Credit, Housing Collateral and Consumption: Evidence from the UK, Japan and the US

Janine Aron
John V. Duca
John Muellbauer
Keiko Murata
Anthony Murphy

Research Department

Working Paper 1002

Federal Reserve Bank of Dallas

Credit, Housing Collateral and Consumption: Evidence from the UK, Japan and the U.S.

Janine Aron, *Department of Economics, University of Oxford, UK*

John V. Duca, *Federal Reserve Bank of Dallas & Southern Methodist University, US*

John Muellbauer, *Nuffield College, University of Oxford, UK*

Keiko Murata, *Tokyo Metropolitan University, Japan*

Anthony Murphy, *Federal Reserve Bank of Dallas, US*

May 2010, Revised May 2011

Abstract: The consumption behaviour of UK, U.S. and Japanese households is examined and compared using a modern Ando-Modigliani style consumption function. The models incorporate income growth expectations, income uncertainty, housing collateral and other credit effects. These models therefore capture important parts of the financial accelerator. The evidence is that credit availability for UK and U.S. but not Japanese households has undergone large shifts since 1980. The average consumption-to-income ratio shifted up in the UK and U.S. as mortgage down-payment constraints eased and as the collateral role of housing wealth was enhanced by financial innovations, such as home equity loans. The estimated housing collateral effect is similar in the U.S. and UK, while land prices in Japan still have a negative effect on consumer spending. Together with evidence for negative real interest rate effects in the UK and U.S. and positive ones in Japan, this suggests important differences in the transmission of monetary and credit shocks in Japan versus the U.S., UK and other credit-liberalized economies.

Keywords Consumption, credit conditions, housing collateral, housing wealth.

JEL Codes E21, E32, E44, E51

[*]We thank Marshall Reinsdorf, Erwin Diewert and two anonymous referees for comments and suggestions. We also thank David Luttrell and Jessica Renier for research assistance. An earlier version was presented at the 2010 American Economic Association meetings, 2010 meeting of the Swiss Society for Financial Market Research, 2010 meeting of the International Association for Research on Income and Wealth (IARIW) and at Statistics Norway in 2009. We are grateful for all of the comments received. The views expressed are those of the authors and do not necessarily reflect those of the Federal Reserve Bank of Dallas, or the Board of Governors of the Federal Reserve System. Any remaining errors are our own.

1. Introduction

The global economic crisis of 2007-9 had its origins in a credit crisis, at the heart of which is asymmetric information between lenders and borrowers, most simply the fear or risk that lenders face about the ability and willingness of borrowers to service debt (see Duca *et al.,* 2010a, Stiglitz and Weiss (1981)). It is now generally accepted that the household credit channel played an important part in the boom preceding the crisis, as well as in accentuating the crisis via a financial accelerator that amplified the shocks emanating from housing and mortgage markets.

The financial accelerator is often neglected in the econometric models that, for the last decade, have been popular with central banks and main-stream macroeconomists. Many dynamic stochastic general equilibrium (DSGE) modellers focused on building macro models with rational expectations and micro-foundations that could generate nominal rigidities by incorporating 'New Keynesian' frictions, primarily price stickiness and adjustment costs. Practical modelling issues resulted in the widespread adoption of micro foundations that too often ignored the asymmetric information revolution of the 1970s and 1980s. These models also neglected the information from the flow of funds and balance sheets, now receiving far more attention from central banks (Gonzalez-Paramo (2009), Eichner *et al.*(2010)).

This paper estimates consumption functions that modernize the approach of Modigliani and Brumberg (1954, 1980) and Ando and Modigliani (1963) for three major economies – the UK, U.S. and Japan – which explicitly incorporate income expectations, uncertainty and credit channel influences. We show that, consistent with theory, these credit channel effects differ across countries and over time. The estimated models take account of household balance sheet data, addressing important measurement issues both for income and for wealth.

Early work attributed much of the fall in the UK household saving rate to credit market liberalisation and increases in house prices and in the 'spendability' of housing and other illiquid wealth (Muellbauer and Murphy, 1989, 1990).[1] Muellbauer and Lattimore (1995) review further research and lay out the foundations of a solved-out consumption function encompassing the classical life-cycle/permanent income hypothesis (LC/PIH) and a credit channel.[2]

---

[1] Elsewhere we have modelled the roles of credit supply and extrapolative expectations in explaining house prices (Muellbauer and Murphy, 1997; Cameron *et al*., 2006; Duca, Muellbauer and Murphy, 2011a).
[2] Although this paper notably influenced the consumption function of the Federal Reserve's FRB-US model for which Brayton *et al*. (1997) carefully modelled expectations, the academic literature in macroeconomics has been dominated by approaches based on Euler equations with representative agents, as in many DSGE models.

This research implies that housing collateral effects on consumption can differ across countries and shift over time due to credit market liberalization. With imperfect capital markets, both the cost and availability of borrowing are affected by agency costs which give rise to down-payment constraints in housing markets. Japelli and Pagano (1994), Engelhardt (1996) and others demonstrate that mortgage down-payment constraints generate an economically significant motive to save. In countries with limited access to consumer and/or mortgage credit, such as Italy and Japan, higher home prices can induce higher saving for down-payments, thereby generating negative housing collateral / 'wealth' effects.

Credit market liberalization creates positive housing 'wealth' effects on consumption by reducing the down-payment constraint and increasing the collateral effect, as found in the UK and U.S. First, credit liberalization lowers the down-payments required of first-time home buyers. Second, it provides those homeowners facing constraints in unsecured credit markets with an improved ability to borrow against housing equity at lower interest rates. This suggests that, *ceteris paribus*, the aggregate household saving ratio is likely to fall as credit markets are liberalized.

It is challenging to measure exogenous shifts in the credit supply function facing households. For the UK, the most systematic estimates are in Fernandez-Corugedo and Muellbauer (2006), who jointly model ten mortgage and other consumer credit indicators - controlling for standard economic and demographic variables, such as incomes, asset prices, interest rates, risk indicators and the age composition of the population - to extract a common latent variable.[3] The resulting general credit conditions index (*GCCI*) is interpreted as a scalar measure of the exogenous shift in credit supply facing UK households.

In models of aggregate UK consumption, we find that *GCCI* significantly affects the log consumption-to-income ratio (approximately the negative of the savings ratio), controlling for income, income expectations, changes in the unemployment rate, interest rates and the composition of household portfolios. Income growth expectations are modelled using a simple but robust income forecasting equation. By interacting *GCCI* with other variables such as housing wealth, other parameters are found to shift with credit market liberalisation, in line with theory. Moreover, by including *GCC*I effects, the other model parameters become stable over the four decades from 1967 to 2005, with co-integration tests easily passed.

We adopt a similar approach to modelling aggregate U.S. consumption, except that we use actual household survey date when modelling income expectations. The long history

[3] The approach parallels the MIMIC approach to estimating latent variables of Goldberger (1974) and Joreskog and Goldberger (1975), and Staiger *et al*.'s (1997) latent measure of the natural rate of unemployment.

of the Thomson Reuters / University of Michigan survey data allows our U.S. income forecasting equation to be based more directly on household evidence than is possible for the UK and Japan. For the U.S. consumption function, as for the UK, there is strong evidence for structural shifts in the consumption-to-income ratio, conditional on income, income growth expectations, interest rates, unemployment changes and household portfolio holdings. These shifts can be plausibly matched to known changes in credit market architecture, particularly since the early 1980s (e.g. Dynan, Elmendorf and Sichel, 2006). Our estimates suggest that co-integration of aggregate U.S. consumption, income, and wealth holdings for the last forty years would be hard to find without accounting for credit market shifts.

In Japan, by contrast, credit market liberalization for households since the mid-1970s appears to have been largely absent. Controlling for income growth expectations using a separate income forecasting equation, no evidence is found of any parameter shifts over the period 1961 to 2008.[4] Consistent with this absence of credit market liberalization, the housing 'wealth' or collateral effect is *negative* for Japan, in contrast to the UK and U.S. Also, given the preponderance of liquid assets held by Japanese households, the aggregate effect of a rise in short-term real interest rates is *positive*, again differing from the UK and U.S.

Although the consumption function equations in this paper are necessarily partial equilibrium or conditional in nature, they have important short and medium term policy implications. This is particularly pertinent to the economic crisis of 2007-9. For example, our UK consumption function made it possible to predict by mid-2008 that the UK would be in recession in the second half of 2008 - given falling house prices, lower real incomes, less credit availability, rising unemployment and lower stock market wealth. The earlier Bank of England view of a weak and unstable relationship between house prices and consumption, probably contributed to some members of the Monetary Policy Committee voting for a *rise* in interest rates as late as August 2008, and the Committee's initially slow policy response to the economic downturn in September and October.

In contrast, while the Federal Reserve Board's U.S. macro model may not have fully accounted for shifts in credit conditions and the short-run consumption response to housing collateral or wealth, it incorporates powerful housing and stock market 'wealth' effects. Along with a greater appreciation of the financial accelerator amongst U.S. policymakers, this may have contributed to an early and decisive monetary policy response to the crisis.

[4] This is confirmed by estimates of a constant-parameter equation for household debt. Co-integration tests are satisfactory, and instrumental variables estimates suggest the absence of endogeneity bias. The results in this paper confirm the earlier results in Muellbauer and Murata (2011) using an extended sample and a better income forecasting model, *inter alia*.

Our results for Japan also explain why the household component of the monetary transmission channel is far weaker there than in the UK or U.S. Had this been better understood in 2001-2004, U.S. monetary policy may have been less concerned about the risk of a Japan-style 'lost decade' over 2000-09. Inter alia, Leamer (2007) and Taylor (2007) argue that, as result, the federal funds rate was kept too low for too long. Perhaps more importantly, there was an unsustainable liberalisation of the mortgage market fuelling an unexpectedly strong (albeit unsustainable) credit, housing, and consumption boom, whose collapse is still playing out.

Although macroeconomic research using aggregate time-series and balance sheet data has been less fashionable in the last two decades, recent events underline its policy relevance. Indeed, Kohn (2008), Goodhart and Hofmann (2008) and others have called for the design and implementation of new central bank macro models with more realistic features. This paper is one step in this direction.

2. The Consumption Function

This section begins by demonstrating the weakness of the housing wealth effect in the classical life-cycle model of consumption. We then discuss the housing credit effect on consumption, operating via lower mortgage down-payments and the increased collateral role of housing. We briefly consider aggregation issues, including changes in demographic structure, before presenting an estimable and realistic, solved-out consumption function, incorporating income expectations and uncertainty as well as credit channel effects.

2.1 Housing wealth effects.

Many argue that there is no housing 'wealth' effect in the standard life cycle model and that any apparent effect arises because housing wealth proxies for omitted expectations of future income (e.g. King (1990) and Pagano (1990)). The lack of a strong positive housing wealth effect in standard frameworks can be shown using a stylised life cycle model of consumption. Let c = real non-housing consumption, p^h = relative price of housing, H = stock of housing, δ = rate of depreciation of housing, r = real interest rate, y^p = permanent real non-property income and A = real financial wealth. In each period, the consumer maximises life-cycle utility defined on the flows of consumption c and on the stock of housing H.

If expected relative house prices p^h and the real interest rate r are constant, then the multi-period, inter-temporal optimization problem reduces to a two-good, single period optimization problem with budget constraint:

$$c + p^{\mathrm{h}}(r+\delta)H = y^{p} + r(A_0 + p^{\mathrm{h}}H_0) \qquad (2.1)$$

where $(r+\delta)H$ = housing services and $p^{\mathrm{h}}(r+\delta)$ = real user cost of housing. We are interested in the effects of change in p^{h} on a constant price index of consumption like the one in the national accounts. This includes imputed rent on housing. Holding base prices fixed and differentiating equation (2.1) with respect to p^{h}, we find:

$$\partial\left(c + p^{h}(r+\delta)H\right)\Big/\partial p^{h} = rH_0 - (r+\delta)H \qquad (2.2)$$

But with $H \approx H_0$, the right hand side of equation (2.2) is negative since δ is positive. This point was overlooked in the classic work by Modigliani and Brumberg (1954, 1980), Friedman (1957, 1963) and Ando and Modigliani (1963).

Of course, the simple implications of equation (2.2) are liable to be somewhat modified in models with finite lives and transactions costs and depend on how well imputed rent is measured in the national accounts. Nevertheless, it is hard to generate a substantial aggregate housing wealth effect from classical life-cycle permanent income theory (e.g. Buiter, 2008). For non-housing consumption, a modest positive effect is likely when there are no credit constraints (Muellbauer, 2007, p. 272).

2.2 The Household Credit Channel

This section reviews how access to credit interacts with house prices, interest rates and income growth expectations to affect consumption and how a change in access to credit alters consumption through two main channels. The first channel concerns the mortgage down-payment constraint. In many countries, mortgage debt is the dominant household liability. Mortgage suppliers set upper limits on loan-to-income and loan-to-value (LTV) ratios to reduce default risk. This forces young households to save for the initial deposit, i.e. to consume less than income, the difference depending on the ratio of house prices to income and on the maximum LTV ratio on mortgages. An easing of credit constraints, in the form of higher LTV ratios, will raise the consumption of these households relative to their incomes (see Japelli and Pagano (1994), Deaton (1999) and Engelhardt (1996)). [5]

Now consider the impact on consumption of higher house prices via the operation of the down-payment constraint. With weak access to credit, potential first-time buyers tend to save more as house prices rise unless they give up on purchasing a house. Increased access to credit will weaken the resulting negative effect on consumption.

The second credit channel operates via housing collateral. In a number of countries, the relaxation of rules and spread of competition has made it easier to obtain loans backed by housing-equity (see Poterba and Manchester, 1989). A rise in house prices then makes it possible to increase debt or to refinance other debt at lower interest rates. Effectively, the liberalization of credit conditions increases the 'spendability' or liquidity of previously illiquid housing wealth. The greater liquidity of housing wealth, along with easier access to credit, gives housing wealth a buffer stock role.

Overall, if existing home-owners have only limited access to home equity loans, the effect on their consumption of higher house prices will be small, when combining the down-payment and collateral mechanisms into a life-cycle framework For example, equation (2.2) implies that existing owners, who are not credit constrained and whose behaviour is governed by the life-cycle model outlined above, will display a small negative response to a permanent increase in real house prices unless they downsize to cheaper accommodation. The same equation, with $H_0 = 0$, also implies that renters will save more when house prices are higher. Hence, the aggregate consumption effect of a rise in real house prices is likely to be negative when access to credit is restricted. The effect then switches from negative to positive as access to credit expands.

In countries like the UK where floating rate debt is important, indebted households are subject to short-term cash flow shocks when nominal interest rates change (see Jackman and Sutton, 1982). Their consumption is thus likely to be influenced by changes in the debt service burden, which can be well tracked by proportional changes in the nominal interest rate weighted by the debt-to-income ratio. Better access to collateral reduces the impact of such changes, as households with positive net equity can more easily refinance to protect their cash flows against rises in nominal interest rates. The negative effect of nominal interest rate changes should thus weaken with credit market liberalization, but become larger in a credit crunch. Finally, greater access to unsecured credit should increase the role of inter-

[5] Note that most potential first-time home-buyers, who are saving for a deposit on a house, are not credit-constrained in the sense of being unable to smooth consumption. The savings they accumulate for the deposit can be adjusted in anticipation of short-term income fluctuations and in response to changes in real interest rates.

temporal substitution, enhancing the role of income growth expectations and, on net, making the real interest rate effect more negative.

2.3 Aggregation and Demographic Effects

In the stylized life-cycle consumption function, with permanent income proxied by current income, micro-level consumption is linear in assets and non-property income:

$$c_{it} = \gamma_{it} A_{it-1} + \lambda_{it} y_{it} \qquad (2.3)$$

where γ_{it} and λ_{it} vary by age. Hence average per capita consumption is:

$$c_t = \tfrac{1}{N_t}\sum_i c_{it} = \left(\frac{\sum_i \gamma_i A_{it-1}}{\sum_i A_{it-1}}\right)\tfrac{1}{N_t}\sum_i A_{it-1} + \left(\frac{\sum_i \lambda_i y_{it}}{\sum_i y_{it}}\right)\tfrac{1}{N_t}\sum_i y_{it} \equiv \tilde{\gamma}_t A_{t-1} + \tilde{\lambda}_t y_t \qquad (2.4)$$

Thus, the consumption function $c_t = \tilde{\gamma}_t A_{t-1} + \tilde{\lambda}_t y_t$ will have non-constant $\tilde{\gamma}$ and $\tilde{\lambda}$ parameters which depend on demography and the distribution of income and wealth by demographic groups. In the long run, Gokhale, Kotlikoff and Sabelhaus (1996) argue that shifts in $\tilde{\gamma}$ and A by age account for some of the secular decline in U.S. saving rate. Similar arguments are common in Japan. However, cross-section evidence suggests that $\tilde{\gamma}$ and $\tilde{\lambda}$ may vary less across households than text book models imply because of uncertainty about time of death (e.g., Bosworth *et al*. (1991) and Murata (1999, ch. 8)).

In practice, $\tilde{\gamma}$ and $\tilde{\lambda}$ evolve slowly with life expectancy and the age distribution of the population, as well as the distribution of *y* and *A* by age. Murata (1999, ch. 5), using calibrations consistent with micro data from the Japanese Family Saving Survey, finds that aggregate consumption models in which $\tilde{\gamma}$ and $\tilde{\lambda}$ are constant have very similar implications and fits to models where they evolve according to the survey data. Furthermore, as households make long-run portfolio decisions, the level and composition of assets is likely to reflect the demographic evolution, implying that shifts in $\tilde{\gamma}$ and $\tilde{\lambda}$ due to demographic change have a less direct impact on consumption. Accordingly, we simplify by assuming that $\tilde{\gamma}$ and $\tilde{\lambda}$ are constant in the next section.

2.4 A Solved Out Consumption Function

Ando-Modigliani-Brumberg or Friedman style consumption functions require an income forecasting model to generate permanent non-property income. Unlike the Euler equation (Hall, 1978), they do not ignore long-run information on income and assets. As a result, the solved out consumption function has advantages for policy modelling and forecasting. When the real interest rate is constant, the basic aggregate life-cycle/permanent income consumption function has the form:

$$c_t = \tilde{\gamma} A_{t-1} + \tilde{\lambda} y_t^P \tag{2.5}$$

where c is real per capita consumption, y^p is permanent real per capita non-property income and A is the real per capita level of net wealth (e.g., Deaton and Muellbauer, 1980, ch. 4.2). This equation has a basic robustness feature missing in the Euler equation. Euler equations require well-informed households to continuously and optimally choose between current and future consumption. Strong multi-country evidence against this fundamental prediction is found by Campbell and Mankiw (1989, 1991) *inter alia*. Equation (2.5) is less restrictive since it is consistent with a fairly rudimentary comprehension of life-cycle budget constraints. Any household with some notion of wanting to sustain consumption will realize that not all of assets can be spent now without damaging future consumption, and that future income affects sustainable consumption. As we shall see, practical applications of equation (2.5) capture these basic ideas.

Dividing equation (2.5) by y and a little manipulation gives:

$$\frac{c_t}{y_t} = \tilde{\lambda}\left(\frac{\tilde{\gamma}}{\tilde{\lambda}}\frac{A_{t-1}}{y_t} + 1 + \frac{y_t^P - y_t}{y_t}\right) \tag{2.6}$$

The right-hand side of equation (2.6) has the form 1 + x, where x is usually a fairly small number.[6] We can then take logs, use the fact that $\ln(1+x) \approx x$ and the approximation $\left(y^P - y\right)/y \approx \ln\left(y^p/y\right)$ to obtain:

$$\ln c_t = \alpha_0 + \ln y_t + \gamma\, A_{t-1}/y_t + \ln\left(y_t^P/y_t\right) \tag{2.7}$$

where $\gamma = \tilde{\gamma}/\tilde{\lambda}$ and $\alpha_0 = \tilde{\lambda}$. Thus, α_0 embodies the evolving distribution of demography and income, while γ embodies the evolving relative influences of the distributions of assets, income and demography, with the last factor proxied by variables such as the population proportions in different age groups.[7] The log ratio of permanent to current income reflects expected income growth and can be proxied using forecasted income using the approximation:

$$\ln\left(y_t^p/y_t\right) \approx E_t\left(\sum_{s=1}^{k}\delta^{s-1}\ln y_{t+s}\right)\Big/\left(\sum_{s=1}^{k}\delta^{s-1}\right) - \ln y_t \equiv E_t \ln yperm_t - \ln y_t \tag{2.8}$$

[6] Where x is not 'small', a second order approximation based on $\ln(1+x) \approx x - \tfrac{1}{2}x^2$ is easy to apply.

[7] However, such variables are typically integrated of order 2 and robust estimates of their effects are therefore not possible, though calibration can be attempted as discussed above.

where $\ln yperm_t - \ln y_t$ is a weighted moving average of forward-looking income growth rates (see Campbell, 1987).

The static consumption function (2.7) can be made dynamic by introducing habits or adjustment costs, resulting in a partial adjustment form of equation (2.7), as derived by Muellbauer (1988) *inter alia*. Allowing for probabilistic income expectations suggests adding a measure of income uncertainty, θ_t, and a risk premium in the discount factor δ in the expected income growth term, $E_t \ln yperm_t - \ln y_t$. As many argue, households are likely to discount the future by more than the real rate of interest (e.g. Hayashi, 1985). If real interest rates are variable, standard theory suggests the real interest rate r_t enters (2.7), with the usual interpretation of inter-temporal substitution and income effects.

These considerations lead to the following generalisation of the canonical rational expectations, permanent income hypothesis (REPIH) model in equation (2.7):

$$\Delta \ln c_t \approx \lambda\left(\alpha_0 - \alpha_1 r_t - \alpha_2 \theta_t + \ln y_t + \alpha_3 (E_t \ln yperm_t - \ln y_t) + \gamma A_{t-1}/y_t - \ln c_{t-1}\right) + \varepsilon_t \qquad (2.9)$$

where λ measures the speed of adjustment. In principle, the coefficients α_3 and γ should depend upon the real interest rate r_t. They may also depend on θ_t, since discount factors applied to expected incomes will increase with income uncertainty, as Skinner (1988), Zeldes (1989), Kimball (1990) and Carroll (1997, 2001) have emphasized. For simplicity we will suppress this complication and the associated potential non-linearities.[8]

In practice, there are several reasons why income growth expectations embodied in $E_t \ln yperm_t - \ln y_t$ are likely to reflect a limited horizon. With aggregate data, it is difficult to forecast income beyond three or so years. Indeed, widely used time series models usually lose most of their forecasting power by that horizon. This suggests that the log of income in the more distant future is best forecast in practice by near-term log income plus a constant. Furthermore, short horizons are suggested if households anticipate future credit constraints according to the buffer-stock theory of saving (see Deaton 1991, 1992). Precautionary behaviour with uncertain 'worst case scenarios' also generates buffer-stock saving, as in Carroll (2001), who argues that plausible calibrations of micro-behaviour can give a practical income forecasting horizon of about three years – the same horizon as Friedman (1957, 1963) suggested.

[8]In principle, the consumption function should include effects arising from aggregation over subgroups when the distributions of wealth and income evolve with life-expectancy and social security provision. It is important to check the parameter stability of the wealth effects in all countries.

The log formulation of the consumption function is very convenient with exponentially trending macro data, since residuals are likely to be homoscedastic. Adding further realistic features, splitting up assets into different types and introducing a role for the credit channel, gives rise to a modern empirical version of the Friedman-Ando-Modigliani-Brumberg consumption function that encompasses the basic life-cycle model in (2.7):

$$\begin{aligned}\Delta \ln c_t \approx \lambda\,(&\alpha_{0t} + \ln y_t - \ln c_{t-1} + \alpha_{1t} r_t + \alpha_{2t}\theta_t + \alpha_{3t}(E_t \ln yperm_t - \ln y_t) \\ &+ \gamma_1\, NLA_{t-1}/y_t + \gamma_2\, IFA_{t-1}/y_t + \gamma_{3t}\, HA_{t-1}/y_t) \\ &+ \beta_{1t}\Delta \ln y_t + \beta_{2t}\Delta nr_t \left(DB_{t-1}/y_t\right) + \varepsilon_t\end{aligned} \tag{2.10}$$

Note that many of the parameters are time varying. The time variation induced by shifts in credit availability is discussed below. NLA_{t-1}/y_t is the ratio of liquid assets minus debt to non-property income, IFA_{t-1}/y_t is the ratio of illiquid financial assets to non-property income, and HA_{t-1}/y_t is the ratio of housing wealth to non-property income. The term $\Delta nr_t \left(DB_{t-1}/y_t\right)$, where nr_t is the nominal interest rate on debt DB_t, captures the cash flow impact on borrowers of changes in nominal rates. The speed of adjustment is λ and the γ parameters measure the marginal propensity to consume (mpc) for each of the three types of assets. The term in the log change of income can be rationalized by aggregating over credit constrained and unconstrained households.

The credit channel enters the consumption function through the different mpc's for net liquid assets (Otsuka, 2004) and for housing, through the cash flow effect for borrowers, and by allowing for parameter shifts arising from credit market liberalization. In principle, credit market liberalization should (i) raise the intercept α_0, implying a higher level of $\ln\left(c/y\right)$; (ii) lower the real interest rate coefficient, thereby raising α_1; (iii) raise α_3 by increasing the impact of expected income growth; and (iv) increase the mpc for housing collateral, γ_3. It should also lower the current income growth effect, β_1 and the cash flow impact of changes in the nominal rate, β_2. We handle these shifts in the UK model by writing each of these time-varying parameters as a linear function of the index of general credit supply conditions, *GCCI,* so that *GCCI* enters the model both as an intercept shift and interacted with several economic variables.

3. The Estimated UK Consumption Equation

We begin by estimating our version of the stylized rational expectations permanent income model in equation (2.10) with quarterly data. Consumption is real per capita consumer spending, including durables. Income is real per capita non-property income.[9] The net worth to income ratio is defined as liquid assets minus mortgage and other consumer debt plus net illiquid financial assets plus housing wealth, using end of previous quarter levels, relative to current income.

--- Table 1 About Here ---

In Table 1, Column 1 shows the text-book REPIH model with habits, equation (2.9) but omitting income uncertainty and the real interest rate, with highly significant estimates of net worth and income growth expectations effects and a speed of adjustment of 0.08 per quarter.[10] The long-run marginal propensity to consume out of net worth is 0.026. The real interest rate, the weighted change in nominal interest rates on debt and the change in the unemployment rate, a proxy for income insecurity, are added in Column 2. The real interest rate is insignificant, but the change in the unemployment rate and the nominal interest rate are both negative and significant. Column 3 relaxes the text book model, by allowing the ratio to income of net liquid assets (liquid assets minus consumer and mortgage debt) to have a different coefficient from illiquid assets and housing wealth. This radically affects the estimated wealth effects, with the estimated marginal propensity to consume (mpc) out of net liquid assets equalling 0.126, rather than the 0.026 figure in Column 1. The estimated mpc for illiquid financial assets is unchanged at 0.026, whilst the housing 'wealth' effect is estimated to be larger at 0.047. The speed of adjustment rises to 0.29 and the improvement in fit clearly rejects the text-book model in Column 1.

Finally, Column 4 allows the relevant parameters of equation (2.10) to shift with the UK general credit conditions index, *GCCI*. The expected shifts in parameters occur, though some are insignificant, and there is a large improvement in fit over Column 3. We show a

[9] This is defined as personal disposable income minus approximately tax-adjusted property income. Permanent income growth is forecast using a time trend, current and lagged four quarter changes in real income, log real income, average log real share prices and, from 1980 on, average log real house prices. Details are available on request.

[10] The specifications in Table 1 include an intercept, dummies for temporary consumption shifts due to VAT (sales tax) anticipations, measures of the change in consumer credit controls for durables purchases and of working days lost in labour disputes. A data appendix with sources, statistics and unit root tests for the UK, U.S. and Japanese variables is available upon request.

parsimonious version of the model. Consistent with the collateral view of housing wealth, the housing wealth-to-income ratio is insignificant, while its interaction effect with *GCCI* is strongly significant, and so we omit the former[11]. The marginal propensity to spend out of housing assets at the maximum value of *GCCI* is 0.043, while that of illiquid financial assets is 0.022, which, in turn, is far below that of net liquid assets at 0.114. These results for the housing assets effect are lower than in most of the literature. We find that a four-quarter moving average of observations on illiquid financial assets fits a little better than the end of previous quarter value, consistent with Lettau and Ludvigson (2004).[12] Since much of illiquid financial assets are in pension funds, this plausibly reflects the slow adaptation of contribution and pay-out rates to changes in asset values.

The real interest rate effect is negative and significant. According to point estimates, not shown, this effect strengthens as *GCCI* rises. The debt-weighted nominal interest rate change, also negative, weakens as *GCCI* rises. With easier access to credit, inter-temporal substitution should play a bigger role, explaining these two results and some evidence of an enhanced role for income growth expectations as *GCCI* rises. Income uncertainty is tracked by the four-quarter change in the unemployment rate, which negatively affects consumption. The interaction effect with *GCCI* is positive, but insignificant, suggesting that higher debt levels have offset the reduction in income uncertainty effects one might have expected from easier access to credit. The quarterly speed of adjustment is 0.37 meaning that over 80 percent of the adjustment of consumption to equilibrium occurs within four quarters.

The parameters of this equation are remarkably stable in charts of recursive estimates. The model can be interpreted in terms of co-integrated variables. Effectively, the log ratio of consumption to non-property income and the three asset-to-income ratios form a co-integrated relationship between four I(1) variables, subject to a shift in the intercept via *GCCI*. Since the real interest rate is arguably I(0) and plays only a marginal role, we can omit it. We carried out a co-integration analysis, in which we treat *GCCI* as an exogenous shift dummy, and include in the equation system I(0) variables such as income growth, forecast growth, and the change in the unemployment rate. We also included the impulse dummies, but outside the co-integration space. With a lag of two, there is only one co-integrating relationship and this is close to the long-run solution implied by the Column 4 estimates. Effectively, this analysis treats current income growth and the forecast of future growth and the unemployment rate as weakly exogenous variables. Evidence for weak exogeneity is found from models for these I(0) variables in which the lagged equilibrium correction term implied by the co-integration

[11] This interaction effect takes the form (housing wealth/income minus its 1980-2005 mean) multiplied by *CCI*.
[12] However, the estimated stock market effect over 1 to 2 years in Lettau and Ludvigson is implausibly small.

vector is insignificant.[13] For the UK, therefore, the pessimism expressed by Lettau and Ludvigson (2004) and Carroll *et al.* (2006) regarding the existence of a cointegrating relationship between consumption, income and assets appears to be misplaced, at least once the *GCCI* effect is included and assets are split into the three components indicated.

A further specification check estimates the model introducing a smooth stochastic trend to capture omitted demographic and other trending effects as discussed below for the U.S. Using STAMP software (Koopman *et al.*, 2006), we find no indication of such a trend. This suggests that, for the UK, the net influence of such omitted effects on consumption is small relative to the large variations in asset prices, credit conditions, unemployment changes and other shocks. The indications are that higher income inequality may have lowered the consumption-to-income ratio while a higher proportion of adults aged over 65 may have raised it. But these trending effects are hard to identify.

--- Figures 1 and 2 About Here ---

Figures 1 and 2 show the long-run contribution to the log consumption-to-income ratio of the three asset-to-income ratios and the credit conditions index, weighting each by its estimated long-run coefficient. As noted below, these are not general equilibrium effects. Figure 1 suggests that a substantial part of the upturn in consumption relative to income can be attributed to the rise in the credit conditions index and that rising collateral values of homes relative to income helps account for relative rises in consumption from 1984 to 1989 and from 1995 to 2005. The role of income growth expectations was far smaller, having a negative effect over 1984-89 and a small positive role over 1995-2005.

Figure 2 further suggests that the upward trend in the value of illiquid wealth holdings relative to income also played in important part in the upward trend in consumption relative to income. However, the rise in debt, reflected in the fall of net liquid assets relative to income has major offsetting effects in the long run. The fact that the estimated marginal propensity to consume out of net liquid assets is substantially higher than that out of other assets is critical. Much conventional discussion of wealth effects focuses on net worth and so misses the special role of liquidity and of debt. UK consumption levels are quite vulnerable to downturns in asset prices, given the level of debt and the fact that debt is hard to reduce in the short-run.

[13] While income is likely to be endogenous for consumption, on the UK data, current quarter growth of real income appears to be weakly exogenous for the log consumption to income *ratio.*

4. U.S. Results

In order to estimate a U.S. version of our consumption function (2.10), we need to measure shifts in the credit supply function facing households. Unlike the UK series of Fernandez-Corugedo and Muellbauer (2006), the closest U.S. data source for tracking mortgage loan-to-value (LTV) and loan-to-income (LTI) ratios is the American Housing Survey. However, the sample is far smaller than in the UK survey of mortgage lenders and the LTV data are only usable from 1979. Nevertheless, as neither ratio rose much from 1979 to 1998, it does suggest that the easing of mortgage credit conditions for US first-time home-buyers may have been less dramatic than for the UK (see Duca *et al.,* 2011a). However, the evidence is that large exogenous shifts in the supply of non-mortgage consumer credit occurred since the early 1970s in the U.S.

4.1 Consumer Credit Index for the U.S.

One data advantage the U.S. has over the UK is the Federal Reserve's long running quarterly Senior Loan Officer Opinion Survey. Using this survey Duca, Muellbauer, and Murphy (2011b) construct an index of U.S. consumer credit conditions (*CCI*). *CCI* is based on a quarterly diffusion index (*CR*) tracking the quarterly change in the willingness of 60 large banks to make consumer instalment loans. This index is negatively and significantly correlated with a diffusion index of the net percentage of banks that tightened credit standards on non-credit card consumer loans, which is available since 1993. Before constructing a levels index from change index, they adjust *CR* for the effects of changes in interest rates and in the macroeconomic outlook using a regression based on screening models. This adjusted index of the relative change in the availability of consumer instalment loans is aggregated into a levels index based on 1966-82 correlations of the index with the growth rate of real consumer loan extensions at banks. The resulting CCI rises greatly during the 1980s, and then rises during the height of the subprime mortgage boom 2004-06, before reversing the gains of the early decade since 2006 (see Figure 3 below).

4.2 Consumption Function Estimates for the U.S.

A sequence of models was estimated for the U.S., similar to those run for the UK. The results are shown in Table 2. Column 1 shows the simplest specification - a traditional life-cycle model with habits. The dependent variable is the change in the log of real per capita consumption. Income is real per capita non-property income (labor plus transfer income) adjusted for temporary taxes à la Blinder and Deaton (1985). The explanatory variables include the income error correction term, $\ln y_t - \ln c_{t-1}$, the change in log income reflecting the possibility that some households simply spend current income, our proxy for permanent income relative to current income, and the ratio of last quarter's net worth to income. Dummies were also added for the imposition of the Carter credit controls in 1980 q2, oil shocks from Mideast disturbances, a proposed but later aborted future tax hike in 1974 q4, a major coal strike in the first half of 1978, as well as for the temporary effect of tax shifting surrounding the major 1987 Tax Reform Act.[14]

To forecast permanent income we use a simple model based on reversion to a split trend, with a slow-down in growth from 1968 as well as a small pickup in 1988 which reverses in 1999, with just two economic drivers. These are the four-quarter change in the three-month Treasury bill yield, which captures the impact of monetary policy, and the Thomson Reuters / University of Michigan survey measure of consumer expectations.

--- Table 2 About Here ---

In the basic model in Column 1, the estimated long-run mpc for net worth is 0.04, near conventional estimates. The speed of adjustment is very low, however, at 0.11 per quarter and the role of current income growth is dominant. The residuals suffer from serious autocorrelation. Column 2 breaks up wealth into its three components with little change in overall fit or the key coefficients in Column 1. The estimated net liquid assets mpc jumps to over 12 percent. The estimated illiquid financial assets and housing wealth mpc's are approximately 5 and 7 percent respectively.

The change in the unemployment rate and the change in the nominal interest rate and the level of the real interest rate are added in Column 3. The interest rate is the auto-finance rate adjusted for depreciation, which reflects special offers sometimes available to borrowers. Both variables are highly significant, the same as in the UK. Surprisingly, the estimated mpc's from net liquid assets, illiquid financial assets and housing wealth are fairly similar. This is implausible since liquid assets such as cash should be more spendable than illiquid

[14] The oil shock dummy equals 1 in 1973 q4, 1974 q1, 1979 q2, and 1990 q4 and 0 otherwise. The 1974 q4 dummy captures the reaction in that quarter to President Ford's proposed tax hike, which he dropped in 1975 q1. The coal strike dummy equals 1 in 1978 q1, -1 in 1978 q2 and 0 otherwise. Tax changes induced capital gains realizations in 1986, causing large one-time capital income tax payments and a plunge in consumption in 1987 q1, followed by a bounce back the following quarter.

financial assets and housing assets which, even if used as collateral for a loan rather than sold outright, can entail large transactions costs.

A plausible explanation is that degree to which U.S. households can use housing wealth as collateral has varied over time. Ignoring changes in the liquidity of housing wealth imparts a downward bias in estimate of the net liquid assets mpc. Prima facie support for this argument is provided by the almost monotonic decline in the ratio of net liquid assets to income since the early 1980's. In part, this decline reflects easier mortgage credit conditions, which tend to boost consumption.[15] A simple check on the model in Column 3 is to add a smooth stochastic trend, estimated with the STAMP software (Koopman *et al*., 2006). The significance of the estimated stochastic trend is evidence for a missing trending factor.

Column 4 explores the role of financial liberalization by adding our consumer credit conditions index, *CCI*, and interacting housing wealth with an index of housing liquidity *HLI*. Duca *et al*. (2010b) estimate *HLI* as a common latent factor / spline in a three-equation model of U.S. consumption, mortgage equity withdrawal and mortgage refinancing. Similar estimates of HLI were obtained from a non-linear, two equation state space model for consumption and mortgage refinancing.

--- Figures 3 and 4 About Here (DESCRIBE) ---

It turns out that both the *CCI* and *HLI* times housing wealth terms are highly significant. Their inclusion in Column 4 lowers the standard error by nearly 20 percent and raises the speed of adjustment from 0.10 to 0.32, implying a better specified long-run solution.[16] The significance of housing wealth interacted with *HLI* and the insignificance of non-interacted housing wealth supports the collateral view of housing wealth in the consumption function. This accords with both theory and micro evidence that observed housing wealth effects result from greater housing collateral (Browning *et al.*, 2009). Finally, Column 5 omits the insignificant level of housing wealth term in Column 4.

--- Table 3 About Here ---

[15] The costs of refinancing fixed interest rate U.S. mortgages fell in the 1990s, as shown by Bennett, Peach and Peristiani (2001) and discussed by Green and Wachter (2007). Duca *et al.* (2010b) estimate the latent costs of refinancing U.S. mortgages since the 1970s. They find these costs shifted down substantially in the late 1990s, making it more feasible to refinance mortgages and increasing the collateral value of housing wealth.

[16] In other runs we instrumented current income using a simple forecasting and obtained very similar results, with the *CCI* and *HLI* times housing wealth terms highly significant. This suggests that the endogeneity of current income is not a major issue.

The much higher mpc of net liquid assets (15 percent) than from illquid financial assets (1 percent) is broadly consistent with microeconomic evidence in Gross and Souleles (2002). Likewise, the higher mpc of housing wealth than from illiquid financial wealth supports the findings of Benjamin *et al.* (2004), Case *et al.* (2005) and Carroll *et al.* (2006). In addition, the similarity in the estimated wealth mpc's for the UK and U.S is noteworthy (Table 3). The mpc for net liquid assets is estimated to be 11 percent in the UK and 15 percent in the U.S. The estimated mpc for illiquid assets are 2 precent in the UK and 1 percent in the U.S. The peak mpc for housing wealth is estimated to be between 4 and 5 percent in both countries, despite a number of structural differences in the two housing markets.[17] Other researchers obtain much higher housing wealth mpc's for the U.S.

An important paper by Slacalek (2009) presents estimates of housing 'wealth' or collateral effects on consumption for a range of countries, using a different methodology. His evidence suggests that institutional differences between countries have large effects on the mpc out of housing wealth – they are larger in countries with more liberal mortgage markets. This concurs with our evidence for the UK, U.S. and Japan. His evidence is also consistent with an upward drift over time in the housing 'wealth" mpc, linked with credit market liberalisation. However, Slacalek's estimated housing 'wealth' mpc's for the UK, the U.S. and other Anglo-Saxon economies are far larger than ours, probably because he does not include any measure of consumer credit conditions in his models.

5. Results for Aggregate Japanese Consumption

Since we are using annual data from 1961 to 2008, we construct permanent income using a five year horizon and a 20% per annum discount rate, equivalent to the 5% per quarter rate used for the UK and U.S. Our permanent income forecasting model includes a trend and split trends from 1973 and 1991 - reflecting a large slowdown in Japanese growth from 1973 and a smaller slowdown in 1991 - as well as the level of log real per capita income. Other variables include the growth in the working age population which has been negative in recent years; the log of U.S. GDP reflecting trade; the real (Topix) share price index; the two year moving average of government debt to GDP, consistent with a partially Ricardian view,

[17] On the one hand, transactions fees (estate agents fees and taxes etc.) are much larger in the U.S. than in the UK, implying greater housing liquidity in the UK. On the other hand, mortgage borrowers with negative equity in the U.K. who hand back the keys to their house are still liable for the mortgage loss. In some U.S. states (e.g., California) there are no deficiency judgements so lenders have recourse only to the house collateralizing the mortgage, which may make housing more attractive to borrow against.

as well as the two-year change in nominal interest which has a negative effect. The income forecasting equation displays remarkable parameter stability over alternative samples.

--- Table 4 About Here ---

We estimate variants of equation (2.10) for Japan using annual data from 1961 to 2008. In the basic REPIH model in Column 1, only current income growth is significant. *Inter alia*, the real interest rate, forecast future income growth, the mpc out of new wealth and the income error correction term, $\ln y_t - \ln c_{t-1}$, are all insignificant. The first two of these variables become significant when we add income uncertainty indicators, the unemployment rate and income volatility, in Column 2. The rate of acceleration of the ratio of the population under 20 years of age to the population aged 20 to 64 is also significant.

In Column 3, net wealth is disaggregated into net financial assets excluding stocks, stocks, and physical assets including housing, which is never significant. However, real land prices are significant at the 10% level and negatively signed. The real interest rate, speed of adjustment / income error correction and forecast future income growth / permanent income terms are all highly significant and positive. The change in the nominal interest rate is always insignificant, unlike the UK and U.S. The positive real interest rate effect in Japan is not a disguised inflation effect as the inflation rate is insignificant when added, while the real rate remains significant.

Finally, in Column 4, we drop the insignificant physical assets-to-income ratio and merged the two financial asset ratios into a single, significant net financial wealth-to-income ratio. The estimated mpc from net financial wealth is 6 percent. We also interact income volatility with forecast income growth rate consistent with theory which suggests that greater income uncertainty should increase the discount on expected growth. When income volatility and its interaction with forecast income growth are included in the equation, only the interaction term is significant so we omit the level of income volatility.

We find no evidence of a shift in the consumption function due to financial liberalization, in contrast to the results for the UK and U.S. This equation is stable when estimated over samples from 1959 to 1992, to 1999, to 2006 and to 2008. In addition, the coefficients on the long-run land price and net financial wealth terms, along with those on income growth and the speed of adjustment do not change as different combinations of uncertainty variables are added. Together, these results suggest that the lack of financial liberalization and a role for housing collateral results in a negative impact of house prices on Japanese consumption.

--- Figures 5 and 6 About Here ---

Lower income growth and the uncertainty indicators explain some of the dramatic decline in the consumption-to-income ratio in the 1970s. The long-run contributions of the four I(1) explanatory variables – the net financial wealth-to-income ratio, the log real land price, the real interest rate and the forecast growth rate of income – are plotted in Figures 5 and 6. These figures show that the rise of the consumption-to-income ratio since the late 1970s is largely driven by a rise in net financial assets that is only partially offset by a rise in real land prices. Interestingly, net financial assets relative to income shows little cyclical variation, as the pension fund component is not very sensitive to the stock market, though its decline in the early 1990s also contributed to the drop in the consumption ratio then.

6. Concluding Comments

Consistent with theory, our empirical findings for the UK, U.S. and Japan demonstrate the importance of credit constraints for consumer spending. We find that the evolution of credit availability differs over time within countries, as well as between them. There have been large changes in credit availability to UK and U.S. households in recent decades. These shifted the consumption function in both countries. Furthermore, financial liberalization enhanced the positive impact of housing wealth on consumption in the UK and U.S. and the role of expected income growth in the UK. In contrast, the Japanese consumption function has been stable, reflecting a lack of household credit liberalization since the 1970s. The latter and differences in the tax code likely account for the restraint on consumption in Japan from rising home prices.

Comparing Japan with the UK and the U.S., the consumption function differences suggest that monetary policy transmission via the household sector is far less powerful in Japan. Large household liquid asset holdings in Japan relative to debt imply that households, particularly older households, feel poorer when short term interest rates fall and so reduce spending. In the U.S. and UK, where debt exceeds liquid assets, higher spending by debtors more than offsets this effect. To the extent that lower interest rates raise house prices, this also has a (small) negative effect on aggregate household spending as Japanese renters are likely to save more in anticipation of higher future rents or of higher mortgage down-payments. In the U.S. and UK, in contrast, greater possibilities for housing equity withdrawal combined with higher house prices to boost spending. The conventional positive effects of lower short term interest rates on household spending via financial asset prices and income growth expectations apply in all three countries.

Our findings suggest that the impact of the large declines in wealth in 2007-9, particularly in housing equity, will have strong and persistent dampening effects on consumer spending in the UK and U.S. Volatile housing wealth also reflects the impact of changes in mortgage credit standards. During the recent recession, negative wealth effects were compounded by a substantial tightening of consumer credit standards in the U.S., a combination not seen since 1974-75, when consumption was unusually weak (Duca *et al.*, 2010a). In both episodes, mortgage availability declined sharply. In the UK, a tightening of credit standards sharply increased loan-to-value and loan-to-income ratios for first-time home-buyers, contributing to large declines in house and other asset prices from historic highs. More recently, cuts in interest rates to historic lows have provided an important counterweight.

Japanese consumer spending is less *directly* affected, if at all, by falling Japanese housing wealth and credit availability. Nevertheless, the global recession, particularly in the U.S., proved detrimental to Japanese household income due to declines in net exports. Moreover, the damage from loan losses at financial institutions has been large enough to induce credit tightening and lower asset prices outside of the U.S. (Greenlaw *et al.*, 2008).

Strong similarities between the consumption functions for the UK and U.S. and their contrast with Japan reflects the importance of institutional differences. This underlines the contribution of modernizing the Ando-Modigliani-Brumberg consumption function for imperfect information, which unlike the textbook Euler equation approach, incorporates credit frictions, uncertainty and income expectations. Household balance sheets, neglected in many macroeconomic models, are critical. Without carefully accounting for evolving credit and wealth relationships, the impact of credit and financial shocks on household spending and the monetary transmission mechanism cannot be properly modelled or understood.

References

Ando, A. and Modigliani, F. (1963), "The 'life cycle' hypothesis of savings: Aggregate implications and tests", *American Economic Review*, 53, 55-84.

Benjamin, J., Chilloy, P. and Jud, D. (2004), "Real estate versus financial wealth in consumption", *Journal of Real Estate Finance and Economics*, 29(3), 341-84.

Bennett, P., Peach, R. and Peristiani, S. (2001), "Structural change in the mortgage market and the propensity to refinance", *Journal of Money, Credit and Banking*, 33(4), 955-75.

Blinder, A. and Deaton, A. (1985), "The time series consumption function revisited", *Brookings Papers on Economic Activity*, 2, 465-511.

Bosworth, B. , Burtless, G. and Sabelhaus, J. (1991), "The decline in saving: Evidence from household surveys". *Brookings Papers on Economic Activity*, 1, 182-256.

Brayton, F., Levin, A., Tryon, R. and Williams, J. (1997), "The evolution of macro models at the Federal Reserve Board", *Carnegie Rochester Conference Series on Public Policy*, 47, 43-81.

Browning, M., Gortz, M. and Leth-Petersen, S. (2009), "House Prices and Consumption: A Micro Study", mimeo, Nuffield College, Oxford.

Buiter, W. (2008), "Housing wealth isn't wealth", Economics - The Open-Access, Open-Assessment E-Journal, Kiel Institute for the World Economy, 4(22), 1-29.

Cameron, G., Muellbauer, J. and Murphy, A. (2006), "Was there a British house price bubble? Evidence from a regional panel," Centre for Economic Policy Research (CEPR) discussion paper no. 5619.

Campbell, J. (1987), "Does saving anticipate declining labor income? An alternative test of the permanent income hypothesis", *Econometrica*, 55(6), 1249-73.

Campbell, J. and Mankiw, N. G. (1989), "Consumption, income and interest rates: reinterpreting the time series evidence", in Blanchard, O. and Fischer, S. (eds.), *NBER Macroeconomics Annual 1989*, MIT Press, Cambridge MA.

Campbell, J. and Mankiw, N. G. (1991), "The response of consumption to income: A cross-country investigation", *European Economic Review*, 35(4), 723-56.

Carroll, C. (1997), "Buffer stock saving and the life cycle/permanent income hypothesis", *Quarterly Journal of Economics*, 112(1), 1-56.

Carroll, C. (2001). "A theory of the consumption function with and without liquidity constraints", *Journal of Economic Perspectives*, 15(3), 23-45.

Carroll, C., Otsuka, M. and Slacalek, J. (2011), "How large is the housing wealth effect? A new approach", *Journal of Money, Credit, and Banking*, 43(1), 55-79.

Case, K., Quigley, J. and Shiller, R. (2005), "Comparing wealth effects: The stock market versus the housing market", *Advances in Macroeconomics*, www.bepress.com/bejm/advances/vol5/iss1/art1.

Deaton, A. (1991), "Savings and liquidity constraints", *Econometrica*, 59(5), 1221-48.

Deaton, A. (1992), *Understanding Consumption*, Clarendon Press, Oxford.

Deaton, A. (1999), "Saving and growth", in Schmidt-Hebbel, K. and Serven, L. (eds.), *The Economics of Saving and Growth*, Cambridge University Press, Cambridge.

Deaton, A. and Muellbauer, J. (1980), *Economics and Consumer Behaviour*, Cambridge University Press, Cambridge.

Duca, J. V., Muellbauer, J. and Murphy, A. (2010a), "Housing markets and the financial crisis of 2007-2009: Lessons for the future", *Journal of Financial Stability*, 6, 203-17.

Duca, J. V., Muellbauer, J. and Murphy, A. (2010b), "Credit market architecture and the booms and busts in U.S. consumption", mimeo, University of Oxford.

Duca, J. V., Muellbauer, J. and Murphy, A. (2011a), "Credit constraints and house prices: Making sense of the U.S. experience", *Economic Journal*, forthcoming.

Duca, J. V., Muellbauer, J. and Murphy, A. (2011b), "A measure of U.S. consumer credit conditions", mimeo, Federal Reserve Bank of Dallas.

Duca, J. V. and Wu, T. (2009), "Regulation and the Neo-Wicksellian approach to monetary policy", *Journal of Money, Credit and Banking*, 41(3), 799-807.

Dynan, K. , Elmendorf, D. and Sichel, D. (2006), "Can financial innovation help to explain the reduced volatility of economic activity?", *Journal of Monetary Economics*, 53, 123-150.

Eichner, M., Kohn, D. and Palumbo, M. (2010), "Financial statistics for the united states and the crisis: what did they get right, what did they miss and how should they change?", Finance and Economics discussion paper no. 2010-20, Board of Governors of the Federal Reserve System.

Engelhardt, G. V. (1996), "Consumption, down payments and liquidity constraints", *Journal of Money Credit and Banking*, 28(2), 255-71.

Fernandez-Corugedo, E. and Muellbauer, J. (2006), "Consumer credit conditions in the UK", Bank of England working paper 314.

Friedman, M. (1957), *A Theory of the Consumption Function,* Princeton University Press, Princeton.

Friedman, M. (1963), "Windfalls, the 'Horizon' and Related Concepts in the Permanent Income Hypothesis," in Christ, C., Friedman, M., Goodman, L., Griliches, Z., Harberger, A., Liviatan, N., Mincer, J., Mundlak, Y., Nerlove, M., Patinkin, D., Telser, G. and Theil, H., (eds.) *Measurement in Economics: Studies in Mathematical Economics and Econometrics in Memory of Yehuda Grunfeld,* Stanford University Press.

Gokhale, Jagdish, Laurence Kotlikoff, and Jane Sabelhaus (1996), "Understanding the Postwar Decline in U.S. Savings: A Cohort Analysis," *Brookings Papers on Economic Activity*, 315-44.

Goldberger, A. S. (1974), "Unobservable variables in econometrics", in Zarembka, P. (ed.), *Frontiers of Econometrics,* Academic Press: New York, 193-213.

González-Páramo, J. M. (2009), "National accounts for monetary policy making: Reflections on the use of the Euro area accounts in the light of the financial crisis", speech to Eurostat 2009 National Accounts Conference, 16 September.

Goodhart, C. and Hofmann, B. (2008), "House prices, money, credit and the macroeconomy", *Oxford Review of Economic Policy*, 24(1), 120-44.

Green, R. and Wachter, S. (2007), "The housing finance revolution", in *Housing, Housing Finance, and Monetary Policy, A Symposium Sponsored by the Federal Reserve Bank of Kansas City,* Jackson Hole, Wyoming, 21-67.

Greenlaw, D., Hatzius, J., Kashyap, A. and Shin, H. (2008), Leveraged Losses: Lessons From The Mortgage Market Meltdown, U.S. Monetary Policy Forum Report No. 2, Rosenberg Institute, Brandeis International Business School and Initiative on Global Markets, University of Chicago Graduate School of Business..

Greenspan, A. and Kennedy, J. (2008), "Sources and Uses of Equity Extracted from Homes," *Oxford Review of Economic Policy*, 24(1), 120-44.

Gross, D. and Souleles, N. (2002), "Do liquidity constraints and interest rates matter for consumer behaviour? Evidence from credit card data", *Quarterly Journal of Economics,* 117, 149-85.

Hall, R. E. (1978), "Stochastic implications of the life cycle-permanent income hypothesis: theory and evidence", *Journal of Political Economy*, 86 (6): 971-87.

Hayashi, F. (1985), "The effect of liquidity constraints on consumption: A cross-sectional analysis", *Quarterly Journal of Economics*, 100(1), 183-206.

Hayashi, F. (1997), "A review of recent literature on Japanese saving", in Hayashi, F. *Understanding Saving: Evidence from the United States and Japan,* MIT Press, Cambridge, MA.

Jackman, R. and Sutton, J. (1982), "Imperfect capital markets and the monetarist black box: Liquidity constraints, inflation and the asymmetric effects of interest rate policy", *Economic Journal,* 92, 108-28.

Jappelli, T. and Pagano, M. (1994), "Saving, growth and liquidity constraints", *Quarterly Journal of Economics,* 109, 83-109.

Joreskog, K. and Goldberger, A. (1975), "Estimation of a model with a multiple indicators and multiple causes of a single latent variable", *Journal of American Statistical Association,* 70, 631-9.

Kimball, M.(1990), "Precautionary saving in the small and in the large", *Econometrica,* 58, 53-73.

King, M. (1990), "Discussion of 'Is the UK balance of payments sustainable?'", *Economic Policy,* 5 (11), 383-387.

Kohn, D. (2008), "Productivity and innovation in financial services", speech at celebration of the 10th anniversary of the Banque Centrale du Luxembourg, Luxembourg, Nov. 12.

Koopman, S. J., Harvey, A., Doornik, J. and Shephard, N. (2006), *STAMP: Structural Time Series Analyser, Modeller and Predictor,* Timberlake Consultants Press, London.

Leamer E. (2007), “Housing is the business cycle”, in *Housing, Housing Finance, and Monetary Policy, A Symposium Sponsored by the Federal Reserve Bank of Kansas City*, Jackson Hole, Wyoming, 149-233.

Lettau, M. and Ludvigson, S. (2004), “Understanding trend and cycle in asset values: reevaluating the wealth effect on consumption”, *American Economic Review*, 94(1), 276-99.

Modigliani, F. and Brumberg, R. (1954), “Utility analysis and the consumption function: An interpretation of cross-section data,” in Kurihara. K. (ed.). *Post-Keynesian Economics*, New Brunswick, NJ: Rutgers University Press.

Modigliani, Franco and Richard Brumberg (1980), “Utility analysis and aggregate consumption functions: An attempt at integration”, in Abel, A. (ed.), *Collected Papers of Franco Modigliani*, Vol. 2, MIT Press, Cambridge.

Muellbauer, J. (1988), “Habits, rationality and myopia in the life-cycle consumption function”, *Annales d'Economie et de Statistique*, 9, 47-70.

Muellbauer, J. and Murphy, A. (1989), “Why has personal saving collapsed?”, *Credit Suisse First Boston Research*, July, 26–77.

Muellbauer, J. and Murphy, A. (1990), “Is the UK balance of payments sustainable?”, *Economic Policy*, 5 (11), 347-95.

Muellbauer, J. (1996), “Income persistence and macro policy feedbacks in the US”, *Oxford Bulletin of Economics and Statistics*, 58, 703-33.

Muellbauer, J. (2007), “Housing, credit and consumer expenditure”, in *Housing, Housing Finance, and Monetary Policy, A Symposium Sponsored by the Federal Reserve Bank of Kansas City*, Jackson Hole, Wyoming, 267-334.

Muellbauer, J. and Lattimore, R. (1995), “The consumption function: a theoretical and empirical overview”, in Pesaran, H. and Wickens, M. (eds.), *Handbook of Applied Econometrics*, Blackwell, Oxford.

Muellbauer, J. and Murata, K. (2011), “Consumption, land prices and the monetary transmission mechanism in Japan”, in Hamada, K., Kashyap, H. and Weinstein, D. (eds.), *Japan's Bubble, Deflation and Long-term Stagnation*, 175-216, MIT Press, Cambridge MA.

Murata, K. (1999), *The Consumption Function in Japan*, DPhil thesis, University of Oxford.

Otsuka, M. (2004), *Essays on Household Portfolio and Current Account Dynamics*, PhD Thesis, John Hopkins University.

Pagano, M. (1990), “Discussion of ‘Is the UK balance of payments sustainable?’”, *Economic Policy*, 5 (11), 387-90.

Poterba, J. M. and Manchester, J. (1989), “second mortgages and household saving”, *Regional Science and Urban Economics*, 19(2), 325-46.

Poterba, J. M. (2000), “stock market wealth and consumption”, *Journal of Economic Perspectives*, 14(2), 99-118.

Slacalek, J. (2009), “What drives personal consumption? The role of housing and financial wealth”, *B.E. Journal of Macroeconomics*, Berkeley Electronic Press, 9(1), 1-35.

Skinner, J. (1988), “Risky income, life-cycle consumption and precautionary saving”, *Journal of Monetary Economics*, 22, 237-55.

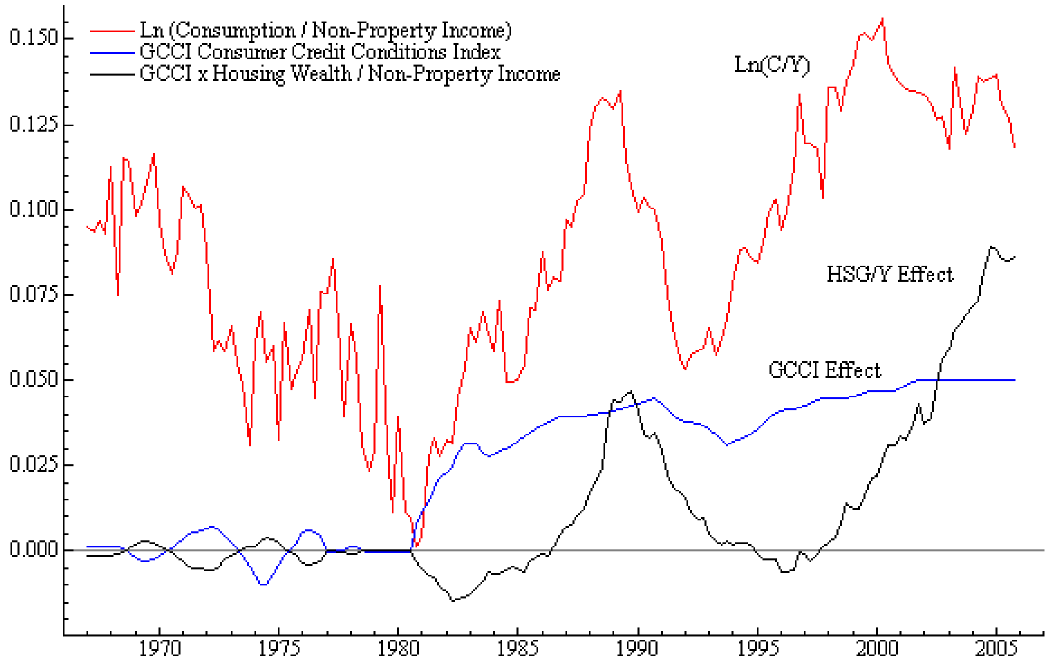
Staiger, D., Stock, J. and Watson. M. (1997), “How precise are estimates of the natural rate of unemployment?”, in Romer, C. and Romer, D. (eds.), Reducing Inflation: Motivation and Strategy, University of Chicago Press.

Stiglitz, J., and Weiss, A. (1981), “Credit rationing in markets with imperfect information”, *American Economic Review*, 71, 393-410.

Taylor, J. (2007), “Housing and monetary policy”, in *Housing, Housing Finance, and Monetary Policy, A Symposium Sponsored by the Federal Reserve Bank of Kansas City*, Jackson Hole, Wyoming, 463-76.

Zeldes, S. (1989), “Optimal consumption with stochastic income: Deviations from certainty equivalence”, *The Quarterly Journal of Economics*, 104(2), 275-98.

Figure 1: Estimated Long-Run Contributions to Log Consumption to Income Ratio of the General Consumer Credit Conditions Index (GCCI) and its Interaction with the Housing Wealth to Income Ratio in the UK.



Note: the UK asset to income ratios are defined as the ratio of the end of last quarter assets to four times current quarterly non-property income. The explanatory variables are multiplied by their estimated coefficients in the long-run solution. See the footnote to Table 1 for the definition of the interaction between GCCI and housing wealth to income ratio.

Figure 2: Estimated Long-Run Contributions to Log Consumption to Income Ratio of Net Liquid Assets (NLA) and Illiquid Financial Assets (IFA) to Income Ratios in the UK

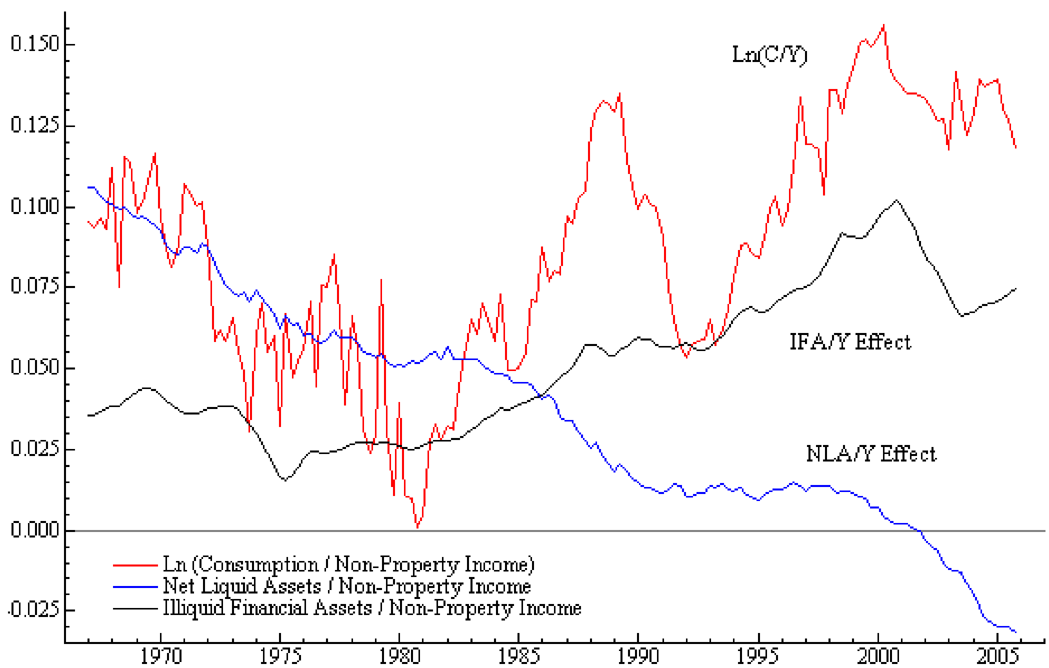


Figure 3: Estimated Long-Run Contributions to Log Consumption to Income Ratio of the Unsecured Credit Conditions Index (CCI) and the Housing Wealth to Income Ratio Adjusted for Housing Liquidity in the U.S.

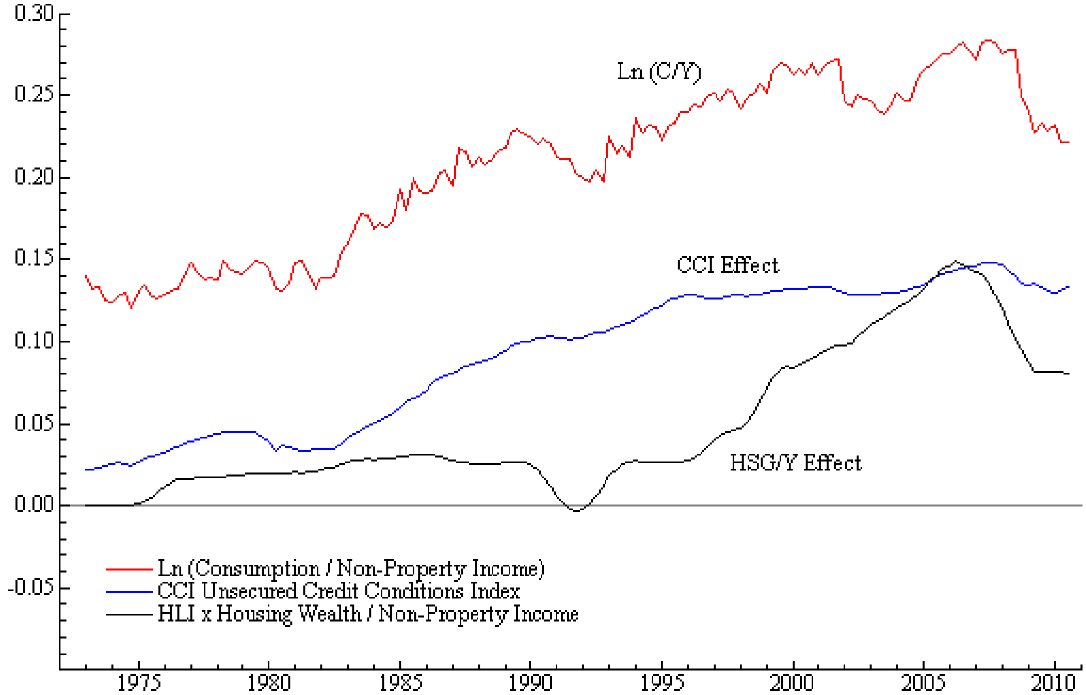


Figure 4: Estimated Long-Run Contributions to Log Consumption to Income Ratio of the Net Liquid Assets (NLA) and Illiquid Financial Assets (IFA) to Income Ratios in the U.S.

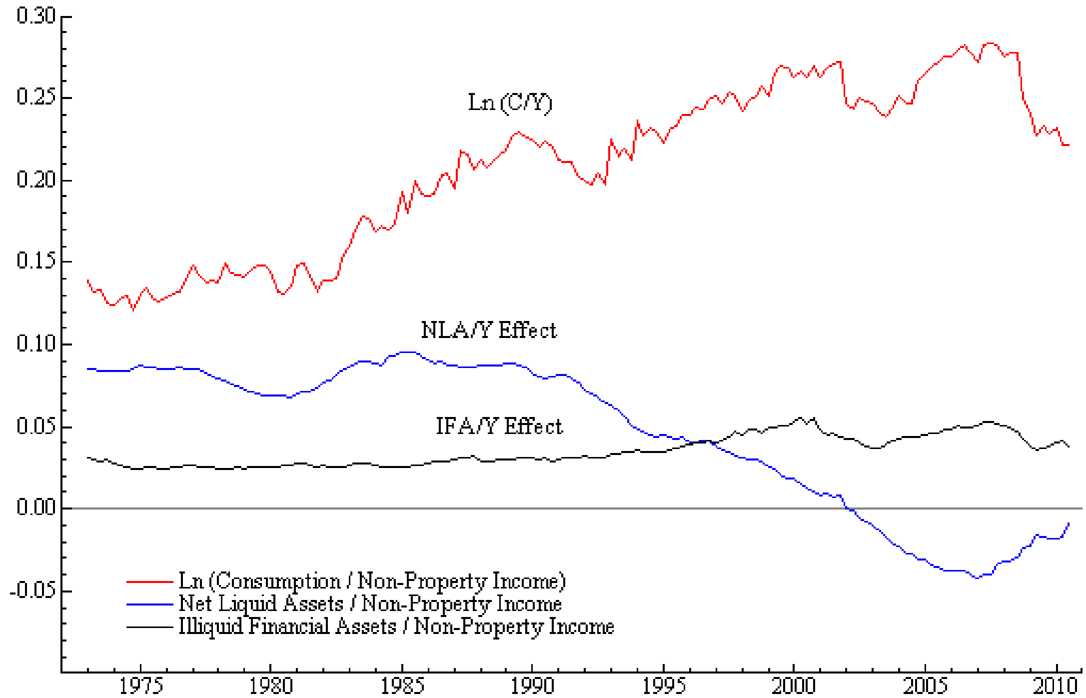


Figure 5: Estimated Long-Run Contribution to Log Consumption to Income Ratio of the Net Financial Wealth (NFW) to Income Ratio and Log Real Land Prices in Japan.

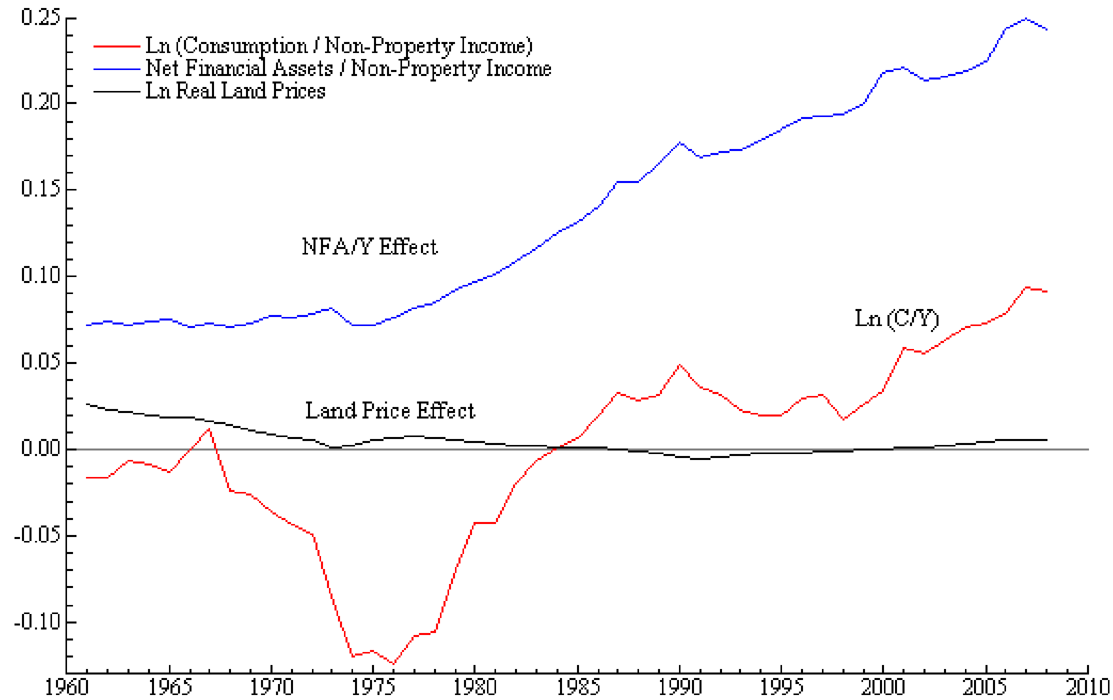


Figure 6: Estimated Long-Run Contribution to Log Consumption to Income Ratio of Real Interest Rate and Forecast Income Growth in Japan.

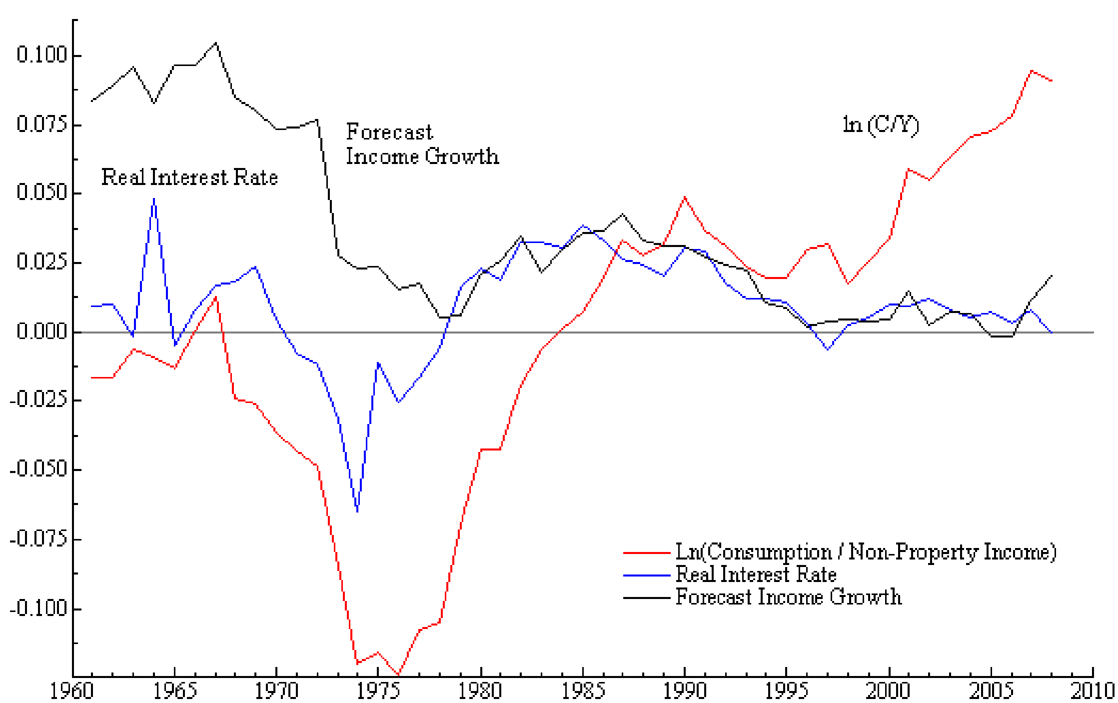


Table 1: UK Consumption Function Estimates for 1967 q1 to 2005 q4

Dependent Variable = $\Delta \ln c_t$	(1)	(2)	(3)	(4)
Constant (α_0)	-0.080** (2.1)	0.016 (0.4)	-0.166*** (7.0)	- 0.041*** (3.0)
$\ln y_t - \ln c_{t-1}$ (λ)	0.078** (2.4)	0.122*** (3.6)	0.291*** (5.9)	0.371*** (7.1)
Real mortgage rate (4 quarter average) x 100 (α_1)	-	0.070 (0.4)	-0.170** (2.4)	-0.155** (2.3)
Forecast future income aka permanent income growth (α_2)	1.061*** (2.7)	0.597*** (3.5)	0.485*** (6.5)	0.201*** (2.9)
Forecast income growth demeaned × *GCCI* interaction (α_3)	-	-	-	0.252 (1.6)
GCCI credit conditions index (α_4)	-	-	-	0.050*** (3.6)
Net worth$_{t-1}$/income ($\gamma_1 = \gamma_2 = \gamma_3$)	0.026*** (4.5)	0.011* (1.8)	-	-
Net liquid assets$_{t-1}$ / income (γ_1)	-	-	0.126*** (7.9)	0.114*** (7.8)
Illiquid financial assets$_{t-1}$ (4 quarter average) / income (γ_2)	-	-	0.026*** (7.2)	0.022*** (7.9)
Housing wealth$_{t-1}$ / income (γ_3)	-	-	0.047*** (8.6)	-
Housing wealth$_{t-1}$ / income (demeaned) × *GCCI* interaction (γ_4)	-	-	-	0.043*** (10.3)
Income growth (β_1)	0.250*** (5.5)	0.175*** (4.0)	0.093** (2.1)	-0.003 (0.1)
Change in unemployment rate (β_2)	-	-0.0043*** (5.6)	-0.0058*** (7.2)	-0.0071*** (8.43)
Change in debt weighted nominal interest rate (β_3)	-	-0.0020*** (3.3)	-0.0029*** (4.8)	-0.0061*** (3.8)
Change in debt weighted nominal interest rate × *GCCI* interaction	-	-	-	0.0041*** (2.1)
Change in credit controls (β_5)	-0.0007*** (2.5)	-0.0010*** (3.7)	-0.0012*** (4.7)	-0.0009*** (3.6)
Lagged change in working days lost to strikes / employment (β_6)	-0.0002*** (3.1)	-0.0002*** (3.7)	-0.0002*** (3.5)	-0.0002*** (3.4)
1968 q1 +1/-1 dummy for preannounced tax increase (β_7)	0.033*** (6.3)	0.033*** (7.0)	0.031*** (7.1)	0.030*** (7.1)
1973 q1 +1/-1 dummy for VAT introduction (β_8)	0.012*** (2.3)	0.012*** (2.4)	0.011*** (2.5)	0.011** (2.4)
1979 q2 dummy for VAT rise (β_9)	0.045*** (8.6)	0.044*** (9.3)	0.040*** (8.9)	0.039*** (9.1)
Q4 seasonal dummy(β_{10})	-0.022 (1.1)	-0.015 (1.3)	-0.008* (1.8)	-0.006** (2.1)

Table 1 (Continued)

	(1)	(2)	(3)	(4)
Standard Error x 100	0.72	0.66	0.62	0.58
Adjusted R^2	0.60	0.67	0.71	0.74
Durbin Watson	1.78	1.98	1.91	1.96
AR1/MA1 (p-value)	0.16	0.92	0.62	0.79
AR4/MA4 (p-value)	0.06	0.76	0.20	0.10
Heteroscedasticity (p-value)	0.52	0.38	0.87	0.05
Chow (1985 q1 break, p-value)	0.55	0.70	0.05	0.89
RESET (p-value)	0.50	0.82	0.29	0.13

Notes: t statistics are shown in parentheses. Statistical significance at the 10%, 5% and 1% levels is denoted by *, ** and *** respectively. The general model in column (4) is:

$$\begin{aligned}\Delta \ln c_t = \lambda\Big\{&\alpha_0 + (\ln y_t - \ln c_{t-1}) + \alpha_1 \bar{r}_t + \alpha_2 (\widehat{\ln y^p_{t+1} - \ln y_t}) + \alpha_3 CCI_t \times \text{demeaned } (\widehat{\ln y^p_{t+1} - \ln y_t}) + \alpha_4 CCI_t \\ &+ \gamma_1 NLA_{t-1}/Y_t + \gamma_2 \overline{IFA}_{t-1}/Y_t + \gamma_3 HSG_{t-1}/Y_t + \gamma_4 CCI_t \times \text{demeaned } HSG_{t-1}/Y_t\Big\} \\ &+ \beta_1 \Delta \ln y_t + \beta_2 \Delta_4 ur_t + \beta_3 \Delta nr_t \times DB_{t-1}/Y_t + \beta_4 CCI_t \times \Delta nr_t \times DB_{t-1}/Y_t + \beta_5 \Delta Credit\,Control_t \\ &+ \beta_6 \Delta WDL_{t-3} + \beta_7\, id1968q1_t + \beta_8\, id1973q1_t + \beta_9\, d1979q2_t + \beta_{10} q4_t + u_t\end{aligned}$$

where $\bar{r}_t$ and $\overline{IFA}_{t-1}$ are four quarter averages. Δnr_t is a weighted average of the changes in the nominal mortgage and unsecured borrowing rates, using the lagged shares of secured and unsecured borrowing as weights.

Table 2: US Consumption Function Estimates for 1973 q1 to 2010 q3

Dependent Variable = $\Delta \ln c_t$	(1)	(2)	(3)	(4)	(5)
Constant (α_0)	0.137 (1.3)	0.037 (0.3)	0.139 (1.6)	0.046^{*} (1.9)	0.047^{*} (1.9)
$\ln y_t - \ln c_{t-1}$ (λ)	0.093^{***} (4.3)	0.091^{***} (3.7)	0.097^{***} (4.3)	0.329^{***} (8.5)	0.326^{***} (8.7)
Real interest rate x 100 (α_1)	-0.007^{**} (2.1)	-0.008^{**} (2.4)	-0.008^{***} (2.8)	-0.004^{***} (3.7)	-0.002^{***} (3.7)
Forecast future income aka permanent income growth (α_2)	0.961^{***} (3.3)	0.948^{***} (2.9)	0.861^{**} (3.1)	0.599^{***} (6.3)	0.593^{***} (6.3)
Consumer credit index, *CCI* (α_3)	-	-	-	0.144^{***} (13.8)	0.146^{***} (16.7)
Net worth$_{t-1}$/income ($\gamma_1 = \gamma_2 = \gamma_3$)	0.039^{***} (7.8)	-	-	-	-
Net liquid assets $_{t-1}$ / income (γ_1)	-	0.125^{**} (3.0)	0.069^{**} (2.2)	0.146^{***} (11.2)	0.148^{***} (12.1)
Illiquid financial assets$_{t-1}$ / income (γ_2)	-	0.051^{***} (4.0)	0.042^{***} (4.1)	0.010^{***} (3.7)	0.010^{***} (3.7)
Housing wealth$_{t-1}$ / income (γ_3)	-	0.069^{***} (3.5)	0.051^{***} (2.9)	0.002 (0.4)	-
Housing wealth$_{t-1}$ / income × housing liquidity index *HLI* (γ_4)	-	-	-	0.081^{***} (9.4)	0.082^{***} (11.5)
Income growth (β_1)	0.208^{***} (3.5)	0.209^{***} (3.5)	0.182^{***} (3.5)	0.076^{*} (1.7)	0.073^{*} (1.7)
Change in unemployment rate (β_2)	-	-	-0.001^{***} (3.5)	-0.001^{***} (2.5)	-0.001^{**} (2.5)
Change in nominal interest rates (β_3)	-	-	-0.005^{***} (5.8)	-0.004^{***} (5.2)	-0.004^{***} (5.2)
Oil shock dummy – 1973 q4, 1974 q1, 1979 q2 & 1990 q4 (β_4)	-0.012^{***} (4.5)	-0.012^{***} (4.6)	-0.012^{***} (5.2)	-0.008^{***} (4.5)	-0.008^{***} (4.5)
1974 q4 dummy for proposed tax increases (β_5)	-0.023^{***} (4.5)	-0.022^{***} (4.4)	-0.019^{***} (4.2)	-0.018^{***} (5.1)	-0.018^{***} (5.1)
Coal strike +1/-1 dummy, 1978 q1 & q2 (β_6)	-0.007^{**} (2.0)	-0.007^{**} (2.1)	-0.007^{**} (2.3)	-0.006^{***} (2.6)	-0.006^{***} (2.7)
1980 q2 dummy for Carter credit controls (β_7)	-0.028^{***} (5.5)	-0.028^{***} (5.5)	-0.017^{***} (3.6)	0.017^{***} (4.5)	0.017^{***} (4.5)
Tax Reform Act +1/-1 dummy, 1987 q1 & q2 (β_8)	-0.008^{**} (2.3)	-0.008^{**} (2.3)	-0.007^{**} (2.2)	-0.007^{***} (2.9)	-0.007^{***} (2.9)

Table 2 (Continued)

	(1)	(2)	(3)	(4)	(5)
Standard Error × 100	0.50	0.48	0.43	0.34	0.34
Adjusted R^2	0.45	0.47	0.59	0.74	0.74
Durbin Watson	1.62	1.38	1.75	2.16	2.16
AR1/MA1 (p-value)	0.02	0.07	0.13	0.13	0.13
AR4/MA4 (p-value)	0.00	0.00	0.01	0.05	0.05
Heteroscedasticity (p-value)	0.20	0.24	0.63	0.55	0.55
Chow (1985 q1 break, p-value)	0.00	0.00	0.01	0.27	0.26
RESET (p-value)	0.12	0.10	0.58	0.91	0.99

Notes: t statistics in parentheses. Statistical significance at the 10%, 5% and 1% levels is denoted by *, ** and *** respectively. The general model in column 4 is:

$$\begin{aligned}\Delta \ln c_t = \lambda\Big\{&\alpha_0 + (\ln y_t - \ln c_{t-1}) + \alpha_1 r_t + \alpha_2 \big(\widehat{\ln y^{p}_{t+1}} - \ln y_t\big) + \alpha_3 CCI_t + \gamma_1\, NLA_{t-1}/Y_t + \gamma_2\, IFA_{t-1}/Y_t + \gamma_3\, HSG_{t-1}/Y_t \\ &+ \gamma_4\, HLI_t \times HSG_{t-1}/Y_t\Big\} + \beta_1 \Delta \ln y_t + \beta_2 \Delta_4 ur_t + \beta_3 \Delta nr_t + \beta_4 Oil\ Shock_t + \beta_5 d1974q4_t + \beta_6 Coal\ Strike_t \\ &+ \beta_7 d1980q2_t + \beta_8 Tax\ Reform\ Act_t + u_t\end{aligned}$$

Table 3: Estimated MPC's for Various Components of Net Worth in the UK and the U.S.

	Net Liquid Assets MPC	Illiquid Financial Assets MPC	Maximum Housing Wealth MPC
UK	0.114	0.022	0.043
U.S.	0.148	0.010	0.046

Notes: The estimated MPC's for the UK and US are from Table 1, Column 4 and Table 2, Column 5 respectively. The housing wealth MPC's were calculated using the maximum estimated values of *GCCI* (UK) and *HLI* (U.S.) respectively.

Table 4: Japanese Consumption Function Estimates for 1961 to 2008

Dependent Variable = $\Delta \ln c_t$	(1)	(2)	(3)	(4)
Constant (α_0)	0.318 (0.7)	0.188 (1.3)	-0.169*** (6.9)	-0.165*** (15.9)
$\ln y_t - \ln c_{t-1}$ (λ)	0.056 (1.4)	0.105*** (3.0)	0.489*** (6.3)	0.461*** (6.8)
Real interest rate (α_1)	0.059 (1.4)	0.027*** (2.3)	0.0073*** (4.8)	0.0083*** (6.3)
Forecast future income aka permanent income growth (α_2)	1.025 (1.2)	0.770** (2.3)	0.471*** (7.6)	0.460*** (7.7)
Forecast income growth × income growth volatility (α_3)	-	-	-	-3.688** (2.2)
Log real land prices (α_4)	-	-	-0.024* (1.7)	-0.021** (2.2)
Net worth$_{t-1}$/income ($\gamma_1 = \gamma_2 = \gamma_3$)	-0.033 (0.6)	-0.012 (0.7)	-	-
Net financial assets excluding shares $_{t-1}$ / income (γ_1)	-	-	0.064*** (11.9)	0.063*** (19.0)
Shares$_{t-1}$ / income (γ_2)	-	-	0.039 (0.7)	
Physical assets$_{t-1}$ / income (γ_3)	-	-	0.003 (0.6)	-
Income growth (β_1)	0.547*** (6.6)	0.448*** (6.1)	0.236*** (3.6)	0.272*** (4.4)
Change in unemployment rate (β_2)	-	-0.017*** (3.3)	-0.015*** (3.1)	-0.013*** (3.4)
Income volatility (β_3)	-	-0.310*** (2.8)	-0.169** (2.0)	-
Acceleration in ratio of population aged under 20 to working age population (β_4)	-	0.871** (2.2)	0.779** (2.5)	0.696** (2.4)
Standard Error × 100	0.99	0.81	0.60	0.57
Adjusted R^2	0.88	0.93	0.95	0.98
Durbin Watson	2.08	2.11	2.24	2.27
AR1/MA1 (p-value)	0.78	0.68	0.35	0.28
AR2/MA2 (p-value)	0.35	0.89	0.23	0.48
Heteroscedasticity (p-value)	0.26	0.70	0.16	0.08
Chow (1979 break, p-value)	0.07	0.06	0.90	0.86
RESET (p-value)	0.45	0.25	0.97	0.37

Notes: t statistics in parentheses. Statistical significance at the 10%, 5% and 1% levels is denoted by *, ** and *** respectively. The general model is:

$$\Delta \ln c_t = \lambda\Big\{\alpha_0 + (\ln y_t - \ln c_{t-1}) + \alpha_1 r_t + \alpha_2 (\widehat{\ln y^p_{t+1}} - \ln y_t) + \alpha_3 (\widehat{\ln y^p_{t+1}} - \ln y_t) \times \sigma_{\Delta \ln y,t} + \alpha_4 \ln pland_{t-1}$$
$$+ \gamma_1 (NFA_{t-1} - SHARES_{t-1})/Y_t + \gamma_2\, SHARES_{t-1}/Y_t + \gamma_3\, PHY_{t-1}/Y_t \Big\} + \beta_1 \Delta \ln y_t + \beta_2 \Delta ur_t + \beta_3 \sigma_{\Delta \ln y,t}$$
$$+ \beta_4 \Delta^2 (POP_{<20,t}/POP_{20-65,t}) + u_t$$