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Disaggregate Wealth and Aggregate Consumption: an Investigation of Empirical Relationships for the G7*

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Abstract

To date, studies of wealth effects on consumption have mainly used aggregate wealth definitions on a single-country basis. This study seeks to break new ground by analysing disaggregated financial wealth in consumption functions for G7 countries. Contrary to earlier empirical work, we find that illiquid financial wealth (i.e. securities, pensions and mortgage debt) tends to be a more important long-run determinant of consumption than liquid financial wealth. These results imply potential instability in consumption functions employing aggregate wealth. Our results are robust using SURE; when testing with a nested specification; and when using a linear model.

I. Introduction

Amongst the numerous time series studies of consumers' expenditure at the aggregate level,[1] the issue of the existence and size of wealth effects in the consumption function has come to the fore strongly in recent years. This is

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[1]See Muellbauer and Lattimore (1995) for a comprehensive survey of empirical work on consumption.

JEL Classification numbers: E21, F15.

partly a consequence of the strength of consumption and low level of savings in the US in the late 1990s (see *inter alia* Ludvigsen and Steindel, 1999; Poterba, 2000; IMF, 2002), which is often thought to be linked to rising share prices and resultant wealth effects. Equally, strong shifts in asset prices, as in the UK in the late 1980s and early 1990s, may have amplified the business cycle partly via wealth effects on consumption, linked to credit expansion (Davis, 1995). Similar arguments may apply to the protracted weakness of consumption in Japan in the 1990s.

In the longer term, interest in wealth holdings and its effect on consumption has been linked to macroeconomic trends. The ratio of wealth to income and consumption is rising in the major industrial economies, suggesting that the ability to draw on wealth to maintain consumption is increasing. In terms of macroeconomic policy, the link from wealth to consumption gives a reason for interest in the effect of asset prices on the aggregate economy. The case is strengthened by the role of asset prices in 'credit cycles' in the late 1980s. The regime shift of financial liberalization may require us to reconsider wealth effects on consumption. Excess sensitivity of consumption to current income is often attributable to liquidity constraints (Campbell and Mankiw, 1991), whilst freer availability of credit tends to reduce the importance of current income as a determinant of consumption. Also with financial liberalization, borrowing may develop a strong independent, and at times even positive, effect on consumption. This may not be consistent with conventional consumption functions, which assume a negative sign for borrowing in the construction of net wealth. In this context, the effect of debt on consumption might also differ according to the level of gearing and the stage of the business cycle, being favourable to consumption at low levels of borrowing, but negative when there is high gearing and a downturn in the economy. As such we may have experienced a decline in the net effect of liquid assets on consumer expenditure.

The institutionalization of savings, and in particular the growing share of household wealth held in the form of life insurance and pension funds (Davis and Steil, 2001), raises the question of whether the relative illiquidity of such wealth may give it a weaker effect on consumption than directly held and liquid wealth. On the other hand, the decreasing elements of insurance (e.g. in the shift from defined benefit to defined contribution pension funds) which would otherwise insulate the beneficiary from a change in the price of assets, could increase the wealth effect from this source. Linked to this, the ageing of the population is helping to drive an increased accumulation of wealth, independent of other motives for savings. There is an interest in how consumption will evolve as population ageing and related accumulation of assets proceed.

As a consequence of these developments, a number of recent studies have investigated the properties of total net wealth as a determinant of consumption (see Poterba, 2000; Davis and Palumbo, 2001). However, rather fewer studies have sought to disaggregate wealth and assess whether rises in sub-components of wealth have differential effects on consumption. And even those that do attempt this are rarely undertaken on a comparative, cross-country basis or with sectoral balance sheet data as opposed to proxies using asset prices. Hence, this paper seeks to break new ground by presenting and comparing data and estimates of the relation between components of personal or household sector financial wealth and consumers' expenditure for the US, UK, Germany, France, Italy, Japan and Canada. We employ a backward looking specification for aggregate consumer expenditure where the long-run relationship between consumption, income and financial wealth is estimated within an error correction model (ECM). Then we seek to disaggregate the existing long-run wealth terms within our established short-run model. In turn we examine whether the typical imposition of an equivalent long-run elasticity for the components of wealth is consistent with the data.

The paper is structured as follows. In Section II we briefly note the theoretical background which could justify inclusion of wealth in a consumption function; we also look at implications of disaggregation, using recent trends in portfolios in the G7 as background. Section III develops our empirical specification. Section IV presents our own results. In the conclusion we seek to draw out some underlying features and policy implications of these outturns.

II. Theoretical and empirical background

In this paper, we estimate benchmark consumption functions based on the 'Life Cycle' hypothesis of Ando and Modigliani (1963) as derived in Deaton (1992). In this model, planned aggregate consumption (C_t^*) is a function of total resources. Total resources are the sum of human wealth (H_t) and net financial wealth (W_{t-1}). Planned consumption can accordingly be expressed as a function of H_t and W_{t-1}:

$$C_t^* = m(H_t + W_{t-1}), \tag{1}$$

where m is the marginal propensity to consume (MPC) out of total resources on average across the population. If we assume that planned consumption does not always equal actual consumption and that human wealth can be proxied by some function, k, of current labour income (i.e. $H_t = kY_t$), we can derive the following relationship for actual consumption (C_t):

$$C_t = aY_t + bW_{t-1} + \varepsilon_t. \tag{2}$$

Further assumptions are warranted to estimate this specification of the consumption function. As suggested by Campbell and Deaton (1989), income in levels is unlikely to be difference stationary. In particular, the first difference of the level of income does not display constant variance: earlier increases in the level of income, in any reasonable sample of data, are likely to be substantially less than increases later in the sample. This non-constant variance would mean any long-run relationship for consumption would be spurious potentially, given that not all of our variables are difference stationary, and a short run ECM for consumption would have non-stationary dynamics. Consequently, we adopt a logarithmic approximation for equation (2) to ensure that income, in natural logs, is difference stationary and hence that our long-run relationship can be non-spurious. The logarithmic approximation is as follows:

$$\ln C_t = c_0 + \alpha \ln Y_t + \beta \ln W_{t-1} + \xi_t. \tag{3}$$

This is the approach adopted in recent work on the US consumption function by Ludvigson and Steindel (1999) and Davis and Palumbo (2001).

The derivation above assumes that non-human wealth is homogeneous, an assumption followed in the bulk of the literature surveyed in Byrne and Davis (2001). However, this is not the case either in terms of liquidity or capital certainty, which suggests that the effects of its sub-components on consumption may vary. Moreover, the household sector is not homogeneous in terms of total wealth of households and stage of the life cycle, for example. Both of these may affect the portfolio composition of the household and its response to changes in the components of wealth.

In this context, *a priori*, it is plausible that in a liquidity-constrained situation, net liquid assets, as well as disposable income, will strongly affect consumption. Where credit is not readily available, the liquidity of an asset is the key aspect of its substitutability for consumer goods, and hence less liquid assets, such as equities and life and pension claims, will have a smaller effect on consumption than liquidity. Moreover, the key component of net liquidity will be (gross) liquid assets rather than indebtedness,[2] as the latter is constrained *ex hypothesi*. A similar argument holds for the income effect on consumption, which will be large for liquidity-constrained consumers. Reflecting this, work on consumption up to the 1980s typically involved liquid assets only as a wealth term (see Hendry and von Ungern-Sternberg, 1981, and the survey in Davis, 1984).

[2]Net liquid assets are typically defined to include total household debt as a negative variable (as it is a liability).

However, when liquidity constraints ease, both the importance of liquid assets and of disposable income may well decline. The former is likely to become less important than total wealth, which indicates resources available over the life cycle for consumption, against which the household may borrow. As noted, consumption functions shifted in the 1990s to such wider definitions of wealth. Whether there will be differences in the view taken by the household in respect of types of wealth, and hence their weight in the consumption decision, is less clear *a priori*, for the following reasons.

On the one hand, liquid assets will remain by definition a closer substitute for consumption than illiquid ones, and especially contractual savings such as life and pension fund claims. Also illiquid assets may be concentrated in fewer hands than liquid ones (although this is less the case when there is a funded pension system). On the other hand, if households tend to shift to a greater proportion of illiquid wealth as their overall wealth increases and liquidity constraints ease (because they are willing to take higher risk on part of the portfolio in exchange for higher return), the valuation of such illiquid wealth could come to the fore as an indicator of consumption. However, transaction costs in the equity market might have a similar effect. Moreover, rises in debt as liquidity constraints ease may also lead to a reduction in the measured effect of liquidity (as a rise in borrowing to finance consumption gives a 'negative wealth effect'). Disposable income, meanwhile, will become at most an indicator of human wealth for life cycle purposes. It is the relative importance of these arguments that this article seeks to test.

As background to the above discussion, we note that there are very marked differences in patterns of wealth holding across the G7, with, in particular, a much lower level of equity and institutional holdings in Continental Europe than in the other G7 countries. This could affect the responsiveness of consumption to wealth via (a) similar coefficients on the components of wealth, but different size of the asset components within the portfolio and/or (b) different coefficients as well as (c) different distributions of wealth across the population. It is important to test for the relative importance of these effects via a disaggregated wealth study, such as the work presented here, so as to calibrate wealth effects more accurately. Besides being relevant across the G7 generally, any remaining differences in wealth effects within Continental Europe will render more complex the task of the ECB to conduct monetary policy.

Figures 1–7 present trends in the composition of gross financial wealth in the G7.[3] Canada, the UK and US show a higher proportion of institutional

[3]These are drawn from the National Flow of Funds Balance Sheet tables prepared by the National Central Banks of France, Germany, Italy, Japan and the US and by the National Statistical Offices in the UK and Canada. Mutual fund holdings are included in bonds or equity.

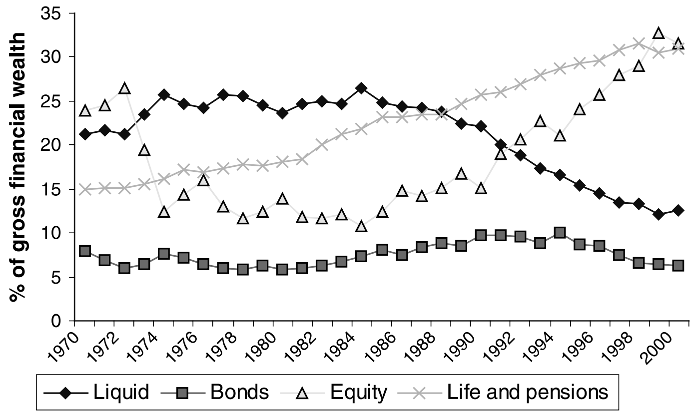

Figure 1. US household sector wealth composition

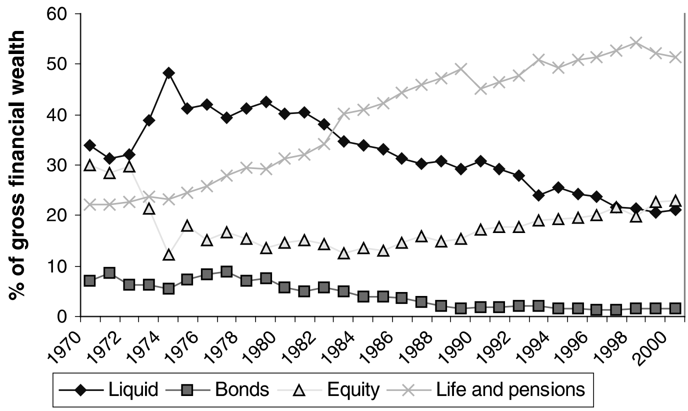

Figure 2. UK household sector wealth composition

saving, while liquid asset holdings tend to be lower. A factor underlying the decline in liquidity, besides relatively higher returns on capital uncertain assets, is financial liberalization, which has reduced the need for precautionary liquid asset holdings. The differences between the US, UK and Canada and Continental European countries should not be exaggerated, as there is some degree of convergence, which, depending on the coefficients in the consumption function, could also imply a convergence of behaviour. In particular, the share of liquid assets in France and Italy, which in the 1970s was over 60% of gross personal sector financial assets, has now more than halved in response to rising equity and institutional investment (in France) and rising equity and bond holdings (in Italy). This pattern is also present in a muted form for Germany, where bond holdings have increased relatively

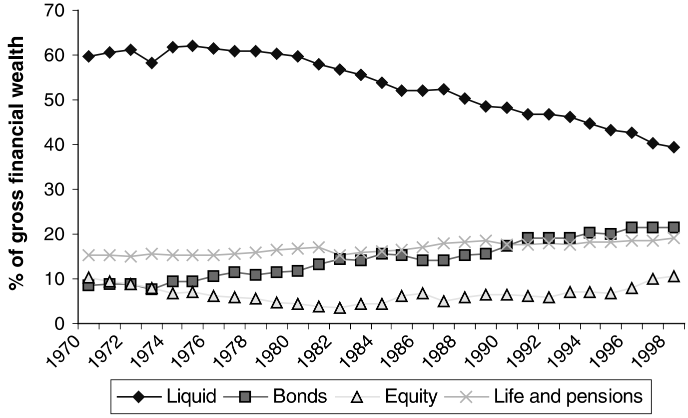

Figure 3. German household sector wealth composition

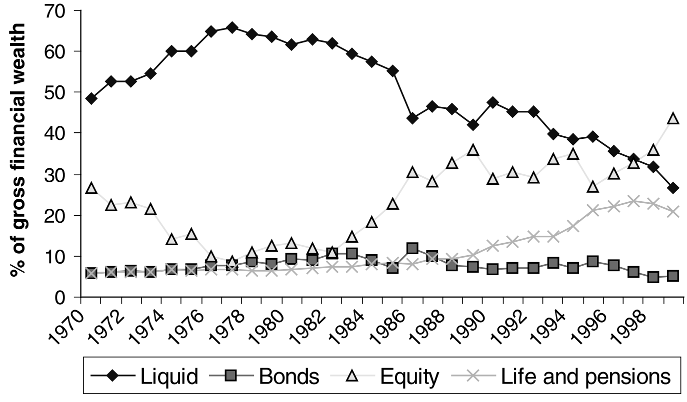

Figure 4. French household sector wealth composition

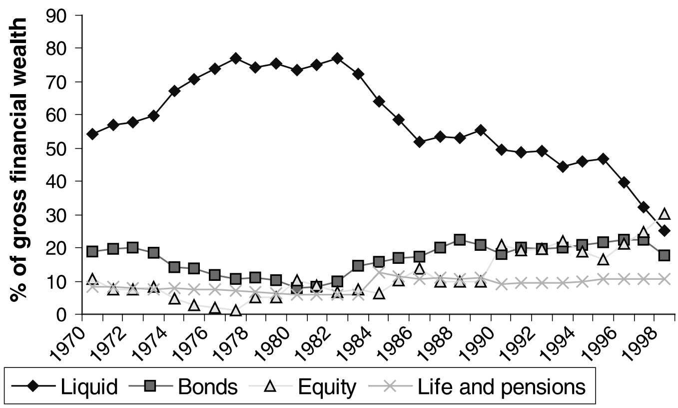

Figure 5. Italian household sector wealth composition

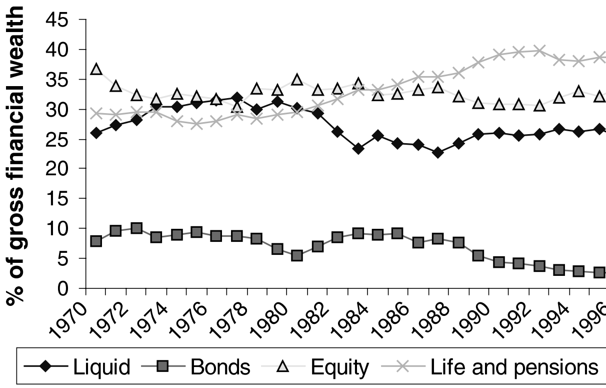

Figure 6. Canadian household sector wealth composition

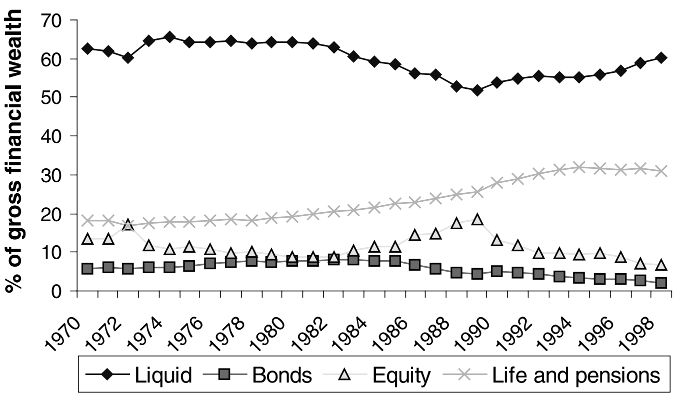

Figure 7. Japanese household sector wealth composition

strongly. Japan is an exception in that liquidity has retained its dominance, accounting for around 60% of portfolios in 1970 and 1998. Whereas convergence had been apparent till 1990, the securities market crises of the 1990s have evidently driven households back to low risk assets. Even the rise in life insurance and pension funds' share of assets reached a plateau in Japan in the early 1990s.

It is also useful to consider patterns of aggregate wealth as a proportion of personal disposable income (PDI). As shown in Byrne and Davis (2001), there is a common pattern all seven countries of a rise in gross wealth relative to PDI, and a generally more muted rise in debt/PDI ratios. Corresponding to these patterns, there has also been an increase in net wealth/PDI. The levels of the gross wealth ratio differ between around five for the US and UK and around three for Continental Europe, Canada and Japan. While this no doubt partly relates to differences in measurement, it may also be linked to the funded pension systems in the US, Canada and the UK shown in the wealth data. A larger part of wealth of households in France, Germany, Italy and Japan is in the form of implicit social security wealth. Meanwhile, debt/PDI ratios are around 1 in Canada, the US, UK, Japan and Germany, while in France they are around 0.6 and 0.3 in Italy.

Only a limited number of studies have disaggregated financial wealth in the consumption function, and they have not generally split financial wealth in the manner that these data patterns suggest, nor have they typically been estimated across a range of countries.[4] Of the existing papers examining disaggregate wealth most concentrate on equity holdings. Brayton and Tinsley (1996) find that the MPC out of stock market wealth in the US is half of other components of net worth. In one notable cross-country study Boone *et al.* (1998) suggest that a 10% fall in stock indexes leads to a fall in consumption of 0.75% in the US, 0.45% in Canada and the UK and a 0.2% fall in other G7 countries. Ludvigsen and Steindel (1999) find, for a reasonably long sample period, that the elasticity is equivalent on stock market and non-stock market wealth. Also, Starr-McCluer (2002), using survey data, questions the existence of any direct stock market effect and emphasize an indirect effect linked to retirement savings in mutual funds and retirement accounts.

III. Data and econometric issues

In the light of the above, we constructed an econometric data set using the flow of funds data illustrated in section II (i.e. liquidity, bonds, shares, life/pension funds, mortgages and other debt) and standard macroeconomic variables (total consumption, C_t, and real PDI, Y_t).[5] We assume that all variables are $I(1)$ or difference stationary in logarithms and this is consistent with Augmented Dickey-Fuller tests of unit root, which are available upon request.

All regressions are estimated by non-linear least squares and the sample period is 1972Q2–1998Q4 for all countries to allow direct comparison. In

[4]According to Brayton and Tinsley (1996), Boone *et al.* (1998) and Starr-McCluer (2002), aggregate studies of the US find that the MPC out of net wealth is between 1 and 7%. There is no strong evidence of any aggregate wealth effect for France according to Bonnet and Dubois (1995) and Grunspan and Sicsic (1997), although Carruth *et al.* (1999) find evidence for a wealth effect using an inflation proxy. Rossi and Visco (1995) find for Italy that the marginal propensity to consume out of net wealth is equal to 3.5%. Generally, the literature finds weaker net wealth effects in Continental Europe and Japan than in the US, the UK and Canada.

[5]All data series are quarterly, seasonally adjusted and from Primark Datastream, where not otherwise stated. The quarterly wealth time series for Germany, France, Italy and Canada is interpolated annual data using changes in broad money for liquid and changes in equity prices for illiquid assets. This is the approach suggested by Chow and Lin (1971).

the presence of non-stationary data we avoid using a static regression approach by utilizing a dynamic ECM, as advocated by Banerjee *et al.* (1993). We note that the results from Inder (1993) suggest estimating the long-run relationship between I(1) variables using an unrestricted ECM via non-linear least squares. He finds evidence that this approach gives precise estimates and valid *t*-statistics, even in the presence of endogenous explanatory variables. As noted below, we also have evidence of a cointegrating long-run relationship between our variables for each of the G7 countries using the Johansen (1988) Trace test. So we believe our estimated coefficients and SEs are accurate, useful for inference and not susceptible to missing long-run variables.

Our econometric approach involved first testing aggregate wealth in the context of 'baseline' consumption specifications. Consequently, the estimated models feature a common error-correction formulation, with the long run having terms in real personal disposable income and real net wealth, as follows:

$$\Delta \ln C_t = \alpha_0 + \alpha_1 \times (\ln C_{t-1} - \beta_1 \times \ln Y_{t-1} - \beta_2 \times \ln W_{t-1}) \\ + \gamma_i \times \Delta \ln C_{t-j} + \gamma_i \times \Delta \ln Y_{t-k}. \tag{4}$$

We initially attempted a full disaggregation with six wealth terms but faced problems with non-linear estimation, with lack of convergence or implausible coefficients. This was to be expected as the data is unlikely to be able to easily distinguish between separate effects from six categories of assets and liabilities. This led us to estimate our consumption function using a simpler disaggregation of wealth. We sought to distinguish between net liquid assets (liquidity less bank lending excluding mortgage debt) and net illiquid assets (shares, bonds, life and pension funds less mortgage debt). We contend that mortgages, being long term in maturity (10–25 years) are much more appropriately included with illiquid than liquid instruments, although we acknowledge that this is not current practice in some countries (such as the UK). In the case of Germany there was a problem with this approach as total mortgage debt exceeded illiquid assets, which required a further disaggregation of German wealth to include a separate variable for mortgage debt.

From equation (4) we consequently derive the following log specification of the consumption function which disaggregates wealth into liquid ($W_{[\mathrm{LQ}]}$) and illiquid ($W_{[\mathrm{ILQ}]}$) components:

$$\Delta \ln C_t = \alpha_0 + \alpha_1 \times \left(\ln C_{t-1} - \beta_1 \times \ln Y_{t-1} - \beta_2 \times \ln W_{[LQ],t-1} - \beta_3 \times \ln W_{[ILQ],t-1}\right) \\ + \text{dynamics}. \tag{5}$$

Two alternatives are also employed to test the robustness of the results. First, we note that the simple disaggregation approach of equation (5) is not nested within our benchmark equation (4). To check that our results are not the result of a non-nested disaggregation of our long-run relationship, we produce a specification that has a separate term for illiquid assets and also one plus the ratio of liquid to illiquid wealth. Algebraically,

$$\ln W = \ln\left(W_{[\mathrm{LQ}]} + W_{[\mathrm{ILQ}]}\right) = \ln\left[W_{[\mathrm{ILQ}]} \times \left(1 + W_{[\mathrm{LQ}]}/W_{[\mathrm{ILQ}]}\right)\right]. \tag{6}$$

This subsequently allows us to embed our nested long-run equation within our dynamic short run equations.

Nevertheless, the log-linear specification poses problems in terms of estimation of elasticities, as suggested by Muellbauer and Lattimore (1995). The MPC out of wealth, notably, is not immediately derivable without further calculation. This is in addition to the problem of the logarithmic specification not providing nested coefficients. Accordingly, we also tested our specification with a linear approach where we divide the disaggregate components of wealth throughout by income. A non-log specification is derived in Rossi and Visco (1995) and such an estimation approach is presented in Muellbauer (1994). In particular, we include the long-run ratios between the disaggregate components of wealth and income in our consumption function.

$$\Delta \ln C_t = \alpha_0 + \alpha_1 \times \left[\ln C_{t-1} - \beta_1 \times \ln Y_{t-1} - \beta_2 \times \left(\frac{W_{[\mathrm{LQ}]}}{Y}\right)_{t-1} - \beta_3 \times \left(\frac{W_{[\mathrm{ILQ}]}}{Y}\right)_{t-1}\right] + \text{dynamics}. \tag{7}$$

Consequently, the estimated coefficients on wealth with this specification provide a direct measure of the MPC from disaggregate wealth.

IV. Results

(i) Summary

Before going into detail on the individual specifications, we bring together the results for our estimated consumption functions with aggregate and disaggregated wealth in Table 1. In this table we only include those terms which are significant at the 5% level and above. The summarized results clearly illustrate that the illiquid component of wealth is dominant in most of the specifications, with the liquid wealth term typically being insignificant. These results were replicated with SURE analysis in Table 4.[6] As regards the MPC out of wealth, the estimates derived from the aggregate logarithmic

[6] When we used real non-property income in our specification, illiquid assets again dominate (see Byrne and Davis, 2001).

TABLE 1

Consumption and Wealth – Summary

		US	*UK*	*Germany*	*France*	*Italy*	*Canada*	*Japan*
Elasticity from total wealth [2]		0.11	0.13	0.09	0.16	0.10	0.16	0.14
Implied MPC from total wealth		0.06	0.02	0.02	0.03	0.02	0.04	0.01
Elasticity from disaggregate wealth [3]	Liquid	×	×	×	×	×	0.03	×
	Illiquid	0.06	0.11	0.05	0.03	×	0.14	0.11
	Mort.			×				
Elasticity from nested disaggregate wealth [5]	Liquid	×	×	×	×	0.10	0.16	×
	Illiquid	0.07	0.09	0.07	0.02	0.10	0.16	0.10
	Mort.			×				
MPC from total linear wealth [6]		0.06	0.02	0.02	×	0.01	0.04	0.02
MPC from disaggregate linear wealth [7]	Liquid	×	×	×	×	×	0.04	×
	Illiquid	0.06	0.03	×	0.03	0.01	0.04	0.04
	Mort.			×				

Notes: ×: where the coefficient on this variable is insignificant at the 5% level. The marginal propensity to consume (MPC) out of total wealth and is calculated by multiplying the elasticity of total wealth by the ratio of consumption to wealth over the entire sample period. Mort. is mortgage holdings in Germany and is separated from illiquid assets where it dominates the other components. Table numbers are in square brackets.

equation are highly consistent with those from the coefficients in the aggregate linear function, as well as the illiquid wealth term in the linear disaggregated function. The US and Canada have higher MPCs than the EU countries, while the latter are highly consistent with each other. Japan is an exception in that the MPC from illiquid wealth is indicated to be higher than that from total wealth. Also of note is that France does not have a significant wealth effect from total linear wealth but this is overturned in linear disaggregation.

(ii) Aggregate wealth

In terms of the detailed results for aggregate wealth (equation 4) shown in Table 2, where there is a variety of significant dynamic terms, there is a reasonable degree of similarity in the specifications for the G7 countries. The error correction term (α_1) suggests a quarterly adjustment of between 5 and 25% of the current disequilibrium between consumption and its long-run determinants. All error correction terms have *t*-statistics greater than 2, and often greater than 3, consistent with a long-run equilibrium relationship. We also tested for cointegration between the variables to fully justify our choice of long-run specification. Using the Johansen (1988) Trace statistic we find at the

TABLE 2

G7 Consumption Functions with Aggregate Wealth

	US	*UK*	*Germany*	*France*	*Italy*	*Canada*	*Japan*
α_0	0.035	−0.028	0.036	−0.005	−0.088	0.056	
	(1.0)	(0.3)	(1.0)	(0.1)	(0.7)	(0.5)	
$\alpha_1 \times \ln C_{t-1}$	**−0.157**	**−0.147**	**−0.258**	**−0.097**	**−0.060**	**−0.267**	**−0.147**
	(3.3)	**(3.5)**	**(4.7)**	**(3.3)**	**(2.7)**	**(4.7)**	**(2.9)**
$\beta_1 \times \ln Y_{t-1}$	**0.847**	**0.862**	**0.852**	**0.805**	**0.992**	**0.790**	**0.828**
	(15.3)	**(10.4)**	**(15.8)**	**(4.0)**	**(5.3)**	**(15.8)**	**(31.1)**
$\beta_2 \times \ln W_{t-1}$	**0.113**	**0.133**	**0.089**	**0.163**	**0.095**	**0.160**	**0.136**
	(4.3)	**(3.9)**	**(3.3)**	**(2.3)**	**(2.2)**	**(8.0)**	**(6.2)**
$\gamma_1 \times \Delta\ln Y_t$	**0.322**	**0.264**	**0.838**		**0.140**	**0.459**	**0.224**
	(5.8)	**(5.5)**	**(22.1)**		**(2.2)**	**(3.3)**	**(5.2)**
$\gamma_2 \times \Delta\ln Y_{t-1}$		**0.187**					
		(3.8)					
$\gamma_3 \times \Delta\ln C_{t-1}$				**−0.254**	**0.392**		
				(2.9)	**(5.3)**		
$\gamma_4 \times \Delta\ln C_{t-2}$	**0.186**						
	(2.7)						
$\gamma_4 \times \Delta\ln C_{t-3}$							**0.204**
							(3.5)
$\gamma_4 \times \Delta\ln C_{t-4}$				**−0.214**			
				(2.5)			
Wald test *P*-value							
$H_0 : \beta_1+\beta_2=1$	0.197	0.931	0.033	0.826	0.555	0.183	0.000
$\bar{R}^2$	0.55	0.62	0.86	0.28	0.64	0.42	0.66
SE	0.005	0.007	0.007	0.007	0.004	0.007	0.007
SC-LM $\chi^2(4)$	5.5	5.1	2.8	5.7	4.6	7.9	4.1
Norm. $\chi^2(2)$	4.2	4.1	3.5	0.5	1.7	2.1	0.1
Het. $\chi^2(1)$	1.9	0.1	0.2	2.1	2.2	0.1	0.9

Notes: Sample period is 1972Q2–1998Q4. The wealth term is constructed by adding the various components of wealth and then taking the natural logarithm of the aggregate term. Japan does not have a constant in its estimated equation due to non-convergence with a constant. Dynamic terms are deleted using a general to specific methodology. *t*-statistics are in parentheses and where >2 they are in bold.

10% significance level, the level often used in this context, that all countries have evidence of one cointegrating vector, except the US where there are two long-run vectors. This indicates that a linear combination of the variables is not spurious for each of our countries and no additional variables are required to form a cointegrating vector.[7]

All income elasticities are significant and vary in size between 0.79 and 0.99. The wealth terms are significant for all countries and vary in size between 0.09 and 0.16. Tests for homogeneity are accepted in most cases

[7]These results are available from the authors upon request.

although not imposed to be consistent with the disaggregate results which follow. All error diagnostics are statistically acceptable in Table 2.[8,9]

One issue raised by Ludvigson and Steindel (1999) is that results such as these may be dependent on the sample period chosen. Interestingly, there is a reduction in the size of the aggregate wealth effect for the US from 0.10 to 0.07, after we exclude the mid-1990s from the sample period used in Table 2.

$$\Delta \ln C_t = \underset{(t=0.8)}{0.028} - \underset{(3.4)}{0.203} \times (\ln C_{t-1} - \underset{(20.1)}{0.900} \times \ln Y_{t-1} - \underset{(2.9)}{0.071} \times \ln W_{t-1})$$

$$+ \underset{(2.7)}{0.206} \times \Delta \ln C_{t-2} + \underset{(5.6)}{0.367} \times \Delta \ln Y_t$$

1972Q2–1992Q4 $\bar{R}^2 = 0.58$, SE $= 0.005$, SC-LM $\chi^2(4) = 2.7$, Norm. $\chi^2(2) = 0.8$, Het. $\chi^2(1) = 1.6$.

Not only does this imply that asset holdings have become much more sizeable for American consumers in the 1990s, but also the response of consumption to a given percentage increase in asset values has also risen. This result stands in contrast to Ludvigson and Steindel (1999) who report a significant wealth effect from a sample of 44 years beginning in 1953 and a relatively small and insignificant wealth effect when they have a sample period which consists of predominantly the 1990s. As pointed out by Poterba (2000), the latter result may be driven by the span of data used (1986Q1–1997Q4). Our analysis below suggests an alternative explanation, that there is parameter instability owing to the changing importance of the sub-components of net financial wealth.

(iii) Disaggregated wealth

The next step in our investigation involves direct disaggregation of the long-run wealth term into sub-components based on liquidity. The results are presented in Table 3. We allow free estimation of the long-run effects of disaggregate wealth based on the short-run model estimated previously.

[8]Tests include Godfrey's test of residual serial correlation [SC-LM $\chi^2(4)$], the Jarque-Bera normality test [Norm. $\chi^2(2)$] and an LM test for heteroscedasticity [Het. $\chi^2(1)$].

[9]The following dummy variables were included. US: 1980Q2 and 1981Q4 = 1, 1973Q2 and 1974Q4 = 1, 1973Q4 = 1, 1973Q3 = 1. The UK: 1979Q2 = 1 and 1979Q3 = −1, 1980Q4 = 1 and 1981Q1 = 1, between 1990Q2 and 1998Q4 = 1, 1980Q2 = 1. Germany: 1990Q1 = 1, 1978Q2 = 1, 1979Q2 = 1, and 1979Q1 = 1. France: 1974Q4 = 1, 1996Q1 = 1, between 1991Q2 and 1998Q4 = 1. Italy: 1993Q1 = 1, 1992Q4 = 1, from 1985Q2 to 1989Q4 = 1, 1974Q3 = 1. Canada: 1982Q1 = 1, 1974Q4 = 1, from 1981Q2 to 1986Q2 = 1, from 1972Q2 to 1979Q1 = 1, from 1996Q1 to 1998Q4 = 1. Japan: 1974Q1 = 1, and 1997Q2 = 1.

TABLE 3

Consumption Functions with Disaggregate Wealth

	US	*UK*	*Germany*	*France*	*Italy*	*Canada*	*Japan*
$\alpha_1 \times \ln C_{t-1}$	**−0.166**	**−0.119**	**−0.319**	**−0.130**	**−0.044**	**−0.265**	**−0.126**
	(3.2)	**(2.5)**	**(4.9)**	**(4.0)**	**(2.0)**	**(4.5)**	**(2.6)**
$\beta_1 \times \ln Y_{t-1}$	**0.928**	**0.950**	**0.856**	**1.044**	**0.896**	**0.787**	**0.887**
	(16.2)	**(7.1)**	**(19.0)**	**(8.3)**	**(2.0)**	**(15.4)**	**(20.2)**
$\beta_2 \times \ln W_{[\mathrm{LQ}],t-1}$	−0.022	−0.065	0.051	0.026	0.103	**0.025**	−0.015
	(0.8)	(0.5)	(1.1)	(0.3)	(0.7)	**(2.2)**	(0.2)
$\beta_3 \times \ln W_{[\mathrm{ILQ}],t-1}$	**0.058**	**0.111**	**0.051**	**0.025**	0.031	**0.136**	**0.108**
	(2.9)	**(2.4)**	**(2.0)**	**(2.6)**	(1.4)	**(6.9)**	**(3.2)**
$\beta_4 \times \ln W_{[\mathrm{MO}],t-1}$			0.008				
			(0.2)				
$\bar{R}^2$	0.54	0.62	0.86	0.28	0.63	0.41	0.66
SE	0.005	0.007	0.007	0.007	0.004	0.007	0.007
SC-LM $\chi^2(4)$	4.4	5.4	1.5	5.2	5.0	8.2	4.1
Norm. $\chi^2(2)$	3.5	4.0	3.8	0.3	1.3	2.4	0.1
Het. $\chi^2(1)$	2.3	0.1	0.1	3.1	2.7	0.0	1.1

Notes: Sample period is 1972Q2–1998Q4. The wealth terms are $W_{[\mathrm{LQ}],t-1}$ which comprises liquid assets less bank borrowing, $W_{[\mathrm{ILQ}],t-1}$ which is illiquid assets less mortgage debt except for Germany and $W_{[\mathrm{MO}],t-1}$ in the case of Germany is solely mortgage debt. t-Statistics are in parentheses and bold where they are greater than 2. Dynamic terms and constants are consistent with Table 2 and henceforth we focus on the long-run results.

In this section, we constructed liquid and illiquid wealth by adding their respective components and then logging the two wealth variables. This implies that the coefficients on the sub-aggregates should not bear a direct relationship to the coefficients on the aggregate wealth terms contained in Table 2. Despite this caveat that the models are non-nested, the coefficients will nevertheless give us a strong indication of the relative importance of the components of wealth. Alternative methods of restricting the coefficients on the logged components on wealth were tried and do not provide qualitatively different results. Nested results are presented in section IV(iv), as well as in a linear specification in section IV(v).

There is a broad pattern of larger and more significant coefficients for illiquid than liquid assets, except for Italy where both are insignificant. In France, the coefficient on illiquid assets is small but significant. The estimated long-run elasticity on illiquid assets varies from 0.03, for France, to 0.14, for Canada. Note that Table 3 does not present the estimated short-run dynamic as we concentrate on the long-run results.

Recent instability in the parameter estimates for liquid wealth is highlighted in Figure 8, where we plot the rolling regression estimate of the US coefficient on liquid wealth and two SE bands with a window of 40 observations. We can see

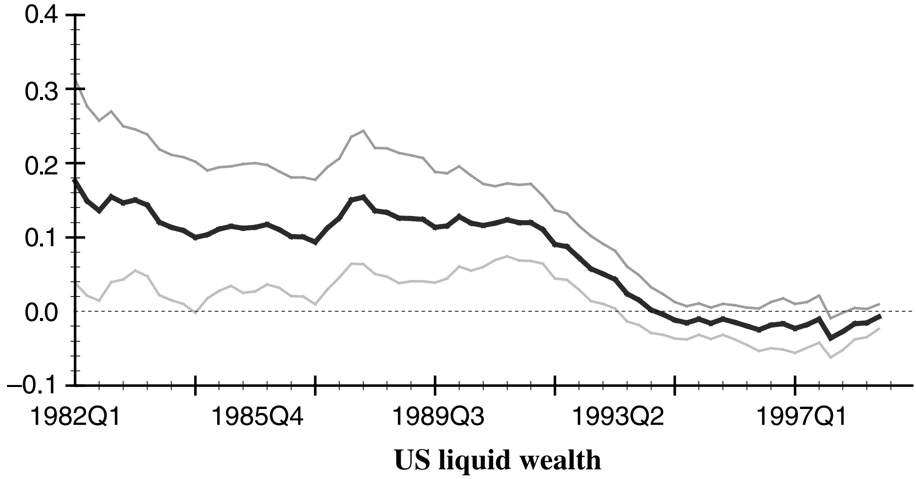

Figure 8. US liquid wealth

quite clearly that there is a fall in the estimated elasticity of liquid assets in the 1990s, with the coefficient becoming not only insignificant but also the wrong sign.[10] This is consistent with our view that the relationship between wealth and income has changed recently, consequent on the removal of liquidity constraints. The later data sample suggests reasons for differences in our results from earlier work. The result also implies that coefficients on aggregate wealth will be unstable as the weights on liquid and illiquid wealth change.

Following Boone *et al*. (1998), we also considered the long-run elasticity of household equity holdings on their own, splitting non-stock market wealth and stock market wealth in our long-run specification for the US. The results are as follows:

$$\begin{aligned}\Delta \ln C_t = {} & \underset{(t=1.7)}{0.095} - \underset{(2.7)}{0.138} \times (\ln C_{t-1} - \underset{(4.8)}{0.731} \times \ln Y_{t-1} - \underset{(1.7)}{0.151} \times \ln \mathrm{NSW}_{t-1} \\ & - \underset{(2.3)}{0.031} \times \ln \mathrm{SW}_{t-1}) + \text{dynamics}\end{aligned}$$

$$\text{1972Q2} - \text{1998Q4}\ \bar{R}^2 = 0.55,\ \text{SE} = 0.005,\ \text{SC-LM}\ \chi^2(4) = 6.7, \text{Norm.}\ \chi^2(2) = 4.0,\ \text{Het.}\ \chi^2(1) = 1.6.$$

Share ownership on its own is significant in the long run, with an elasticity of 0.031. This estimated coefficient is smaller than the coefficient on the

[10]The estimated coefficient in Figure 8 should be divided by the estimated US error correction term in Table 2 to allow the results to be directly comparable.

combined variable of net liquid assets, bonds, life assurance and pension minus mortgage debt – which has an elasticity of 0.151. These results support the view that stock market wealth has a strong effect on consumption and that other illiquid assets are also important.[11]

As a means of checking the robustness of our results, and to aid inference, we proceeded to undertake Zellner's (1962) seemingly unrelated regression (SUR) estimation. If the disturbances across countries are correlated, then there are useful efficiency gains from using SURE. We may expect that the residuals are correlated in this context due to, for example, common shocks to the residuals having possibly a common impact on G7 consumption. Table 4 contains our main results on disaggregate wealth data using SURE. The illiquid wealth coefficient is always larger than the liquid coefficient and significant. As in the previous estimation results, there is a significant effect from Canadian liquid assets. Nevertheless, this is only a quarter of the size of the coefficient on illiquid assets. The parameters on income and wealth from

TABLE 4

SUR Estimation Results with Disaggregate Wealth

	US	*UK*	*Germany*	*France*	*Italy*	*Canada*	*Japan*
$\alpha_1 \times \ln C_{t-1}$	**−0.169**	**−0.121**	**−0.343**	**−0.135**	**−0.052**	**−0.271**	**−0.117**
	(3.6)	**(2.8)**	**(5.6)**	**(4.4)**	**(2.5)**	**(5.1)**	**(2.5)**
$\beta_1 \times \ln Y_{t-1}$	**0.158**	**0.113**	**0.298**	**0.143**	0.051	**0.217**	**0.104**
	(3.2)	**(2.9)**	**(5.1)**	**(3.4)**	(1.5)	**(4.4)**	**(2.8)**
$\beta_2 \times \ln W_{[LQ],t-1}$	−0.004	−0.003	0.015	0.001	0.004	**0.008**	−0.002
	(1.0)	(0.3)	(1.1)	(0.0)	(0.7)	**(2.6)**	(0.3)
$\beta_3 \times \ln W_{[ILQ],t-1}$	**0.009**	**0.011**	**0.016**	**0.003**	**0.002**	**0.035**	**0.012**
	(3.3)	**(3.5)**	**(2.3)**	**(3.1)**	**(2.0)**	**(5.9)**	**(2.6)**
$\beta_4 \times \ln W_{[MO],t-1}$			−0.001				
			(0.1)				
R^2	0.58	0.65	0.87	0.33	0.66	0.46	0.69
SE	0.005	0.007	0.006	0.006	0.004	0.007	0.007
DW	2.0	1.9	2.1	2.0	2.2	2.3	2.0
Het. $\chi^2(1)$	1.8	0.1	0.0	2.6	2.5	0.1	1.0

Notes: The sample period is 1972Q2–1998Q4. The wealth terms are $W_{[LQ],t-1}$ which comprises liquid assets less bank borrowing, $W_{[ILQ],t-1}$ which is illiquid assets less mortgage debt except for Germany and $W_{[MO],t-1}$ in the case of Germany is solely mortgage debt. To make our SUR estimation and single equation results comparable it is necessary to divide the long-run coefficients (β_i) by the error correction term (α_1). *t*-Statistics are in parentheses and bold where they are >2.

[11]We also considered how sensitive our results were to the definition of bond holdings as illiquid assets (see Byrne and Davis, 2001). For Germany it was found that once we defined them as liquid assets the illiquid asset component became insignificant whereas the liquid component became significant. Italy was also found to have an insignificant liquid asset component once we change the definition of bonds. For the other countries we do not find the results change by any quantitative and qualitative amount.

SUR estimation have to be divided by the error correction coefficient before they become comparable with the coefficients in Table 2 and the long-run coefficients as displayed in Table 4 should not be compared across countries without this further transformation. We went on to conduct formal tests using the SUR estimates to examine the equivalence of coefficients across countries, first dividing the long-run relations by the coefficient on the error correction term. We can then implement joint Wald Tests using n individual results, which are distributed as a $\chi^2(n)$ statistics under the null of cross-sectional homogeneity.

Overall the joint null hypothesis that the long-run coefficients on income, liquid and illiquid wealth are equal for the G7 is strongly rejected with a test statistic of $\chi^2(18) = 61.179$ (P-value $= 0.000$). Nevertheless, there is quite a considerable similarity amongst our countries' coefficients as the large majority of probability values accept the null of homogeneity of individual coefficients across countries.[12] Indeed, we accept the null hypothesis that all liquid wealth statistics are equivalent with a test statistic of $\chi^2(6) = 5.369$ (P-value $= 0.497$). This suggests that not only are liquid assets generally unimportant as a component of the impact of wealth on consumption but they also display similar behaviour across countries. We then tested the joint hypothesis that the coefficients on illiquid wealth were homogenous and this was accepted at the 5% significance level with a test statistic of $\chi^2(5) = 10.317$ (P-value $= 0.068$), although we excluded Canada as it appears to be an outlier in this part of the analysis. Additionally, we tested whether this group of countries' estimated liquid wealth coefficients could be restricted to zero and this was accepted with a test statistic of $\chi^2(6) = 2.757$ (P-value $= 0.839$). This supports our major hypothesis. Tests were also undertaken for similar behaviour for the UK and the current members of the Euro Area in our sample. For the individual coefficient on illiquid wealth we accept the null of homogeneity with a test statistic of $\chi^2(3) = 3.959$ (P-value $= 0.266$).[13]

(iv) Nested specification

The results with the nested specification derived in equation (6) are presented in Table 5. All equations are well specified and pass residual congruency tests at the 5% significance level. Examining a restricted long-run specification

[12] See Byrne and Davis (2001).

[13] The Wald Test results on poolability of European countries were sensitive to the inclusion of liquid wealth in the long run. This suggests that the similarities between Germany and the other countries are not total. Nevertheless, the estimated coefficients for illiquid wealth from SURE presented in Table 4 are neither qualitatively nor quantitatively different when we exclude liquid wealth in the long-run.

TABLE 5

Consumption Functions with Nested Disaggregate Wealth

	US	*UK*	*Germany*	*France*	*Italy*	*Canada*	*Japan*
$\alpha_1 \times \ln C_{t-1}$	**−0.152**	**−0.134**	**−0.283**	**−0.133**	**−0.060**	**−0.267**	**−0.134**
	(3.2)	**(3.4)**	**(5.0)**	**(4.2)**	**(2.7)**	**(4.7)**	**(3.6)**
$\beta_1 \times \ln Y_{t-1}$	**0.890**	**0.900**	**0.865**	**1.072**	**0.992**	**0.790**	**0.877**
	(18.7)	**(10.8)**	**(19.9)**	**(12.8)**	**(5.3)**	**(15.8)**	**(49.9)**
$\beta_2 \times \ln W_{[\mathrm{ILQ}],t-1}$	**0.070**	**0.094**	**0.068**	**0.023**			**0.101**
	(4.2)	**(3.4)**	**(3.8)**	**(3.1)**			**(6.5)**
$\beta_2 \times \ln W_{t-1}$					**0.095**	**0.160**	
					(2.2)	**(8.0)**	
Wald Test *P*-values							
$H_0 : \beta_1+\beta_2=1$	0.210	0.927	0.012	0.235	0.555	0.183	0.000
$\bar{R}^2$	0.55	0.62	0.86	0.28	0.64	0.42	0.67
SE	0.005	0.007	0.007	0.007	0.004	0.007	0.007
LM-SC $\chi^2(4)$	5.2	5.6	1.6	5.1	4.6	7.9	3.9
Norm. $\chi^2(2)$	3.7	4.1	3.8	0.3	1.7	2.1	0.1
Het. $\chi^2(1)$	2.1	0.1	0.2	3.6	2.2	0.1	1.0

Notes: Sample period is 1972Q2–1998Q4. The wealth terms are illiquid assets less mortgage debt $W_{[\mathrm{ILQ}],t-1}$ except for Germany. In the case of Germany, we have illiquid wealth excluding mortgage debt. W_{t-1} comprises net total liquid and illiquid assets. *P*-values < 0.05 reject the null for the Wald tests. *t*-Statistics are in parentheses and bold where they are >2.

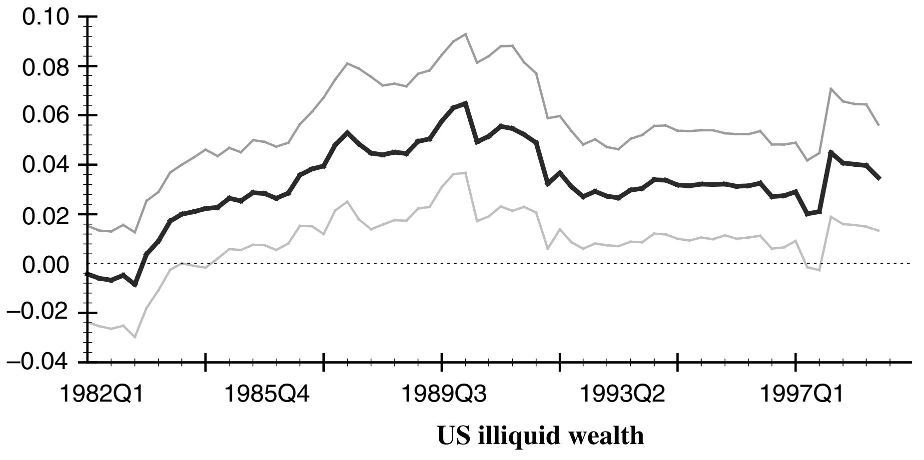

Figure 9. US illiquid wealth

(removing highly insignificant coefficients one by one) we find that for the US, UK, Germany, France and Japan there is evidence of a significant effect from illiquid wealth while the implicit effect from liquid wealth is zero due to an insignificant *t*-statistic. This replicates our previous results on

TABLE 6

Consumption with Linear Wealth

	US	*UK*	*Germany*	*France*	*Italy*	*Canada*	*Japan*
α_0	0.020	−0.007	0.018	0.012	−0.147	0.076	
	(0.6)	(0.1)	(0.5)	(0.1)	(1.1)	(0.7)	
$\alpha_1 \times \ln C_{t-1}$	**−0.172**	**−0.140**	**−0.245**	**−0.083**	**−0.069**	**−0.222**	**−0.092**
	(3.5)	**(3.5)**	**(4.5)**	**(2.8)**	**(3.0)**	**(3.8)**	**(2.3)**
$\beta_1 \times \ln Y_{t-1}$	**0.967**	**0.993**	**0.956**	**0.957**	**1.143**	**0.943**	**0.980**
	(36.5)	**(16.3)**	**(35.3)**	**(5.4)**	**(10.8)**	**(19.8)**	**(176.2)**
$\beta_2 \times (W/Y)_{t-1}$	**0.060**	**0.018**	**0.023**	0.029	**0.013**	**0.043**	**0.015**
	(4.7)	**(3.8)**	**(2.9)**	(1.8)	**(2.7)**	**(6.3)**	**(2.9)**
$\bar{R}^2$	0.56	0.63	0.85	0.26	0.65	0.36	0.65
SE	0.005	0.007	0.007	0.007	0.004	0.007	0.007
SC-LM $\chi^2(4)$	4.6	3.5	2.2	6.4	4.6	7.9	4.4
Norm. $\chi^2(2)$	3.1	4.4	4.1	0.2	1.7	0.7	0.0
Het. $\chi^2(1)$	2.0	0.1	0.1	1.3	2.2	0.3	1.0

Notes: Sample period is 1972Q2–1998Q4. The wealth term is constructed by adding the various components of wealth and then dividing by income. Dynamic terms are deleted where insignificant using a general to specific methodology. *t*-Statistics are in parentheses and bold where they are >2.

disaggregation in Table 3 and implies that removal of liquid wealth does not radically change the results. The results for Italy and Canada suggest an important role for both kinds of wealth holdings, and hence a combined wealth term was included in this long-run specification.[14]

For the US we also estimated the nested model presented in Table 5 using rolling regressions with 40 observations. The estimated illiquid wealth elasticity and two SE bands are given in Figure 9. Since the mid-1980s the elasticity is broadly stable, consistently greater than zero and by more than two SEs. This highlights the recent importance of illiquid wealth in the consumption function for the US. Note that the estimated coefficient for the US should be divided by US error correction term above, before it is of comparable size to the illiquid wealth coefficient in Table 5.

(v) Linear wealth

Following the discussion in section III, there may be some difficulty in interpreting the disaggregate elasticities from a log specification. As a robustness check we undertook linear estimation of our consumption function. Aggregate wealth results are provided in Table 6 and these replicate previous

[14]It is worthwhile comparing the disaggregate results from Tables 3 and 5. This suggests excluding liquid wealth from the long-run relationship does not substantially alter the importance of illiquid wealth for most countries.

TABLE 7

Consumption with Linear Disaggregate Wealth

	US	*UK*	*Germany*	*France*	*Italy*	*Canada*	*Japan*
α_0	0.021	0.004	−0.014	0.022	−0.109	0.076	
	(0.6)	(0.1)	(0.4)	(0.2)	(0.7)	(0.7)	
$\alpha_1 \times \ln C_{t-1}$	**−0.175**	**−0.113**	**−0.339**	**−0.105**	**−0.067**	**−0.223**	**−0.096**
	(3.4)	**(2.5)**	**(5.1)**	**(3.3)**	**(2.8)**	**(3.6)**	**(2.4)**
$\beta_1 \times \ln Y_{t-1}$	**0.968**	**0.995**	**0.970**	**0.970**	**1.100**	**0.944**	**0.983**
	(37.0)	**(13.1)**	**(15.8)**	**(7.2)**	**(6.9)**	**(19.7)**	**(153.2)**
$\beta_2 \times (W_{[LQ]}/Y)_{t-1}$	0.045	−0.030	0.024	−0.038	0.025	**0.044**	−0.008
	(0.6)	(0.6)	(1.4)	(1.2)	(0.9)	**(2.1)**	(0.5)
$\beta_3 \times (W_{[ILQ]}/Y)_{t-1}$	**0.058**	**0.026**	0.012	**0.025**	**0.014**	**0.042**	**0.036**
	(3.4)	**(2.4)**	(1.9)	**(2.1)**	**(2.6)**	**(5.3)**	**(2.7)**
$\beta_4 \times (W_{[MO]}/Y)_{t-1}$			0.000				
			(0.1)				
$\bar{R}^2$	0.56	0.63	0.86	0.27	0.65	0.36	0.66
SE	0.005	0.007	0.007	0.007	0.004	0.007	0.007
SC-LM $\chi^2(4)$	4.4	3.1	1.1	4.8	4.7	7.9	4.5
Norm. $\chi^2(2)$	3.1	4.3	4.1	0.4	1.7	0.7	0.1
Het. $\chi^2(1)$	2.1	0.2	0.1	4.6	2.3	0.3	1.2

Notes: Sample period is 1972Q2–1998Q4. The wealth term is disaggregated by dividing the various components of wealth based on liquidity by income. Dynamic terms are deleted where insignificant using a general to specific methodology. *t*-Statistics are in parentheses and bold where they are >2.

results using a log specification. There is little to differentiate between the results presented in Tables 2 and 6 in terms of $\bar{R}^2$ and regression SEs. Only in the case of France is there any indication that the linear aggregate wealth effect is insignificant at the 5% level.

Table 7 suggests that there is a significant long-run impact from illiquid wealth for all our countries (although in Germany it only exists at the 10% significance level). The German results are sensitive to the estimation of a separate coefficient on mortgage wealth in the long-run relationship (indeed both disaggregate wealth terms become significant when we include mortgage debt in illiquid assets).[15] In the case of Italy, the log consumption function suggests that illiquid wealth is insignificant in determining consumption (t-statistic $= 1.4$ in Table 3). However, we find that the illiquid wealth term is significant in the disaggregate linear specification. For France we find that although a net wealth effect is insignificant in the long-run, the illiquid wealth effect is significant. The MPCs derived from the coefficients in Tables 6 and 7

[15] In such circumstance the MPC on German liquid wealth is 0.037 ($t = 4.3$) and on illiquid wealth 0.014 ($t = 2.3$).

are consistent between aggregate and illiquid wealth. However, a corollary is that the aggregate wealth term may be unstable when the ratio of liquid to illiquid wealth changes.

V. Conclusion

There is considerable evidence from the results presented in this paper that the estimated coefficients on the various sub-components of financial wealth differ significantly for the G7 consumption functions, thus implying that indiscriminate aggregation of wealth is inappropriate. Furthermore, there is evidence that illiquid wealth typically dominates the effect of conventional liquid assets, which is insignificantly different from zero with the latest data. The different effects from the components of wealth may be seen as justified in the context of a liberalized financial system, where consumption is focused on life cycle considerations and lifetime wealth is held increasingly in securities and institutional investment. In contrast, liquid assets are held largely for transaction purposes and changes in their volume are less strongly related to consumption than previously was the case.

We concentrated our analysis on a log specification, but as we checked the robustness of our results using a nested framework and with a linear specification; this strengthens our conclusion of the importance of illiquid wealth for consumption. Also, the linear approach allows us to directly assess the impact of illiquid wealth on consumption. In particular, we find that disaggregate wealth consistently dominates for the US, UK, France and Japan and a 10% fall in illiquid wealth holdings leads to a fall in consumption of 0.6, 0.3, 0.3 and 0.4%, respectively. Across specifications for Canada and, to a lesser extent, for Italy there is an equivalent effect from different kinds of wealth, where a 10% fall in net wealth results in a 0.4 and 0.1% fall in consumption for these two countries respectively. For Germany the results were sensitive to the actual definition of liquid assets, with bond holdings and mortgage debt influential.

An important implication of these results is that conventional studies of wealth effects on consumption may give biased and potentially unstable estimates as the share of illiquid relative to liquid assets changes. This could lead to important errors in predictions. For example, if illiquid wealth falls and liquid wealth rises to offset it (perhaps due to precautionary accumulation), an aggregate function would predict no change in consumption while the disaggregate function would predict a decline. This difficulty could be of particular relevance for the US in the short term, and also in the longer term for a wider range of countries as financial market liberalization proceeds, securities markets develop and portfolios resemble those of the US to a greater degree. Moreover, the process of ageing, which is already underway, is also driving portfolios towards holding more illiquid assets.

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