States, Japan, Germany, France, Italy, the United Kingdom, Canada, and Australia. Some of the results reported in Sections 4 and 5 were just reported in an early working paper version by Blundell-Wignall, Browne and Cavaglia (1991).

[4] These countries are the United States, Japan, France, Italy, the United Kingdom, and Canada. The results for total consumption less purchases of durables do not yield overall inferences which are sufficiently different from those for total consumption to warrant reporting, perhaps because of the emphasis on the time

Instrumental variables estimation was used to examine whether the degree of excess consumption sensitivity (which is measured by the λ parameter in equation (15)) is linked to imperfect capital markets.[5] This linkage is examined mainly by exploiting the knowledge that financial liberalisation has progressed over time, particularly in the 1970s and 1980s. If the liberalisation process seen in financial markets has caused liquidity constraints to be progressively relaxed, then estimating equation (15) for successive time periods should tend to indicate a generally reduced size and significance of the λ parameter. Three time periods are chosen, the 1960s, the 1970s and the 1980s (see Table 1 for precise dates). These, of course, may not be the economically most relevant time periods, given the gradual and diffuse nature of deregulation which makes it difficult to identify clear structural shifts. Nevertheless, with this approach less emphasis is placed on the cross-country ranking of λ than in the J and P paper. Formal tests can be applied to differences in λ values for individual countries between decades.

(a) Nested Model Estimates for Consumption

Table 1(p27) presents the results from estimating equation (15), in its change in logarithms form, for eight large OECD countries. Overall, the results suggest the excess sensitivity parameter λ varies across countries, as does its pattern over time. The United States, Japan, Canada and Australia show evidence of declining liquidity constraints. For the United States economy, there is no evidence that liquidity constraints were lessened in the 1970s compared to the 1960s. However, the λ parameter falls from a significant value of 0.47 in the 1970s to a statistically insignificant value of 0.25 in the 1980s, which is

variation of λ. These results are available, however, on request. It is also worth recalling that use of total consumption is less likely to create the illusion of excess sensitivity due to the stock adjustment of durables (Bernanke 1985).

[5] The instruments actually employed for the lag level of disposable income are three lags of this variable itself, as well as three lags on the lag level of personal consumption, the unemployment rate, and total exports, all in per capita terms, as well as contemporaneous population and a time trend.

consistent with reduced liquidity constraints. For Japan, the magnitudes of the estimated λ parameter suggest that liquidity constraints in the 1980s are less severe than they were in either the 1960s or 1970s. The size and pattern of estimated coefficients for Canada and Australia are quite similar. The λ estimate for Australia is significant in all three decades, but it declines in value during the 1970s and 1980s. The pattern for Canada conforms closely to expectations, since Canada was one of the first countries in the sample to deregulate its financial markets.

The results for Germany, France and Italy are similar in that they show no evidence of declining liquidity constraints. For Germany, the λ parameter is significant at the 5 per cent level for all three subsample periods, and increases in value from 0.37 in the 1960s to 0.67 in the 1970s and to 0.98 in the 1980s. Given that German households are known to have a strong preference for saving, increasing precautionary saving behaviour in the more uncertain environments of the 1970s and 1980s, together with the presence of liquidity constraints, might explain our results. For France, a significant excess sensitivity parameter (λ=0.40) is estimated when the extended period of the 1960s and 1970s together is used (see final column of Table 1). The robust errors estimate of the 0.31 coefficient for France in the 1980s suggests little easing of liquidity constraints in this period. Similarly, the Italian results show little change in the estimated values of λ (within the range of 0.43 to 0.47) between decades.

The United Kingdom results appear to have little in common with the two groups of countries discussed so far. The PIH appears to be accepted by the data in the 1960s and 1970s, but the size and significance of the λ parameter both increase during the 1980s. Possible reasons for this are discussed in Section 6.

Campbell and Mankiw (1989) observe that an estimate of the excess sensitivity parameter close to zero, even if statistically significant, supports the PIH, while a large value of this coefficient suggests rejection. According to this criterion, the pattern of the λ parameters across decades generally suggests easing liquidity constraints in four of the countries considered, but not in Germany, France, Italy and the United Kingdom. However, it is possible to test more rigorously whether declines in the λ parameter in successive time periods are significant in a statistical sense. The results, based on the unit normal distribution, are given in Table 2. The null hypothesis is that the λ value in the most recent period is smaller than in the preceding period. These results are a strong qualification to any interpretation based on the Campbell and Mankiw criterion. Only in a few instances (indicated by asterisks) does this test accept the null hypothesis of falling liquidity constraints.

(b) Asymmetries in the Nested Model

Implicit in the above tests is the proposition that liquidity constraints encountered will have the same effect regardless of the direction of movement in disposable income. This assumption may be unrealistic, since reductions in income are likely to be more constraining than increases -- at least for large changes in income. If the consumer is rationed in credit markets and current income falls substantially then consumption **must** fall. If income increases under the same credit market conditions, consumption may or may not increase. This asymmetry may be important for individuals with little or no non-human wealth, for whom necessary expenditure is a very large percentage of disposable income. Therefore, the presence of this type of asymmetry can be taken as further evidence of liquidity constraints. Moreover, as these constraints unwind over time with financial liberalisation, so should the **magnitude** of the asymmetry.

To test this proposition, the current values of the change in disposable income are transformed as follows:

$$\Delta \ln \hat{y}_t^+ = \Delta \ln \hat{y}^d{}_t \text{ if } \ln \hat{y}^d{}_t > \ln y^d{}_{t-1}$$
$$\text{zero, otherwise}$$

$$\Delta \ln \hat{y}_t^- = \Delta \ln \hat{y}^d{}_t \text{ if } \ln \hat{y}^d{}_t < \ln y^d{}_{t-1}$$

zero, otherwise

The hats indicate instruments obtained on y^d using the exogenous variables already noted (see footnote (5)). The test equation (15) is now rewritten as follows:

$$\Delta \ln C_t = \mu' + \lambda_1 \Delta \ln \hat{y}_t^+ + \lambda_2 \Delta \ln \hat{y}_t^- + (1-\lambda)\omega_t \tag{16}$$

Tests for asymmetric behaviour based on equation (16) are carried out with a simple t-test. The null hypothesis is assumed to be $\lambda_2 > \lambda_1$ and the alternative that $\lambda_2 \leq \lambda_1$. The results are presented in Table 3.

The null hypothesis of significant asymmetries is accepted at the 5 per cent level for the United States, Germany and Canada for the 1960s; Japan, Italy and Canada for the 1970s; and France, Italy and Australia in the 1980s. Furthermore, when asymmetry was accepted, the magnitude of the response of consumption to a fall in income was, in most cases, several multiples of the effect of an equivalent increase in income. In no case was the effect on consumption of an increase in income significantly greater than that for a fall. These results lend further support to the liquidity constraints interpretation of the excess sensitivity parameter. They also further corroborate the evidence already reported of easing liquidity constraints over successive decades. This conclusion is inferred on the basis of the magnitude of the average differences in the λ_1 and λ_2 estimates in successive subperiods (2.27, 0.29 and 0.22 for the 1960s, 1970s and 1980s, respectively[6] (see the bottom panel of Table 3)).

(c) Real and Nominal Interest Rate Effects

An attempt to add more realism to the nested consumption model in equation (15) is made by relaxing the assumption that the real interest rate is constant. This allows for changing intertemporal substitution in

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[6] These apparently systematic patterns may be biased by the very large λ_2 estimate for Italy for the 1960s. If, in fact, there were very few observations on Δy^- it might be operating as a dummy variable picking up some other influence.

consumption. The exclusion of the real interest rate from the test equation could conceivably be leaving real disposable income growth to pick up this changing intertemporal substitution effect, i.e. if real income growth and real interest rates are correlated over time. Nominal interest rate changes, in the absence of real interest rate changes, should of course have no effect on consumption unless households are prevented from consuming from their permanent income by imperfect capital markets. It has been argued that such an imperfection in personal credit markets is reflected in banks' practice of using virtually constant repayment-to-current income ceilings as criteria for loan qualification.[7] An increase in the nominal rate of interest, for a fixed real rate, may cause this ceiling to be breached and the potential borrower to be denied a loan. Thus if the liquidity constraints theory is valid, consumption can be constrained by variations in both current disposable income and in nominal borrowing rates of interest. Unless nominal interest rate effects are controlled for, variations in the imperfectly measured real rates could capture liquidity constraint effects coming from changes in nominal rates.

To capture both intertemporal substitutions and liquidity constraints phenomena arising from interest rate changes, equation (15) is accordingly amended as follows:

$$\Delta \ln C_t = \mu + \lambda \Delta \ln \hat{y}^d{}_t + \alpha r_{t-1} + \gamma \Delta i_t + \omega_t \tag{17}$$

An increase in last period's real borrowing rate of interest r_{t-1} reduces that period's consumption relative to that of the current period ($\alpha>0$). An increase in the nominal interest rate i_t in this period for a fixed real rate will tighten household liquidity constraints and dampen consumption expenditure ($\gamma<0$) if capital markets are imperfect.

7 Wilcox (1989) refers to a recent American Bankers Association textbook on consumer lending which suggests that a borrower's capacity to repay a loan can be measured by the payment-to-income ratio. Wilcox concludes that, in practice, this means the current payment-to-current income ratio.

The cross-country results from estimating equation (17) are given in Table 4. They indicate support for significant changing intertemporal substitution effects for the United States (for the 1960s/1970s subperiod and also for the 1980s) and for the United Kingdom (for the 1970s and the 1960s/1970s subperiods). There is also some weaker evidence for Japan favouring changing intertemporal substitution effects. Note, however, that for Italy in the 1960s this effect was significantly negative. These results are of some interest in the light of the failure of many recent empirical studies to uncover a significant positive intertemporal elasticity of substitution (see, for example, Mankiw, Rotemberg and Summers (1985), Hall (1978), Campbell and Mankiw (1989) and Bayoumi and Koujianou (1989)). Nominal interest rate effects arising from liquidity constraints are significant for Japan (1960s/1970s), for Germany (1980s), for France (1970s), and for Italy and Canada (1960s). However, for the United States, nominal interest rate increases appear to have promoted consumption in the 1980s.

There is, therefore, some evidence of both intertemporal substitution effects and liquidity constraint effects arising from real and nominal interest rate changes, respectively. However, in no instance does the presence of these effects require any substantive amendment of the conclusions already arrived at with respect to the changing pattern of liquidity constraints across countries or over time. Hence liquidity constrained consumers cannot be explained away by allowing households to redistribute their consumption over time in response to changes in the intertemporal relative price.

5. CROSS-COUNTRY CORRELATION OF RESIDUALS AND POOLED RESULTS

A worrying aspect of the Table 1 results is the general inability to reject the null hypothesis that the declines in λ are significant according to the unit normal test. This could be due to the use of an inefficient estimation procedure. An econometric issue not addressed in most of the previous literature concerns the possible importance of cross correlations between the error terms in the test equations for the countries being studied. Thus, for example, income and consumption shocks in one country could be translated through standard international linkages to income and consumption shocks in others. Alternatively, common shocks (e.g. oil price changes) could be important. It is possible to take account of this problem by pooling the data for a number of countries and using a Seemingly Unrelated Regression Estimation (SURE) procedure to estimate individual country parameters. This should enable more efficient estimates of the λ parameter.

More efficient estimates might also be obtained if the λ parameter can be estimated jointly for a number of countries where financial deregulation is thought to have been broadly similar, provided such a restriction is accepted by the data. This requires the countries considered suitable for pooling to be chosen on *a priori* grounds, i.e. on the basis of what is known about the deregulatory policies in each.

(a) Country Groupings Based on Information about Deregulation

Financial regulations generally fall into two broad categories:

(i) "rate/quantity" regulations on bank deposits and loans, including ceilings on bank deposit rates and quantitative measures that have similar effects (credit ceilings, capital controls, etc.); and

(ii) "powers" regulations governing the extensiveness of activities of individual financial institutions and their competitiveness.

In general terms, there are considerable differences in emphasis between countries in the extent to which "rate/quantity" regulations have been removed and/or "powers" regulations still apply. Developments are summarised in Table 5. The United States, the United Kingdom, Canada and Australia moved relatively early and with some rapidity in removing rate/quantity and powers regulations. While some powers regulations still apply, their financial system may be described as highly competitive. Japan too has made important steps in the 1980s, removing capital controls at the beginning of the decade, and gradually introducing market alternatives to regulated bank deposits throughout the decade. Developments proceeded more cautiously in France and Italy, with capital controls being removed only gradually throughout the 1980s and rate/quantity and powers regulations still applying fairly extensively over the full sample period used here. While Germany was one of the first countries to remove rate/quantity regulations in the 1960s and 1970s, it has been relatively slow to implement "powers" deregulation. As a result, competition between German banks has remained muted, and short-term financial instruments paying market returns have not been readily available as alternatives to bank deposits throughout the 1970s and 1980s.

On this basis our sample of countries can be divided into two groups.[8] In the first group, the United States, Japan, the United Kingdom, Canada and Australia are classified as countries that have implemented substantial liberalisation policies. The second group consists of the continental European countries, Germany, France and Italy, which have been much slower to deregulate. Having decided on this separation of countries, the test equations can be estimated for each group of countries as a system using a SURE technique and the instrumental variables for income already described above.

(b) Empirical Results

The results of this estimation procedure for group 1 are displayed in Table 6. No cross-equation parameter constraints are imposed in the top panel, and the results differ from those in Table 1 only to the extent that they take into account possible contemporaneous cross-correlation of residuals. Panel 2 of the table displays the results that emerge when the excess sensitivity parameter is constrained to be the same across countries in each subperiod. The validity of this constraint is tested using a likelihood ratio test. Finally, the joint constraint of equal slopes and intercepts is imposed across countries for each subperiod and again

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[8] See Blundell-Wignall, Browne and Manasse (1990) and further references therein.

the validity of this constraint is tested using the chi-squared test based on likelihood ratios. This result is reported in panel 3 of the table. The United Kingdom is excluded because the data did not accept the restriction that its λ value was the same as the other countries in the group.

For the other countries in group 1, the excess sensitivity parameters have changed somewhat in value relative to the Table 1 estimates and, as expected, the corresponding sample standard errors have fallen in all cases. The results shown in the top panel show declining λ values in the 1980s compared to either the 1970s or the 1960s for all countries. The revised results for the unit normal tests are shown in Table 7. In contrast to the earlier results shown in Table 2, the decline in λ in the 1980s compared to either the 1970s or to the 1960s is significant for both the United States and Japan. In the case of both Canada and Australia, the value of λ is significantly lower in the 1980s compared to the 1960s.

While the constraint that λ be identical across countries cannot be rejected for any subperiod, the additional constraint that the drift parameter also be the same across countries is rejected for two of the four subperiods. Applying the unit normal tests to the jointly estimated λ values in panel 2 of Table 6, significant liquidity constraint relaxation for this group of countries cannot be rejected for the 1980s compared to the 1960s (at the 5 per cent level) nor for the 1980s relative to the 1970s (at the 10 per cent level). However, no significant reduction in liquidity constraints is indicated for the 1970s compared to the 1960s, despite a substantial fall in the magnitude for the group excess sensitivity parameter. Abstracting from issues of statistical significance, and focusing on the magnitude of the common λ estimates for Group 1 (panel 2), the results say that the number of households which experienced liquidity constraints fell from 38 per cent in the 1960s to 29 per cent in the 1970s and to 14 per cent in the 1980s.

For the second group of countries (Germany, France and Italy), shown in Table 8, the pooled individual country results are broadly similar to those in Table 1. However, the likelihood ratio test rejected the imposition of a common λ across countries for the 1980s. None of the remaining constraints can be rejected at the 5, or even 10 per cent level of significance. Therefore, applying the normal tests to examine the significance of changes in the cross-country constrained λ for the 1980s compared to the other subsamples is clearly invalid. The 1970s common λ value is fractionally higher though not significantly different from that of the 1960s.

6. TIME-VARYING PARAMETER ESTIMATES

To further explore the value of λ over time, the instrumental variables equation (15) is estimated as a rolling regression. The initial sample period is chosen to reproduce the value of λ for the 1960s reported in Table 1. Its value is then allowed to vary by adding an observation, while keeping the total number of data points in each regression unchanged (ie dropping the earliest observation from the previous regression).[9] Only the countries in which financial liberalisation was thought to be important over the sample period are considered here, including the United Kingdom, which was excluded for technical reasons from the SURE procedure. The results, shown in Chart 1, permit further interpretation of the earlier findings.

What is particularly striking is the apparent correlation of changes in λ with factors other than financial liberalisation which influenced liquidity constraints prior to the 1980s. For all five countries, the excess sensitivity parameter declines in the late 1960s and/or the early 1970s (though much less so in the case of Australia). This corresponds with the easing of monetary policy at the time, which saw liquidity in the form of money balances expand rapidly. In the case of the United

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[9] Note that this does not necessarily reproduce the Table 1 1980s λ estimates as the last observation of λ in the graph. This is because the length of the "1960s" sample period differs (depending on data availability) for each country.

Kingdom, there was also a credit explosion in the wake of the introduction of Competition and Credit Control in 1971. The ready availability of money balances reduced liquidity constraints independently of the degree of financial regulation. However, the first oil shock in 1973 and 1974, and the firming of monetary policy at the same time, appears to have reversed these developments. The well-known transmission of a world-wide downturn in activity, in the presence of regulated capital markets, saw a marked rise in the dependence of consumption on current income - a phenomenon probably associated with increased precautionary saving behaviour.

These common patterns in response to shocks, it is worth noting, are consistent with the rationale given for the importance attached to the pooled (SURE) results presented in Section 5.

During the second half of the 1970s and/or throughout the 1980s, financial liberalisation became much more widespread in all of the countries considered here (Table 5). Moreover, following the inflation and income shocks of the 1970s, financial innovations to avoid existing regulations had in any case become more widespread. In the case of the United States, the Volker disinflation from 1979 to 1981 was associated with some rise in the excess sensitivity parameter. But following major deregulatory moves in the early 1980s, the sensitivity of consumption to current income appears to have moved into a phase of a sustained decline, despite major changes to the stance of monetary policy, the stock market crash and other shocks. In the cases of Japan, Australia and Canada there are also sustained declines from either the late 1970s or early 1980s, in spite of major nominal and real shocks during the 1980s.

Only the case of the United Kingdom presents something of a puzzle. From the mid 1970s to the mid 1980s, there is a sustained decline in the excess sensitivity parameter, in much the same way as for the other countries in this group. However, from about 1987 there is a marked reversal of this trend, a phenomenon which probably explains why the UK was rejected for pooling in Section 5. The reasons for this pattern are unclear, but one possibility is that financial institutions themselves may impose liquidity constraints. Throughout the second half of the 1980s, there was a remarkable build-up of debt in relation to net worth within the UK household sector.[10] Asymmetric information problems in these circumstances, and in the absence of official regulations, may lead to equilibrium credit rationing by financial institutions.[11] Alternatively, after a period of excessive borrowing following financial liberalisation, a debt overhang may have generated more conservative attitudes on the part of UK households.

This possibility underlines a more general qualification to the finding that excess sensitivity of consumption to income declined in the late 1970s and throughout the 1980s in a number of countries that liberalised their financial markets. Financial liberalisation in the presence of pre-existing excess demand for credit by households could be associated with once-for-all portfolio re-adjustments and period-specific apparent declines in the excess sensitivity parameter which might later be reversed. While this must be considered a very real possibility, it is nevertheless the case that countries in this group other than the United Kingdom all show evidence of increased consumption smoothing, in spite of widely differing experiences with regard to household indebtedness and shocks to real income experienced in the 1980s. Thus, for example, Australian households were relatively conservative in their borrowing through the 1980s (Callen 1991), and still show strong evidence of greater consumption smoothing in the face of major adverse movements in the terms of trade in the middle of the 1980s. It is possible, therefore, that the UK household sectors' excessive use of

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[10] UK experience over this period is summarised by Franklin et al (1989).

[11] The credit rationing proposition resulting from asymmetric information is most clearly exposited in Stiglitz and Weiss (1981). Bernanke and Gertler (1989) present a model in which cyclical variations in the level of economic activity are amplified via the effects of agency costs on the price of external funding. The greater the level of corporate net worth the lower are agency costs. But net worth generally varies procyclically, aggravating deadweight agency costs, reducing investment and magnifying the extent of the downturn in activity. This effect reverses itself for an upturn in activity. Similar models have been presented by Greenwald and Stiglitz (1989), Greenwald and Stiglitz (1990) and Williamson (1987).

credit markets may itself be a period specific and a relatively unique phenomenon.

7. CONCLUSIONS

According to the econometric evidence presented in this paper, aggregate consumption in the United States, Japan, Canada and Australia seems to be less responsive to fluctuations in income in the 1980s than in the 1970s or 1960s. The favourite candidate for explaining this phenomenon is the combined and reinforcing effects of financial liberalisation (the progressive lessening of the extent and intensity of official regulations in both national and international capital markets) and innovation. Further support for this conclusion was also evident in the findings for countries where *a priori* information suggests a lack of financial liberalisation over the sample period. While inter-country comparisons of precise λ values are unreliable, for reasons discussed in section III, there is nevertheless little evidence to suggest declines in this parameter in the cases of Germany, France and Italy.

These conclusions are based on empirical tests which hinge on an equation derived on the basis of several maintained hypotheses. A finding of excess sensitivity of current consumption to disposable income could be the consequence of a failure of one or more of those hypotheses. However, the broad pattern displayed by the excess sensitivity parameter, both across countries and over time, is best accounted for by the changing pattern of deregulation and financial innovation in most cases. This conclusion obtains further corroboration from tests that reveal some asymmetric household consumption behaviour in response to increases and decreases in income combined with a tendency for the magnitude of these asymmetries to fall systematically over time. Nor are the results altered when allowance is made for changing intertemporal substitution effects. One of the main contenders for the rejection of the random walk consumption model is a failure of the rational expectations maintained hypothesis. But for many countries the results obtained would imply that consumers were becoming systematically less myopic over time. That this could occur independently of financial liberalisation seems unreasonable.

To the extent that greater consumption smoothing possibilities are associated with increasing financial liberalisation, there are likely to be important implications for policy. Household demand will depend much more on the behaviour of real interest rates, financial prices and expectations, which operate via wealth and intertemporal substitution effects, than on any shift in liquidity constraints that can be manipulated easily by the authorities. Perceptions about the longer-run goals of monetary policy actions which might influence permanent income are therefore more likely to be important influences on household demand than any fluctuations in interest rates perceived to be temporary. Furthermore, short-run fiscal policy measures may be a relatively less useful tool for demand management in circumstances where liquidity constraints are not binding. This is because transitory income is less relevant for spending decisions as the proportion of the population who are liquidity constrained declines.

Table 1: Slope (λ) OLS (instrumental variables) Estimates

$$\Delta \ln C_t = \mu + \lambda \Delta \ln \hat{y}^d t + e_t$$

	1960s (-1969:4)		1970s (1970:1-1979:4)		1980s (1980:1-)		1960s/1970s (-1979:4)	
Individual country results								
United States	0.50		0.47		0.25		0.27	
(1956:1-1988:4)	(0.18)**	(2.46)	(0.12)**	(2.21)	(0.18)	(1.98)	(0.12)**	(1.96)
	(0.19)**		(0.17)**		(0.22)		(0.18)	
Japan	0.42		0.31		0.14		0.34	
(1960:1-1988:1)	(0.17)**	(1.98)	(0.07)**	(2.31)	(0.09)	(1.93)	(0.07)**	(2.14)
	(0.17)**		(0.08)**		(0.08)		(0.05)**	
Germany	0.37		0.67		0.98		0.56	
(1960:1-1988:1)	(0.13)**	(2.74)	(0.17)**	(2.67)	(0.20)**	(2.71)	(0.12)**	(2.79)
	(0.10)**		(0.20)**		(0.10)**		(0.13)**	
France	0.48		0.12		0.31		0.40	
(1963:1-1988:1)	(0.28)	(2.18)	(0.19)	(2.45)	(0.20)	(2.51)	(0.21)*	(2.44)
	(0.40)		(0.17)		(0.15)**		(0.24)	
Italy	0.47		0.43		0.46		0.66	
(1960:1-1988:4)	(0.17)**	(0.40)	(0.19)**	(1.97)	(0.08)**	(0.96)	(0.14)**	(1.80)
	(0.33)		(0.15)**		(0.11)**		(0.12)**	
United Kingdom	0.08		0.12		0.14		0.09	
(1963:1-1988:4)	(0.19)	(2.73)	(0.12)	(2.45)	(0.12)	(2.05)	(0.12)	(2.52)
	(0.13)		(0.10)		(0.07)**		(0.09)	
Canada	0.30		0.21		0.16		-0.01	
(1960:1-1988:4)	(0.16)*	(2.81)	(0.16)	(2.29)	(0.14)	(1.33)	(0.16)	(2.31)
	(0.16)*		(0.11)*		(0.13)		(0.14)	
Australia	0.37		0.25		0.20		0.20	
(1960:1-1988:4)	(0.08)**	(1.57)	(0.14)*	(2.03)	(0.10)**	(2.32)	(0.10)**	(1.74)
	(0.08)**		(0.08)**		(0.07)**		(0.07)**	

Note: Unadjusted standard errors and Durbin-Watson statistics are reported below the co-efficient. A second standard error is reported below these -- these are robust errors calculated with an autocorrelation correction of 4, as the data is quarterly. One and two asterisks indicate difference from zero at the 10 and 5 per cent levels.

Table 2: Unit Normal Tests of the Hypothesis of Declining λ Values in Later Relative to Earlier Periods - OLS (instrumental variables)

	$\hat{\lambda}60>\hat{\lambda}70$		$\hat{\lambda}70>\hat{\lambda}80$		$\hat{\lambda}60/70>\hat{\lambda}80$		$\hat{\lambda}60>\hat{\lambda}80$	
				Individual country results				
United States	0.14,	0.12	1.02,	0.79	0.09,	0.07	0.98,	0.86
Japan	0.60,	0.59	1.49*,	1.50*	1.75**,	2.12**	1.46*,	1.49*
Germany	-1.40,	-1.34	-1.18,	-1.39	-1.80,	-2.56	-2.56,	-4.31
France	1.06,	0.83	-0.69,	-0.84	0.31,	0.32	0.49,	0.40
Italy	0.16,	0.11	-0.15,	-0.16	1.24,	1.23	0.05,	0.03
United Kingdom	-0.18,	-0.24	-0.12,	-0.16	-0.26,	-0.39	-0.27,	-0.41
Canada	0.40,	0.46	0.24,	0.29	-0.79,	-0.87	0.66,	0.68
Australia	0.74,	1.06	0.29,	0.47	0.00,	0.00	1.33*,	1.60*

Note: Test statistics using unadjusted and adjusted SE estimates are reported respectively. The test statistics presented in the table are calculated as follows:

$$Z = \frac{\hat{\lambda}_E - \hat{\lambda}_L}{\sqrt{\hat{\sigma}_E + \hat{\sigma}_L}}$$

where λ_E and λ_L are the estimated coefficients for the relevant earlier and later periods and σ_E and σ_L are the corresponding variance estimates. Z is approximately normally distributed with zero mean and unit variance for moderately large samples (a condition fulfilled here with 40 observations for most subperiods). The critical values for the normal distribution at the 5 per cent and 10 per cent levels are 1.65 and 1.29 respectively. Z values in excess of these lead to acceptance of the null hypothesis of declining λ's. One asterisk indicates that the null cannot be rejected at the 10 per cent level, and two that it cannot be rejected at the 5 per cent level. The absolute values of Z are presented in the table.

Table 3: Asymmetric Behaviour

Test equation: $\Delta lnC_t = \mu' + \lambda_1 \Delta lny_t^+ + \lambda_2 \Delta lny_t^- + (1-\lambda)\omega_t$

Figures in parentheses are absolute values of standard errors

	1960s			1970s			1980s		
	λ_1	λ_2	H_0	λ_1	λ_2	H_0	λ_1	λ_2	H_0
United States	0.40**	1.89**		0.41**	0.60**		0.04	0.63	
	(0.18)	(0.88)	A	(0.18)	(0.23)	R	(0.23)	(0.48)	R
Japan	0.32	1.19		0.04	0.75**		0.04	0.22	
	(0.20)	(2.77)	R	(0.11)	(0.13)	A	(0.13)	(0.19)	R
Germany	0.15	1.37**		0.72**	0.43		1.31**	0.70**	
	(0.19)	(0.58)	A	(0.20)	(0.72)	R	(0.41)	(0.37)	R
France	0.60	0.54		0.22	0.07		-0.17	0.84*	
	(0.37)	(1.09)	R	(0.33)	(0.94)	R	(0.32)	(0.45)	A
Italy	0.39**	14.80		0.01	0.63**		0.26**	0.75**	
	(0.19)	(13.42)	R	(0.21)	(0.25)	A	(0.12)	(0.21)	R
United	0.59*	-0.34		0.06	0.27		0.33*	-0.07	
Kingdom	(0.31)	(0.29)	R	(0.20)	(0.31)	R	(0.18)	(0.34)	R
Canada	0.11	1.37*		-0.07	0.79		0.05	0.21	
	(0.22)	(0.73)	A	(0.22)	(0.50)	A	(0.20)	(0.34)	R
Australia	0.42**	0.26		0.25	0.49		0.05	0.47*	
	(0.10)	(0.18)	R	(0.17)	(0.46)	R	(0.16)	(0.25)	A
Average λ_1 and λ_2 estimates across countries	0.37	2.64		0.21	0.50		0.24	0.46	

Note: The null hypothesis H_0 is $\lambda_2 > \lambda_1$ and the alternative that $\lambda_2 \leq \lambda_1$. A value of the Z statistic (see Table 2) in excess of 1.65 or 1.29 indicates acceptance (A) of H_0 at the 5 and/or 10 per cent levels respectively. R indicates rejection of H_0.

Table 4: Estimates for Equation (17) in the Text, i.e.

$$\Delta \ln C_t = \mu + \lambda \Delta \ln \hat{y}_t + \alpha r_{t-1} + \gamma \Delta \hat{i}_t + \omega_t$$

Absolute t values in parentheses

		1960s	1970s	1980s	1960s/1970s
United States	λ	0.36*	0.43**	0.20	0.23*
		(2.02)	(3.67)	(1.17)	(1.66)
	α	0.001	0.0007	0.001**	0.001*
		(0.75)	(1.31)	(2.54)	(1.93)
	γ	0.003	-0.0005	0.003*	-0.0003
		(0.53)	(0.21)	(1.84)	(0.15)
	DW	2.15	2.10	1.96	1.99
Japan	λ	0.44**	0.28**	0.16*	0.32**
		(2.57)	(3.35)	(1.90)	(4.15)
	α	0.001	0.0003	0.001	0.0004
		(1.55)	(0.71)	(1.34)	(1.49)
	γ	-0.024	-0.01	-0.001	-0.015**
		(1.64)	(1.38)	(0.42)	(2.42)
	DW	1.88	1.96	2.11	1.87
Germany	λ	0.36**	0.66**	1.07**	0.53**
		(2.48)	(3.20)	(5.13)	(3.51)
	α	-0.001	-0.0003	-0.001	-0.001
		(1.20)	(0.28)	(0.93)	(1.05)
	γ	0.17	-0.0001	-0.006**	-0.0001
		(0.99)	(0.11)	(2.19)	(0.14)
	DW	2.61	2.58	2.54	2.65
France	λ		0.16	0.28	
			(0.83)	(1.39)	
	α		0.001	0.001	
			(1.14)	(1.40)	
	γ		-0.005**	0.001	
			(2.16)	(0.76)	
	DW		2.61	2.37	

Table 4 (continued)

Italy	λ	0.73**	0.30	0.48	0.86**
		(4.92)	(1.30)	(6.14)	(5.00)
	α	-0.002**	0.0003	-0.0002	-0.0002
		(4.72)	(1.02)	(1.05)	(1.07)
	γ	-0.024*	0.001	0.001	0.001
		(3.83)	(1.20)	(0.90)	(0.61)
	DW	1.35	1.87	0.84	1.86
United Kingdom	λ	0.06	0.102	0.17	0.08
		(0.32)	(0.82)	(1.34)	(0.71)
	α	0.001	0.001*	0.0002	0.001*
		(0.32)	(1.85)	(0.40)	(1.70)
	γ	-0.012	-0.004	-0.001	-0.005
		(0.81)	(0.87)	(0.72)	(1.13)
	DW	2.97	2.75	2.11	2.70
Canada	λ	0.28*	0.23	0.17	0.15
		(1.68)	(1.33)	(1.19)	(0.93)
	α	0.0003	-0.0002	0.001	-0.0002
		(0.31)	(0.34)	(0.71)	(0.38)
	γ	-0.01*	-0.002	-0.0004	-0.002
		(1.80)	(0.72)	(0.33)	(0.92)
	DW	2.98	2.22	2.08	2.26

Note: One and two asterisks indicate difference from zero at the 10 and 5 per cent levels. Interest rate data for Australia were not available for a sufficiently long time period to complete the tests, nor were interest rate data available for France for the 1960s.

Table 5: Financial Liberalisation in the 1970s and 1980s

	Rate/Quantity Deregulation of Intermediaries	Powers Deregulation Competition Between Intermediaries	Foreign Exchange Deregulation
	(Rapid Liberalisation in the 1980s)		
United States	Mainly in the late 1970s and early 1980s.	From the mid-1970s important.	Always deregulated in 1970s and 1980s.
Japan	Carried out gradually through the 1980s.	Gradual introduction of new instruments, mainly in 1980s.	For all the 1980s (not 1970s).
United Kingdom	Controls widely used until 1980.	Being gradually carried out mainly from the mid 1980s. Cartel-like behaviour evident.	Removed controls in 1979.
Canada	Always deregulated in 1970s and 1980s.	Always deregulated in 1970s and 1980s.	Always deregulated in 1970s and 1980s.
Australia	Controls widely used until early 1980s.	Regulations eliminated and foreign bank competition introduced in mid 1980s.	Removed controls from 1983.
	(Countries Slow to Liberalise)		
Germany	Always deregulated in 1970s and 1980s.	Strongly controlled and little deregulation in 1970s or 1980s. Cartel-like behaviour evident.	Always deregulated in 1970s and 1980s.
France	Controls widely used in 1970s and 1980s.	Being gradually carried out mainly from the mid 1980s. Cartel-like behaviour evident.	Controls widely used and only in late 1980s phasing out begins.
Italy	Credit ceilings used until 1983.	Ready availability of short Treasury Bills since 1975, but strong regulation of intermediaries. Cartel-like behaviour evident.	Highly regulated in 1970s and most of the 1980s - some recent easing.

Source: OECD.

Table 6: Pooled Results: Group 1
(United States, Japan, Canada and Australia)

		1960s	1970s	1980s	1960s/1970s
United States	μ′	0.003	0.002	0.005	0.004
		(0.001)	(0.001)	(0.001)	(0.001)
	λ	0.42**	0.43**	0.01	0.25*
		(0.16)	(0.11)	(0.15)	(0.12)
Japan	μ′	0.010	0.008	0.005	0.009
		(0.003)	(0.002)	(0.001)	(0.002)
	λ	0.42**	0.28**	0.14	0.30**
		(0.16)	(0.07)	(0.08)	(0.07)
Canada	μ′	0.003	0.004	0.004	0.005
		(0.002)	(0.002)	(0.002)	(0.002)
	λ	0.46**	0.25	0.16	0.14
		(0.16)	(0.16)	(0.12)	(0.16)
Australia	μ′	0.004	0.003	0.002	0.004
		(0.001)	(0.001)	(0.001)	(0.001)
	λ	0.35**	0.19	0.15	0.15
		(0.07)	(0.14)	(0.11)	(0.12)
Log likelihood		525.2	536.4	483.7	1012.6
United States	μ′	0.004	0.003	0.004	0.004
		(0.001)	(0.001)	(0.001)	(0.001)
Japan	μ′	0.010	0.008	0.005	0.010
		(0.002)	(0.002)	(0.001)	(0.002)
Canada	μ′	0.003	0.004	0.004	0.004
		(0.002)	(0.002)	(0.002)	(0.001)
Australia	μ′	0.004	0.003	0.002	0.004
		(0.001)	(0.004)	(0.001)	(0.001)
	λ	0.38**	0.29**	0.14**	0.24**
		(0.06)	(0.05)	(0.05)	(0.05)
Log likelihood		525.0	535.4	483.6	1011.8
United States, Japan, Canada, Australia	μ′	0.004	0.003	0.004	0.004
		(0.001)	(0.001)	(0.001)	(0.001)
	λ	0.47**	0.33**	0.14**	0.34**
		(0.06)	(0.05)	(0.05)	(0.05)
Log likelihood		517.7	532.6	481.8	1002.5

Note: A SURE estimation technique and the same instrumental variables as for Table 1 results are used here. Standard errors are shown in parentheses.

Table 7: Unit Normal Tests of the Hypothesis of Declining λ Values in Later Relative to Earlier Periods -- Pooled Estimates

	$\hat{\lambda}60>\hat{\lambda}70$	$\hat{\lambda}70>\hat{\lambda}80$	$\hat{\lambda}60/70>\hat{\lambda}80$	$\hat{\lambda}60>\hat{\lambda}80$
United States	0.05	2.25**	1.25	1.87**
Japan	0.80	1.31*	1.51**	1.56**
Canada	0.93	0.45	-0.10	1.50*
Australia	1.02	0.22	0.00	1.53*
Joint λ	1.15	2.12**	1.28*	3.07**

Note: The test statistics presented in the table are calculated as follows:

$$Z = \frac{\hat{\lambda}_E - \hat{\lambda}_L}{\sqrt{\hat{\sigma}_E + \hat{\sigma}_L}}$$

where λ_E and λ_L are the estimated coefficients for the relevant earlier and later periods and σ_E and σ_L are the corresponding variance estimates. Z is approximately normally distributed with zero mean and unit variance for moderately large samples (a condition fulfilled here with 40 observations for most subperiods). The critical values for the normal distribution at the 5 per cent and 10 per cent levels are 1.65 and 1.29 respectively. Z values in excess of these lead to acceptance of the null hypothesis of declining λ's. One asterisk indicates that the null cannot be rejected at the 10 per cent level, and two that it cannot be rejected at the 5 per cent level. The absolute values of Z are presented in the table.

Table 8: Pooled Results: Group 2 (Germany, France and Italy)

		1960s	1970s	1980s	1960s/1970s
Germany	μ'	0.008	0.003	-0.003	0.005
		(0.002)	(0.002)	(0.002)	(0.002)
	λ	0.31**	0.65**	1.11**	0.50**
		(0.12)	(0.18)	(0.19)	(0.14)
France	μ'	0.005	0.006	0.003	0.004
		(0.003)	(0.002)	(0.001)	(0.002)
	λ	0.50	0.12	0.25	0.50
		(0.26)**	(0.19)	(0.18)	(0.20)**
Italy	μ'	0.006	0.006	0.005	0.006
		(0.002)	(0.002)	(0.001)	(0.001)
	λ	0.43**	0.41**	0.42**	0.55**
		(0.15)	(0.16)	(0.10)	(0.12)
Log likelihood		257.40	409.40	401.20	651.80
Germany		0.007	0.005		0.004
		(0.002)	(0.002)		(0.001)
France	μ'	0.006	0.004		0.004
		(0.003)	(0.002)		(0.001)
Italy		0.007	0.007		0.006
		(0.002)	(0.001)		(0.001)
	λ	0.38**	0.41**		0.52**
		(0.09)	(0.11)		(0.08)
Log likelihood		257.20	407.30		651.80
Germany	μ'	0.007	0.006		0.005
France		(0.001)	(0.001)		(0.001)
Italy	λ	0.38**	0.40**		0.53**
		(0.09)	(0.11)		(0.08)
Log likelihood		257.10	405.70		650.60

Note: A SURE estimation technique and the same instrumental variables as for Table 1 results are employed here. Standard errors are shown in parentheses.

APPENDIX

Data and Replication Results

Data for eight OECD countries, including the United States, Japan, Canada, Australia, the United Kingdom, Germany, France and Italy are collated.

Quarterly total consumption expenditure, disposable income and exports data are obtained from the OECD National Accounts. Total personal consumption less purchases of durables is calculated for the subset of countries (namely, the United States, Japan, France, Italy, the United Kingdom and Canada) which have the appropriate statistics available. All data are measured in per capita terms.

Unemployment, population and interest rate statistics are also obtained from the OECD.

The sample period for each country extends from 1960:1 to 1988:4 with the exception of Japan with a sample period ending in 1988:1.

The unsuccessful attempts to replicate J and P's results are reported only for the two largest countries in their sample, namely, the United States and Japan. Table A1 details our estimates for these countries over both J and P's original sample period and an updated period. J and P source the *National Accounts*, OECD, Volume II, Detailed Statistics, 1986 for data on Sweden, the United Kingdom, Japan, Italy, Spain and Greece. Equivalent data for the US were taken from *The Economic Report of the President* (1986).

As mentioned in the text, J and P define consumption to "... exclude expenditure on durables, ... [namely] appliances, furniture and means of transportation", and use the first lag of non-durables consumption, personal disposable income, government expenditure and exports as instruments. However, in order to reproduce J and P's instrumental variables co-efficient estimates for the US, it was necessary to exclude personal consumption expenditure both on durable goods and on non-durable services. For Japan, it was necessary to use national disposable income (which includes the business sector) rather than personal disposable income, and total government outlays as an instrument rather than the national accounts measure of government expenditure (used for the USA). With these inconsistent definitions, it was possible to reproduce J and P's original estimates shown in the last two columns of Table A1. The estimates from consistently defined data are also shown in Table A1, both for the original sample period as used by J and P, and for a larger, more up-to-date sample.

This replication exercise also revealed that the estimates of a_1 are often very close to 1, suggesting that J and P's model (equation (12) in the text) collapses to a linear equation in first differences:

$$\Delta C_t = a_0 + \lambda(\Delta y^d{}_t) + e_t \tag{A1}$$

A.1: Tests for Non-Stationarity and Non-Zero Drift

Augmented Dickey-Fuller (1979, 1981) and Phillips-Perron (1988) Z_t tests for non-stationarity are set out in Tables A2 and A3. The ADF procedure also allows a test of the significance of the constant term which indicates the presence of non-zero drift. Notwithstanding the small sample problems associated with these tests,[12] both the ADF and Phillips-Perron Z_t tests find a unit root in consumption, income and remaining instruments for both the US and Japan at the 5% significance level. The Z_t test reports US Consumption of Non-Durables to be trend stationary at the 10% significance level and the ADF statistic on the constant term cannot reject non-zero drift for this variable alone.

12 Given the unworkably small size of J and P's samples, the ADF and Phillips-Perron tests are reported over our extended estimation periods. It should be noted, however, that these tests evidence a drop in power when applied to samples of less than 100 observations.

Table A1
The Excess Sensitivity of Consumption: Japelli & Pagano's NLIV Method

Country	Parameters	Starting Values	NLIV Estimates	(t-statistic)	Error Sum of Squares	Japelli & Pagano NLIV Estimates	Japelli & Pagano (t-statistic)
U.S.	a0	0.00	45.30	(0.75)			
(1961-1984)	a1	1.00	1.00	(46.70)			
	λ	0.20	0.42	(4.77)	36309.83	0.21	(2.30)
	a0	10.00†	6.20	(0.07)			
	a1	10.00	0.75	(5.44)			
	λ	10.00	0.77	(15.61)	84328.19		
(1961-1990)	a0	0.00	7.34	(0.13)			
	a1	1.00	1.02	(35.67)			
	λ	0.20	0.51	(5.62)	61929.45		
	a0	10.00†	-163.00	(-1.93)			
	a1	10.00	0.75	(6.88)			
	λ	10.00	0.86*	(15.61)	149654.38		
Japan	a0	0.00	-0.33	(-0.01)			
(1972-1983)	a1	1.00	1.04	(12.11)			
	λ	0.20	0.51	(4.87)	2848.33	0.34	(5.00)
	a0	10.00†	-484.38	(-4.08)			
	a1	10.00	0.67	(5.80)			
	λ	10.00	1.76	(4.93)	6258.03		
(1972-1989)	a0	0.00	3.33	(0.20)			
	a1	1.00	1.03	(27.24)			
	λ	0.20	0.51	(5.74)	3183.25		
	a0	10.00†	-175.28	(-2.52)			
	a1	10.00	0.62	(481.00)			
	λ	10.00	1.08	(16.25)	4677.31		

† J and P's model displayed some sensitivity to parameter starting values. Grid searches reveal two optima both for Japan and the US. However, in each case, estimates originating from the more theoretically plausible starting values also displayed the lower error sum of squares.

* The estimated λ increases for the US when the sample period is extended to the 1990s. Since this longer and more recent period encompassed increasing deregulation of financial markets, it would seem more reasonable to expect any liquidity constraints to be declining rather than increasing.

Table A2: ADF Tests for Non-Stationarity

Country	Updated Sample	
	Constant Term	ADF Test
Japan *[1972-1989]*		
CND	0.25	0.50
PDY	1.14	-0.71
GE	0.68	-0.40
X	2.00	-1.89
United States *[1959-1990]*		
CND	3.71**	-3.51
PDY	2.12	-1.35
GE	2.47	-2.16
X	2.38	-3.04

Where CND is consumption of non-durables, PDY is personal disposable income, GE is government expenditure and X is exports.

Underlined statistics denote those regressions which display a significant trend. Critical values at the 5% level for these statistics are -3.60 and 3.20 on the constant term; all other tests find an insignificant trend and face critical values at the 5% level of -3.00 and 2.61 on the constant term.

** The null of non-stationarity is rejected at the 5% level.

Table A3: Phillips-Perron Z_t Tests for Non-Stationarity

Country	Variable	Sample Period	Z_t With Trend	Z_t No Trend
United States	CND	1959-1990	-2.86*	-0.25
	PDY		-1.88	-0.99
	GE		-1.96	-1.51
	X		-2.40	-0.68
	Δ_1CND	1960-1990	-3.66**	-3.73**
	Δ_1PDY		-5.22**	-5.06**
	Δ_1GE		-3.92**	-3.69**
	Δ_1X		-3.74**	-3.73**
Japan	CND	1972-1989	-2.44	0.69
	PDY		-3.39	-0.93
	GE		-2.19	-0.85
	X		-2.31	-2.31
	Δ_1CND	1973-1989	-3.00	-2.84*
	Δ_1PDY		-5.02**	-4.69**
	Δ_1GE		-2.99	-3.03**
	Δ_1X		-2.60	-2.78*

** The null of non-stationarity is rejected at the 5% level with critical values of -3.60 and -3.00 for the test with a trend and excluding a trend respectively.

* The null of non-stationarity is rejected at the 10% level with critical values of -3.24 and -2.63 for the test with and without a trend respectively.

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