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

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Abstract: The contrasting consumption behaviour of UK, US and Japanese households is examined in this paper through the lens of modernised Ando-Modgliani style consumption functions. Income growth expectations are treated systematically and income uncertainty and credit channel features are incorporated. These models therefore capture important parts of the financial accelerator. The evidence is that, since 1980, credit availability for UK and US but not Japanese households has undergone large shifts. Moreover, there is UK and US evidence both for a shift in the average consumption to income ratio as down-payment constraints eased, and for a shift in the collateral role of housing wealth as home equity loans became more freely available. Point estimates suggest the housing collateral effect is larger in the US than the 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 US and positive ones in Japan, this suggests important differences in the transmission of monetary and credit shocks.

Keywords consumption, credit conditions, housing collateral and housing wealth

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1. Introduction

The global economic crisis of 2008-9 has its origins in a credit crisis. At the heart of a credit crisis is asymmetric information between lenders and borrowers, most simply the fear that lenders have about the ability and willingness of borrowers to service their debts. It is now obvious to almost all observers that the household credit channel played an important part in the boom which preceded the crisis, as well as in accentuating the crisis which began in the sub-prime mortgage market.

Figure 1 illustrates the multiple transmission channels of the original mortgage and housing crisis, though to avoid visual clutter, omits the reverse transmission from lower economic activity back to housing and mortgage markets, consumer spending, bank balance sheets, other asset prices and credit spreads. Four main channels are illustrated: the first is via residential construction. The second is via housing collateral and direct credit influences on consumption. The third and fourth channels both operate via the lowered capital base of banks and other financial firms feeding into credit standards and credit spreads, and more generally into the risk appetite of investors. This articulates the down-phase of the financial accelerator. However, the same mechanisms, if at a slower pace, operate in the up-phase, where an initial expansion of credit availability, due to financial innovation or low global interest rates, tends to be amplified.

--- Figure 1 About Here ---

The types of econometric models which, for the last decade, were popular with central banks and with main-stream macro economists, often neglected the financial accelerator. Many dynamic stochastic equilibrium models focused on building rationally-based macro models that could generate nominal rigidities by incorporating 'New Keynesian' frictions, primarily price stickiness and adjustment costs. The practical need to construct first-generation general equilibrium models led to the general adoption of micro assumptions that too often ignored the asymmetric information revolution of the 1970s and 1980s for which George Akerlof, Michael Spence and Joe Stiglitz shared the 2001 Nobel Prize in economics.

In a recent speech, the Vice Chairman of the Federal Reserve, Donald Kohn criticized these models[1]:

> "The recent experience indicates that we did not fully appreciate how financial innovation interacted with the channels of credit to affect real economic activity--both as credit and activity expanded and as they have contracted. In this regard, the macroeconomic models that have been used by central banks to inform their monetary policy decisions are clearly inadequate. These models incorporate few, if any, complex relationships among financial institutions or the financial-accelerator effects and other credit interactions that are now causing stresses in financial markets to spill over to the real economy.

Rather, these models abstract from institutional arrangements and focus on a few simple asset-arbitrage relationships, leaving them incapable of explaining recent developments in both credit volumes and risk premiums. Economists at central banks and in academia will need to devote much effort to overcoming these deficiencies in coming years."

At the present time, no ready-made model exists which fully captures the linkages and feedbacks shown in Figure 1. This necessarily implies that work is needed on the individual elements of such a model without, initially at least, a general equilibrium solution. Two important equations are for house prices and for consumer spending. Elsewhere, we have modelled the critical role of shifts in credit supply in explaining house prices and of extrapolative expectations in the overshooting of house prices[2]. In this paper, we present estimates of consumption functions for three major economies in the tradition of Modigliani and Brumberg (1954, 1980) and Ando and Modigliani (1963) but more explicitly incorporating both income expectations and credit channel influences, the latter of which can differ across these countries and over time.

Our work in this area began with Muellbauer and Murphy (1989, 1990) which explained the fall in the UK household saving ratio in partly in terms of UK credit market liberalisation and the increase both in house prices and in the 'spendability' of

[1] http://www.federalreserve.gov/newsevents/speech/kohn20081112a.htm
[2] Muellbauer and Murphy (1997), Cameron et al. (2006) and Duca et al. (2008).

housing and perhaps other illiquid wealth[3]. Our further research was summarized in Muellbauer and Lattimore (1995) which explained the foundations of a solved out consumption function encompassing both classical life-cycle/permanent income theory and credit channel features[4]. This and other related research have important implications for how the role of housing collateral in consumption decisions can differ across countries and structurally shift over time due to credit market liberalization. In a perfect capital markets world, higher home prices will not affect non-housing consumption because higher home prices perfectly capitalize expectations of higher future rents. However, with imperfect capital markets, the price and availability of borrowing are affected by agency costs that give rise to down payment constraints in housing markets. Research by Japelli and Pagano (1994) and Englehardt (1996) and others has shown that mortgage down-payment constraints generate an economically significant motive to save. In countries with low access to consumer and mortgage credit, such as Japan and Italy, higher home prices can induce higher saving for downpayments, thereby generating negative consumption effects. Credit market liberalization can generate two positive effects on consumption that can lead to strong housing wealth effects, such as in the UK and the U.S. First, credit liberalization lowers the typical down-payment required of first-time home buyers. Second, it provides households who face constraints in unsecured credit markets a greater ability to borrow against housing equity at lower interest rates. It therefore seemed likely that the aggregate household saving ratio, conditional on income, income expectations, interest rates and household wealth, would fall with credit market liberalization. Aron and Muellbauer (2000) built this effect into an extended consumption function and provided estimates for South Africa of a jointly estimated two equation model for consumption and household debt demonstrating the importance of credit liberalisation.

Measuring the shift in the credit supply function facing households was the subject of several unpublished papers in the 1990s, culminating in the most systematic estimates to date in Fernandez-Corugedo and Muellbauer (2006). In this paper, ten UK credit indicators were jointly modelled, controlling for standard economic and demographic variables, such as incomes, asset prices, interest rates, risk indicators

[3] The comments by King (1990) and Pagano (1990), instead took the view that an exogenous shift in income growth expectations accounted for most of the fall in the saving ratio.

[4] Although this paper had a notable influence on the consumption function of the Federal Reserve's FRB-US model and Brayton et al. (1997) devoted thorough attention to modeling expectations in FRB-US, the academic literature in macroeconomics has been dominated by approaches based on Euler equations with representative agents, as in many DSGE models.

and age composition of the population, to extract a latent variable,[5] the credit conditions index (CCI) interpreted as a scalar measure of the shift in credit supply facing households.

In Aron et al. (2008), summarized in Muellbauer (2007), we demonstrate that this credit supply shift had clearly significant effects on the consumption-to-income ratio, conditional on income, income expectations, changes in the unemployment rate, interest rates and household portfolios measured on quarterly data. Income growth expectations are modelled through an income forecasting equation. By interacting CCI with several other variables such as housing wealth, we provide evidence of other parameter shifts with credit market liberalisation, in line with theoretical priors. Moreover, we showed that given these CCI effects, the other parameters of the model are stable over 1967-2005, and co-integration tests are passed with flying colours.

In Japan, credit market liberalization for households since the mid-1970s appears to have been largely absent. In Muellbauer and Murata (2008), drawing on earlier work by Murata (1999), we apply the same general consumption model to annual data for Japan for 1961 to 2006. Again income growth expectations are controlled for using a separate income forecasting equation. We find no evidence of any parameter shifts, confirmed by estimates of a constant-parameter equation for household debt. Co-integration tests for this consumption function are highly satisfactory and instrumental variables estimates suggest a remarkable absence of endogeneity bias. Consistent with this absence of credit market liberalization, we find that for Japan the housing wealth or collateral effect is *negative*, in contrast to the UK and US. 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*, also in contrast with the UK and US.

Given this evidence, we turn to the U.S. consumption function, extending the very preliminary estimates in Muellbauer (2007). In properly modelling the role of income expectations, we compare estimates of a U.S. income forecasting equation with those for the UK and Japan. The long historical run of household survey data from the Michigan Survey allows our forecasting equation to be based more directly on household evidence than is possible for the other two countries. Nevertheless,

[5] The approach has parallels with the multiple indicator- multiples causes (MIMIC) approach to estimating latent variables of Goldberger (1974) and Goldberger and Joreskog (1975).

there are interesting common factors and some differences between the three countries.

For the consumption function, we find strong evidence for structural shifts in the consumption-to-income ratio, conditional on income, income growth expectations, interest rates, unemployment changes, and portfolio holdings. These structural shifts plausibly link with changes in credit market architecture, particularly since the early 1980s. Our estimates suggest that co-integration of aggregate US consumption, income and portfolio holdings for the last forty years will be hard to find without taking account of these credit market shifts.

We have emphasised the necessarily partial equilibrium or conditional nature of these consumption and income forecasting equations. Nevertheless, they have powerful applications to short and medium term policy formation, particularly in the economic crisis of 2008-9, even without full articulation of the rest of the feedbacks illustrated in Figure 1. For example, the UK estimates suggest that the housing collateral effect on consumer spending in recent years has been of the order of \$3 for every \$100 of variation in gross housing wealth, with about 80 percent of the effect experienced within four quarters. The combination of falling house prices, lower real incomes, less credit availability, rising unemployment, and lower stock market wealth made it possible to say with complete certainty by mid-2008 that the UK would be in recession in the second half of 2008. The earlier Bank of England view that there was 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 and for the slow policy response to the economic downturn of the MPC in September and October.

In contrast, while the FRB-US model does not take full account of shifts in credit conditions and may not have perfectly estimated the short-run response of consumption to housing collateral or wealth, it incorporates powerful housing and stock market 'wealth' effects. It seems likely that this aspect, coupled with a greater appreciation for the financial accelerator amongst U.S. policymakers, helps account for the early and decisive nature of the monetary policy response to the emerging crisis in the U.S.

Finally, our empirical evidence for Japan helps explain why the household component of the monetary transmission channel is far weaker in Japan than it is for the UK and the U.S. Had this been more clearly understood in 2001-2004, it seems likely that U.S. monetary policy would have been less concerned about the risk of a Japan-style 'lost decade' in the US. It is now widely agreed (e.g., Leamer (2007) and Taylor (2007)) that the federal funds rate was kept too low for too long in this period. Perhaps more importantly, there was an unusual and unsustainable liberalisation of the mortgage market that fuelled an unexpectedly strong credit, housing, and consumption boom, whose collapse is now playing out.

We conclude that although research of this type on aggregate time-series data has been deeply unfashionable in macroeconomics for the last two decades, it has crucial policy relevance. It also has a contribution to make in establishing some empirical considerations that may provide some guidance in the reconstruction of central bank models for which Donald Kohn, Charles Goodhart and others are calling.

2. Consumption Theory Background

2.1 Housing wealth effects.

We begin by demonstrating the weakness of the housing wealth effect in classical life-cycle theory. Let c = real non-housing consumption, p^h = relative price of housing, H = stock of housing, δ = rate of deterioration 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 c and on the stock of housing H.

Suppose expected relative house prices p^h and the real interest rate r are constant. Then the multi-period inter-temporal optimization problem is just a two-good problem with budget constraint:

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

where $(r+\delta)H$ = housing services and $p^{h}(r+\delta)$ = real user cost. 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) w.r.t. p^{h}, we find:

$$\partial[c + p^{h}_{0}(r+\delta)H] / \partial p^{h} = rH_{0} - (r+\delta)H \qquad (2.2)$$

But with $H \approx H_0$, the RHS of equation (2.2) is negative, since δ is positive. This point seems to have been overlooked in the classic work by Modigliani and Brumberg (1954), Friedman (1957, 1963) and Ando and Modigliani (1963). 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 place much store on a substantial aggregate housing wealth effect from classical life-cycle permanent income theory.

2.2 The Household Credit Channel

This section discusses how access to credit interacts with house prices, interest rates and income growth expectations to influence consumption and how a change in access to credit changes consumption through two main mechanisms. In many countries, mortgage debt is the dominant household liability. The first mechanism concerns the mortgage down-payment constraint. Suppliers of mortgage credit set upper limits to loan-to-income and loan-to-value 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 minimum deposit as a fraction of the value of the house[6]. A reduction in credit constraints in the form of a reduction in the minimum deposit as a fraction of the value of the house, will raise the consumption of these households relative to income (see Japelli and Pagano (1994) and Deaton (1999), and micro evidence in Engelhardt (1996)).

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-

[6] Note that most potential first-time home-buyers saving for a housing deposit are not credit-constrained in the sense of being unable to smooth consumption. The savings they are building up for a future housing deposit can be run down or increased in anticipation of shorter-term income fluctuations and in response to changes in real interest rates.

time buyers save more with higher house prices (unless they give up on house purchase). Increased access to credit will weaken the resulting negative effect on consumption of higher house prices.

Next, consider the second credit channel mechanism operating 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 the lower interest rates, given collateral backing. Effectively, the liberalization of credit conditions increases the "spendability" or liquidity of such previously illiquid housing wealth. The greater liquidity of housing wealth, with easier access to credit, gives housing wealth a buffer stock role.

Overall, combining the down-payment and collateral mechanisms with the life-cycle view relevant for some households, if existing owners have only limited access to home equity loans, the effect on their consumption of higher house prices will be small. Taking equation (2.2) literally implies that existing owners, who are not credit constrained and whose behaviour is governed by the life-cycle model outlined above, will have a small negative response to a permanent increase in real house prices unless they downsize to cheaper accommodation. By life-cycle theory, renters save more with higher house prices, as is implied by equation (2.2) when H_0 is zero. Hence, given the above discussion of the down-payment constraint, the aggregate consumption effect of a rise in real house prices is likely to be negative when access to credit is restricted, but switches 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 shocks to cash flows 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 represented by proportional changes in the nominal interest rate, weighted by the debt-to-income ratio. Better access to collateral will reduce the impact of such changes, as households with positive net equity can more easily refinance to protect cash flows against rises in nominal interest rates. The negative effect of nominal interest rate changes weighted by the debt-to-income ratio, 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-temporal substitution, enhancing the role of income growth expectations and, on balance, making the real interest rate effect more negative.

2.3 Aggregation and the Incorporation of Demographic Effects

In stylized solved out life-cycle consumption functions where we proxy expected or 'permanent' income by current income, micro-level consumption is given by a linear function of assets and non-property income:

$$c_i = \gamma_i A_i + \lambda_i y_i \tag{2.3}$$

where γ_i and λ_i vary by age. Hence aggregate or average per capita consumption is:

$$\overline{c} = \frac{\sum c_i}{N} = \left(\frac{\sum \gamma_i A_i}{\sum A_i}\right)\overline{A} + \left(\frac{\sum \lambda_i y_i}{\sum y_i}\right)\overline{y} \tag{2.4}$$

Thus $\overline{c} = \gamma^* \overline{A} + \lambda^* \overline{y}$ will have non-constant γ^* and λ^* depending 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 γ^* and A_i by age account for some of the secular decline in US saving rate. Similar arguments are common in Japan. However, cross-section evidence suggests that γ^* and λ^* may vary less across households than text book models might imply, Bosworth et al. (1991) and Murata (1999, ch. 8), for example, because of uncertainty about time of death. γ^* and λ^* are likely to evolve slowly over time as the age distribution, distribution of y and A by age and life expectancies evolve. Murata (1999, ch.5), using calibrations broadly consistent with micro data from the Japan Family Saving Survey, finds that aggregate consumption models in which γ^* and λ^* are constant have very similar implications and fit as models where they evolve according to sample survey data. Furthermore, as households make long-run portfolio decisions, the level and composition of assets is likely to reflect the demographic evolution, implying less direct impact on consumption of shifts in γ^* and λ^* due to demographic change.

2.4 A Solved Out Consumption Function

The Friedman-Ando-Modigliani consumption function requires an income forecasting model to generate permanent non-property income. Unlike the Euler equation, it does not ignore long-run information on income and assets. The solved out consumption function has advantages for policy modelling and forecasting. This basic aggregate life-cycle/permanent income consumption function has the form:

$$c_t = \gamma^* A_{t-1} + \lambda^* y_t^P \qquad (2.5)$$

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. This equation also has a basic robustness feature missing in the Euler equation. Euler equations require well-informed households continuously trading off efficiently between consuming now and consuming next period. Equation (2.5) is also 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 has a bearing on 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:

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

The right-hand side of equation (2.5) has the form 1+ x, where x is usually a fairly small number. We can then take logs, using the fact that $\ln(1+x) \approx x$ when x is mall and $\ln\left(y^P/y\right) \approx \left(y^P - y\right)/y$ .We then see that:

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

where $\gamma = \gamma^* / \lambda^*$ and $\alpha_0 = \lambda^*$. Thus, α_0 embodies the evolving distribution of income and demography, while γ embodies the evolving relative influences of the asset and income distribution and demography. One might attempt to proxy the former by the inclusion of demographic variables such as the population proportions in different age groups. The log ratio of permanent to current income reflects expectations of income growth and in practice can be proxied by functions of forecasted income growth rates.

The difference between log permanent and log current income in (2.7) can be expressed as

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

where $\Delta \log ym_{t+k}$ is defined as a weighted moving average of forward-looking income growth rates, see Campbell (1997). To dynamise the static form of equation for instance to introduce habits or adjustment costs, implies a partial adjustment form of equation (2.5) or (2.7), see Muellbauer (1988) for a derivation.

Further, extending the model to probabilistic income expectations, suggests the introduction of a measure of income uncertainty, θ_t and allows the discount factors in expected income growth, measured by $E_t \Delta \log ym_{t+k}$ to incorporate a risk premium, allowing the possibility that households may discount the future more heavily than by the real rate of interest (see Hayashi, 1985). If real interest rates are variable, standard theory suggests the real interest rate r_t enters the model, with the usual interpretation of inter-temporal substitution and income effects.

This gives the following generalisation of the canonical REPIH model in equation (2.7):

$$\Delta \ln c_t \approx \beta(\alpha_0 - \alpha_1 r_t - \alpha_2 \theta_t + \ln y_t + \alpha_3 E_t \Delta \ln ym_{t+k} + \gamma A_{t-1} / y_t - \ln c_{t-1}) + \varepsilon_t , \qquad (2.9)$$

where β measures the speed of adjustment. In principle, the coefficients α_3 and γ should depend upon the real interest rate, r_t; they should also depend on θ_t , since discount factors applied to expected incomes will increase with income uncertainty, as Skinner (1988), Zeldes (1989), and Carroll (1997, 2001b) have emphasized. For simplicity we will suppress this complication and the associated potential non-linearities.[7]

In practice, there are a number of reasons why income growth expectations embodied in $E_t \Delta \log ym_{t+k}$ are likely to reflect a limited horizon. With aggregate data it is difficult to forecast income beyond about three years. Indeed, widely used time series models have usually lost most of their forecasting power by then. 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, shorter 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 (2001b), who argues that plausible calibrations of micro-behaviour can give a practical income forecasting horizon of about three years - as Friedman (1957, 1963) suggested.

The log formulation is very convenient with exponentially trending macro data, since residuals are likely to be homoscedastic. Adding further realistic features, such as habits, a role for variable interest rates and income uncertainty, 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 consumption function that encompasses the basic life-cycle model given by (2.7).

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

The time variation in some of the parameters induced by shifts in credit availability is discussed below. $E_t \Delta \log ym_{t+k} = \log(y_t^P / y_t)$ measures income growth expectations. *NLA/y* is the ratio of liquid assets minus debt to non-property income, *IFA/y* is the

[7] In principle, the aggregate consumption function should also include effects arising from aggregation over subgroups when evolutions take place in distributions of wealth and incomes, see Section 2.3 above, in life-expectancy and in social security provision. We suspect that, over the 1967-2005 period, and given the magnitude of aggregate shocks, the UK is less sensitive to such omissions than many countries, but it is important to check the parameter stability of the wealth effects in all countries.

ratio of illiquid financial assets to non-property income, *HA/y* is the ratio of housing wealth to non-property income; $\Delta nr_t \left(DB_{t-1} / y_t \right)$,where *nr* is the nominal interest rate on debt DB, measures the cash flow impact on borrowers of changes in nominal rates; the speed of adjustment is β, and the γ parameters measure the MPCs 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, (see Muellbauer and Lattimore, 1995). Note that $\beta = 1,\ \alpha_{1t} = \alpha_{2t} = 0,\ \gamma_1 = \gamma_2 = \gamma_{3t},\ \beta_{1t} = \beta_{2t} = 0$ and $\alpha_{3t} = 1$ are the restrictions which result in the basic life-cycle/permanent income model equation (2.7).

The credit channel enters the consumption function through the different MPCs for net liquid assets (Otsuka, 2006) and for housing; through the cash flow effect for borrowers; and by allowing for possible parameter shifts stemming from credit market liberalization. Credit market liberalization should **(1)** raise the intercept α_0, implying a higher level of $\ln\left(c/y\right)$; (2) lower the real interest rate coefficient, thereby raising α_1; (3) raise α_3 by increasing the impact of expected income growth; and (4) 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 the change in the nominal rate, β_2. In our work on the UK, Aron et al. (2008), we handle these shifts by writing each of these time-varying parameters as a linear function of an index of credit supply conditions, *CCI* so that *CCI* enters the model as an intercept shift and in interaction with several economic variables.

3. UK Results

3.1 Income-Forecasting Equations

The dependent variable in the income forecasting equation, $\Delta \log yperm$, is defined as the difference between 'log permanent' and log current income given by (2.8), where the discount factor is 0.85 and the horizon k is 3 years, as originally suggested by Friedman (1963), see Carroll (2001) for discussion. With a discount value of 0.85, truncating the geometric formula for permanent income after 12 quarters introduces only a slight approximation error. To forecast $\Delta \log \mathit{yperm}$, we examined a range of alternative informational assumptions. At one extreme, we regress it simply on Δlog y and its lags, which would be the reduced form of an AR process in Δlog y. However, we allow for the possibility of longer lags by considering also $\Delta_4$log y at lags of 4 and 8 quarters. The only significant lag is a negative effect at lag 8, suggesting some kind of reversion in growth rates, but this is not a very stable relationship. The next simplest is to introduce a trend and the level of log y. This suggests strong trend reversion, with some persistence in the annual growth rate, and fits better. A further extension is to introduce changes in interest rates to reflect the influence of monetary policy on growth and levels of real asset prices. Given the widely discussed potential of asset prices to be proxies for income growth expectations, see King (1990), Pagano (1990), Attanasio and Weber (1994), Poterba (2000) and Attanasio et al. (2006), it is important that we control for this effect. Interestingly enough, the log real stock market index is not significant in this formulation and the log real house price index only begins to be relevant from 1981, after the advent of UK credit market liberalisation. The estimated equation with these elements is used to generate a 'naïve' forecast.

At the other extreme, we posit a long-run relationship for log y as a function of a linear trend (+), real interest rates (-), changes in nominal interest rates (-), the logs of real oil prices (-), share prices (+) and real house prices (+), the rate of tax on income (-), the rate of unionization (+) since greater union power should raise the share of labour income, and some national accounts ratios. These include the ratio of the government surplus to GDP where a higher ratio in the long run should allow lower tax rates or higher government spending, though offset in the short run by the negative 'Keynesian' effect of fiscal contraction, and the ratio of the trade deficit to GDP, since trade deficits have in the past constrained growth. However, there was a profound shift in fiscal policy around 1980, with the coming into power of the Thatcher government. This would be expected to have reinforced the positive role of the government surplus, and with the Burns-Lawson doctrine[8], to have led to trade deficits no longer mattering for fiscal policy. We find strong evidence for both

[8] The doctrine states that with free global capital flows, governments should concern themselves with budget deficits, but not with trade deficits and let these be a matter for the private sector. Terence Burns as chief economic advisor and Nigel Lawson as chancellor, made the doctrine official policy. Exchange controls were removed in 1979.

hypotheses by testing for interaction effects with pre and post 1980 dummies. We also test for a shift in the early 1980s in the role of real house prices, to be consistent with the shifting role of housing wealth in consumption with credit market liberalization. We confirm the absence of a positive real house price effect on income before the early 1980s, as in the 'naïve' model. We also checked for world growth and real exchange rate effects but failed to find stable relationships.

The long run level effects discussed all enter as 4-quarter moving averages, though for oil prices, the lags are even longer. Using a general-to-specific reduction procedure with HAC t-ratios and F-tests, we check for short run dynamics from changes in interest rates, where negative effects are confirmed, and growth rates of income and real oil and asset prices, in part to check for dynamic mi-specification due to the choice of 4-quarter moving average level effects.[9]

3.2 The Consumption Equation

We draw on Aron et al. (2008) to summarize key evidence for UK consumption. We begin by estimating our version of the text-book rational expectations permanent income model given by equation (2.10), with quarterly data. Consumption refers to real per capita consumer spending, including durables. Income is real per capita non-property income. The net worth to income ratio is defined as liquid assets minus debt plus illiquid financial assets plus housing wealth, taken as the 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 total wealth and income growth expectations effects and a speed of adjustment of 0.16 per quarter.[10] The long-run marginal propensity to consume out of net worth is obtained by dividing its coefficient 0.0036 by the speed of adjustment 0.16, to give 0.022. Column 2 shows one relaxation of the text book model, in which

[9] Details will be provided in the next version of this paper.

[10] All specifications reported in Table 1 also include an intercept, dummies for temporary shifts in consumption due to sales tax anticipations, a measure of the change in consumer credit controls for durables purchases, and a measure of working days lost in labor disputes.

the ratio to income of net liquid assets, defined as liquid assets minus debt, is permitted to have a different coefficient from illiquid assets. This radically affects the size of the wealth effects, with the marginal propensity to consume out of net liquid assets equalling 0.11 and that out of illiquid assets equalling 0.033, rather than the 0.022 implied by column 1. The speed of adjustment rises to 0.23 and the improvement in fit clearly rejects the text-book model in column 1. In column 3, we report on estimates of equation (2.10) again without including CCI or its interaction with any other variables. The additional variables are the change in the unemployment rate, a proxy for income insecurity, the real interest rate, the weighted change in nominal interest rates on debt, and a separate housing 'wealth' effect. Though the real interest rate is insignificant, the other effects are all significant and the marginal propensity to consume out of housing wealth effect is apparently larger at 0.036 than out of illiquid financial assets at 0.023. Clearly, the superior fit of this model rejects the restrictions embodied in columns 1 and 2.

Finally, we show a specification in column 4 in which we allow the relevant parameters of equation (2.10) to shift with the UK index of credit conditions, CCI.. The expected shifts in parameters all occur, though some are insignificant. Overall, the improvement in fit is significant relative to column 3. We show a parsimonious version of the model. The housing wealth-to-income ratio is insignificant, while its interaction effect[11] with CCI is strongly significant, and so we omit the former. The marginal propensity to spend out of housing assets at the maximum value of CCI (normalized at 1) is 0.032, while that of illiquid financial assets is around 0.019, which, in turn, is far below that of net liquid assets, at around 0.11. These results for the housing assets effect are lower than many found in the literature. We find that a four-quarter moving average of observations on illiquid financial assets fits better than the end of previous quarter value, consistent with findings by Lettau and Ludvigson (2004).[12] Since much of illiquid financial assets lies in pension funds, this plausibly reflects the slow adaptation of contribution and pay-out rates to changes in asset values.

[11] This interaction effect takes the form (housing wealth/income minus the mean value of this ratio from 1980 to 2005) multiplied by CCI. The post-1980 mean value of the housing wealth-to-income ratio is 3.08, compared to a 2005Q4 value of 4.71.

[12] However, over a one or two year horizon, the estimated stock market effect on consumption of Lettau and Ludvigson is implausibly small.

The real interest rate effect is negative, but significant only at the 10 percent level. According to point estimates, not shown, the evidence is that it strengthens as CCI rises. The debt-weighted nominal interest rate change, also negative, weakens as CCI rises. With easier access to credit, inter-temporal substitution should play a bigger role, explaining, as noted above, the enhanced role for income growth expectations, for which there is also evidence here. Income uncertainty is represented by the four quarter change in the unemployment rate, which has a negative effect on consumption. The interaction effect with CCI is positive, but quite insignificant, suggesting that higher debt levels may have offset the reduction in income uncertainty effects one might have expected from easier access to credit. The speed of adjustment is 0.33 meaning that 80 percent of the adjustment of consumption to income and the other explanatory variables is complete after four quarters.

The parameters of this equation are remarkably stable as the charts of recursive estimates shown in Aron et al. (2008) reveal. 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 CCI. Since the real interest rate is arguably I(0) and in any case plays only a marginal role, we can neglect it here. We carried out a co-integration analysis, in which we treat CCI as an exogenous shift dummy, and include in the equation system I(0) variables such as income growth and forecast growth and the change in the unemployment rate and 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 2 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 vector is insignificant[13]. For the UK, therefore, the pessimism expressed by Lettau and Ludvigson (2004) and Carroll et al. (2006) for the existence of a cointegrating relationship between consumption, income and assets appears to be misplaced, at least once the CCI effect is included and assets are split into the three components indicated.

[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*.

A further specification check on the model is to estimate it introducing a smooth stochastic trend, to capture omitted demographic and other trending effects, see discussion below for the US. Using the STAMP software (Koopman, Harvey, Doornik and Shephard, 2006), we find no indication of such a trend, in contrast to the US, see below. This suggests that the net influence of such omitted effects on consumption is small for the UK in this period, 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 2 and 3 About Here ---

Figures 2 and 3 show the long-run contribution to the log consumption-to-income ratio of the three asset to income ratios and of the credit conditions index, weighting each by its estimated long-run coefficient. As discussed further below, it should be noted that these are not general equilibrium effects. Figure 2 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 some of the upturn in consumption relative to income from 1984 to 1989 and much of the upturn from 1995 to 2005 can be attributed to the rises in the collateral values of homes relative to income.

Figure 3 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 seen in Figure 3, has major offsetting effects in the long run[14]. The fact that the estimated marginal propensity to consume out of net liquid assets is substantially higher than that out of other assets is quite important here. Much conventional discussion of wealth effects focuses on net worth and so misses the special role of liquidity and of debt. Figure 3 suggests that, as discussed in the

[14] As Ed Lazear and William White noted at the 2007 Jackson Hole Symposium, if the house price effect on consumption is mainly a collateral effect, payback time has to come. This is reflected in Figure 3.

Introduction, UK consumption levels are now quite vulnerable to downturns in asset prices, given that debt is hard to reduce in the short- run[15].

4. Results for Japan

4.1 The income growth forecasting equations.

For Japan, we work with annual data. In practice, it made little difference whether we use the one year ahead growth rate of income, or a weighted average of up to three years ahead, to model $\ln(y_t^P / y_t)$. Results are shown for the former. We follow a general to specific methodology in paring down a very general model to a parsimonious form. The general model includes a trend, a split trend from 1973 for the slowdown in Japanese growth which then occurred and the level of log real per capita income. Other variables include log US GDP, the log real exchange rate, log real oil prices, log real asset prices, the real interest rate, the change in the nominal interest rate, and the government surplus and debt to GDP ratios. Table 2 reports the parsimonious specifications found after testing down.[16]

--- Table 2 About Here ---

There is strong evidence for reversion to the split growth trend and the 3-year moving average of government balance/GDP has a positive coefficient. The table also shows an alternative specification in which lags in the ratio of government debt to GDP replace government balance/GDP. The lagged government debt/GDP ratios have highly significant coefficients, negative in the long run, as shown in the last column in Table 3. The US log GDP is also significant and real oil prices are then not significant, though they would be if US log GDP were omitted. The influence of the US as leading economy and key trading partner for Japan is thus confirmed. The change in nominal interest rates over the previous three years has a negative effect on next

[15] Note that at mid-2007 illiquid financial assets relative to income were substantially above the 2005 ratio shown in Figure 3, but have fallen sharply since.

[16] With the help of Autometrics software, Doornik (2007). Since the residuals reveal some heteroscedasticity, reflecting the greater volatility of pre-1975 growth, robust standard errors are reported

year's household income growth, suggesting that, overall, monetary policy has some negative effect on growth. However, the real interest rate is insignificant, and positively signed. Columns 2 and 3 show the income forecasting equation fitted over alternative samples, showing remarkable parameter stability.

4.2 Results for Aggregate Consumption

Our aim is to estimate for Japan variants of equation (2.10) discussed in section 2. We use annual data from 1961 to 2006. In a slight modification, we also include the lagged log real land price. It quickly becomes apparent that the ratio of physical assets to income and the real land price have negative coefficients and we therefore report equations in which each is included separately. Further, we cannot reject the hypothesis that the marginal propensities to spend out of deposits and illiquid financial assets are the same, and are equal to minus the coefficient on household debt. This may be because "deposits" includes a substantial amount of longer term time deposits which are therefore not so liquid. At any rate, we can work with a net financial wealth, which is always very significant and with a long run MPC of around 0.05 to 0.07 (see Table 3).

--- Table 3 About Here ---

The income uncertainty indicators enter with the correct signs. The measure of income volatility is significant as shown in the first column of Table 3. In the second column the cross term of income volatility and forecast income growth rate was added, as the theory discussed in Section 2 suggests that greater income uncertainty should lead to a bigger discount on expected growth. When both are included, the cross term was found to be significant while income volatility insignificant. The change in the unemployment rate is not significant at the 5% level, probably because of its more limited variability in Japan, in contrast to its far more significant role in the UK and the US. However, the sign is negative and the magnitude of the coefficient is not far below UK and US estimates.

The change in the nominal interest rate is always insignificant, unlike in the UK, but the level of the real rate has a strongly significant positive effect. This is not a disguised inflation effect as the inflation rate is insignificant when included, while the real rate remains significant. In Aron et al. (2008) and Muellbauer (2007) we find negative real interest rate effects in similar specifications estimated for the UK and the US.

Finally, the log change in income has a positive and significant effect. This is also in contrast to the UK and US findings, where this effect is not significant. The argument comes from applying the Campbell-Mankiw aggregation of credit constrained and unconstrained households to a solved out consumption function, see Muellbauer and Lattimore (1995). On this interpretation, the proportion of total income in income constrained households π is given by $(1-\pi)\beta = \beta_1$, where the coefficient on the change in log income is β_1 and β is the speed of adjustment. Given a speed of adjustment of 0.359 and β_1 estimated at 0.332 from Table 3 column 4, this suggests just over half of Japanese consumption comes from households who are, or behave as if they were, income constrained. This is not far from previous estimates of this proportion for Japan, see Hayashi (1997). However, given the somewhat unsatisfactory micro foundations for the Campbell-Mankiw story, it is probably a mistake to interpret this too literally in terms of credit constraints (see Carroll, 2001 and Aron et al., 2008).

In Muellbauer and Murata (2008), various charts of the fit of the equation over the full sample and recursive parameter estimates are presented. The stability properties of the equation are also evident in Table 3, when the equation is estimated over different samples. Together, these provide clear support for the long-term relevance of the model, though in short samples the real land price effect loses significance, given its lack of short-term variability.

This raises the question of whether there may have been a structural break in the coefficient on log real land price. We test for this by interacting the lagged log real land price with two step dummies, one zero up to 1980 and one from 1981; the other zero up to 1990 and one from 1991[17]. The results hardly alter when the step dummy beginning in 1991 was replaced by the one beginning in 1981. The coefficient on the step dummy interaction effect is not significant, with a t-ratio of 0.7. The point estimate is consistent with a small amelioration in the negative impact of

[17] To avoid a jump in the interaction effect in, for example, 1991, the 1991 step dummy is multiplied by the lagged log land price index minus its 1990 value.

land prices on consumption after 1991 (and indeed after 1981). But we can easily accept the hypothesis of constancy of the negative real land price effect.

--- Figures 4 and 5 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 shown in Figures 4 and 5. It is clear from this figure that the rise of the consumption to income ratio is very much driven by the rise in net financial assets owned by households, only somewhat offset by the rise in real land prices. Interestingly, net financial assets relative to income shows rather 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.

It is a striking fact that demographic variables, such as the share of the population aged 25 to 44 or 65 and over, are jointly and individually insignificant when included in the consumption equation.[18] This does not mean, of course, that demographic developments are irrelevant for aggregate consumption in Japan. The accumulation of financial wealth in Japan in the past has surely been, in part, driven by the ageing of the population and lengthening life-expectancy. Consumption or saving, *conditional* on such portfolio accumulations, is always less likely to be so sensitive to demographic structure.

5. US Results

5.1 Income Forecasting Equations

As for the UK, we estimate equations for DLYPERM (earlier called $\Delta \log yperm$), the deviation of log 'permanent' income from current log income,

[18] Moreover the signs often make little sense. What does make more sense is the inclusion of the proportion of the adult population aged 25 to 44. These are the main savers for a housing deposit. A rise in their proportion tends to lower consumption relative to income, though the effect still only has t=-1.1.Micro evidence, see Hayashi (1997), indicates that older Japanese households tend to carry on saving until their 80s, but perhaps the 25 to 44 group are the biggest savers.

LRYN, see equation (2.8). Income is real per capita non-property income as constructed for the FRB-US model.[19] The same quarterly discount factor of 0.85 is adopted as for the UK and two models, 'naïve' and 'sophisticated' are estimated.

The basic idea behind the naïve model is that households tend to project into the future income growth experienced in the past four years, measured by D16LRYNMA, the 16 quarter change in the 4-quarter moving average of LRYN, but take other factors into account. Chief among these is CEXP, the Michigan Survey based measure of how confident households are about financial conditions in the next year. Another factor is the change in short term interest rates in the previous four quarters, D4TB3, a fall (monetary easing) should improve growth prospects. A further factor is D4LRPPIM, the 4-quarter change in the log ratio of PPI for raw materials, including energy, to finished goods: a rise typically leads to cash flow problems for companies and increases in prices of goods, which with sticky wages, leads to short term real income contraction.

The model for DLYPERM was obtained from a more general model also including changes in the unemployment rate and changes in the log real S&P500 index – but these had 'wrong' signs or were insignificant.

In contrast, a more sophisticated model assumes that households believe in trend-reversion of income in the long run and also are aware of a wider range of factors. Trends are captured by a linear trend, a split trend beginning in 1968, capturing a growth slowdown, and the log of labour productivity, LLPRODMA, the last two measured as 4-quarter moving averages. Equilibrium correction is captured by LRYN, with a highly significant negative coefficient. Thus, when LRYN is above trend, future growth will tend to be lower, other things being equal. There is also evidence that higher recent growth *rates* tend to be followed by lower growth, as reflected in negative coefficients on 4-quarter growth rates at t, t-4 and t-8.

As in the naïve model, the Michigan Survey expectations measure is strongly significant, and enters as a 4-quarter moving average, CEXPMA. Annual changes in the unemployment rate, D4UNR, have negative effects on income growth, entering both at t and at t-4. As in the naïve model, monetary policy, as measured by the 4-quarter change in the three month T Bill yield, enters significantly at lags t and t-4.

[19] Non-property income is defined as tax-adjusted labor income plus transfer income. The particular variant used is adjusted, following Blinder and Deaton (1985), for temporary tax changes, and for the 1994Q1 blizzard.

The 4-quarter change in the log relative price of raw materials is also relevant at t and at t-4.

Stock market prices have no significant forecasting power, however, probably because survey expectations already capture their effect. In the absence of a reliable long run of house price data, cyclical fluctuations in the housing market not already in the survey expectations measure are captured by housing starts scaled by population. This enters as a 4-quarter moving average HOUSTCMA, lagged 4 quarters. It seems plausible that housing starts reflect optimism about the future and perhaps credit availability. Finally, as in the UK and Japan, there is some evidence of a 'Ricardian' effect in that large federal budget deficits tend to be followed by slower personal income growth. This is measured by a positive coefficient on the federal government surplus relative to GDP, entering as a 4-quarter moving average, RGFSURMA, lagged 4 quarters.

Figure 6 shows actual and fitted or forecast values of DLYPERM from naïve and sophisticated models, respectively.

5.2 Consumption Function Estimates for the US.

The first issue we address is the measurement of shifts in the credit supply function facing households. The closest US source for micro data on mortgage loan-to-value (LTV, or its equivalent, downpayment to value) and on loan-to-income ratios (LTI), used by Fernandez-Corugedo and Muellbauer in the UK context, is the American Housing Survey. However, the sample is far smaller than the UK survey of mortgage lenders and LTV data are usable only from 1979, too short a period to analyse consumption from 1966. Nevertheless, as neither LTV nor LTI ratios rose much from 1979 to 1998 it does suggest that, for first time buyers in the US before the 2000s, the easing of credit conditions may have been less dramatic than for the UK, see Duca, et al. (2008).

One data advantage the US has over the UK is the Federal Reserve's quarterly Senior Loan Officer Opinion Survey. For the U.S., a credit conditions index (CCI) is constructed from a quarterly diffusion index (*CR*) tracking the net relative change in bank willingness to make consumer instalment loans over the prior three months across 60 large banks in this survey. This index is negatively and significantly correlated with a diffusion index from the same survey of the net percentage of banks that tightened credit standards on non-credit card consumer loans, available since 1993. Before constructing a levels index from this relative change index, we first adjust it for identifiable effects of interest rates and the macroeconomic outlook by estimating an empirical model based on screening models (see Appendix). 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 7). While the index is tailored to consumer instalment credit rather than mortgage markets, it can serve as a first approximation to a more general credit conditions index.

We have estimated a sequence of models for the US of a similar form to the UK consumption functions shown in Table 2. Table 4, column 1 reports the simplest specification based on a simple extension of traditional life-cycle models with habits. The dependent variable is the log change in real per capita consumption. The regressors include log y-log c_{-1} , the log change in real per capita non-property income (to reflect the possibility that some households simply spend income), our naïve and sophisticated proxies for permanent income relative to current income, and the ratio of end of previous quarter net worth to income. The estimated long-run MPC for net worth is 0.05, a plausible number, but the speed of adjustment is very low at 0.05 per quarter, while the role of current income growth is dominant. The residuals suffer from serious autocorrelation. Column 2 adds the change in the unemployment rate, the change in the nominal interest rate and the level of a real interest rate, as in the UK. The interest rate is the auto-finance rate which reflects special offers sometimes available to borrowers. The first two of these variables are highly significant, as in the UK and reduce some of the specification problems seen in column 1. Much of the interest rate effect may reflect that special auto sales finance programs induce large, short-term inter-temporal substitution.

In column 3, the credit conditions index is added and the three main assets are disaggregated into net liquid assets, housing wealth and illiquid financial assets, as for the UK. The credit conditions index is highly significant and the speed of adjustment increases dramatically to 0.27, suggesting a far better specified long-run solution. The role of current income growth is now far reduced, while the weight on the sophisticated permanent income growth proxy is now positive (in contrast to columns 1 and 2). As far as the wealth effects are concerned, the data suggest that housing wealth has a far higher MPC than either net liquid assets or illiquid assets. On the surface, this is quite implausible since cash should be more spendable than housing assets, which, even if used as loan collateral rather than sold outright, are subject to the transactions costs of arranging loans.

One possibility is that short-run effects of higher housing wealth may be being confused with long-run effects. Another is that credit market liberalisation may have altered the spending power of housing wealth or collateral. Finally, and perhaps most convincingly, if the credit conditions index inadequately proxies shifts in credit availability to households, a downward bias in the coefficient on net liquid assets is likely to result. Since the early 1980s, RNLA1, the ratio of net liquid assets to income has fallen almost monotonically, in part, no doubt, under the influence of easier credit conditions. Since the latter has a positive effect on consumption, a downward bias on the coefficient on RNLA1 is likely.

Column 4 explores the role of housing wealth dynamics and interaction effects, beginning from more general dynamic specifications. It suggests that the acceleration of the housing wealth to income ratio is relevant as well as the level. The interaction of the 4-quarter change in housing wealth with the credit conditions proxy appears with a t-ratio of 1.6. However, the interaction of the housing wealth/income ratio with the credit conditions proxy is quite insignificant, t=0.3. All of these complications only marginally reduce the long-run MPC of housing wealth or collateral, however. In all specifications the interaction of the income growth expectations term with the US CCI is insignificant.

To provide prima facie evidence on the hypothesis that the credit conditions proxy may be inadequate, the same model was estimated using a stochastic trend in place of US CCI, in the STAMP software, see Koopman et al. (2006). The results are shown in column 5. The speed of adjustment rises sharply to 0.507, while the long-run MPCs are now far more plausible. They are 0.072 for net liquid assets, and 0.018 for illiquid financial assets (in which illiquid pension wealth has a large weight). For housing wealth or collateral, only the interaction effect with credit conditions now matters giving an MPC of 0.055 for housing collateral at the very peak of credit availability, but normally rather less. The dynamic effects are still present with t-ratios of 1.6 and 1.7 respectively[20]. Current income growth is now insignificant, while the model puts a larger weight on the sophisticated permanent income proxy.

Figure 8 shows the fitted stochastic trend. It differs in several respects from the credit conditions proxy, for example, showing a more pronounced downturn in the early 1990s and a more extended upturn from around 2002. Of course, this version of the model is not fully coherent, since the interaction effects use the CCI based on the Senior Loan Officer Survey, rather than the stochastic trend.

A standard criticism of time series estimates of solved out consumption functions is that there are many slowly moving, correlated factors that could be affecting consumption. These include demographic trends, evolutionary changes in the inequality of income and wealth and changes in social security and pensions systems, cohort-specific evolutionary shifts in attitudes in time preferences and risk, as well as long-term shifts in credit conditions. We cannot exclude the possibility that the estimated stochastic trend reflects some of these factors as well as shifts in credit market architecture.

--- Table 4 About Here ---

A range of alternative specifications, with and without the stochastic trend, are all consistent with the view that the marginal propensity to consume out of housing wealth in recent years has substantially exceeded that out of illiquid financial wealth. This supports the claim by Case et al. (2005) that the housing wealth (or collateral) effect in the US exceeds the stock market wealth effect. These results are also consistent with Benjamin et al (2004) and with Carroll et al. (2006). As noted above, the latter separate assets into two kinds, broadly illiquid financial wealth and the rest, largely net liquid assets plus housing wealth.

In the UK, the marginal propensity to consume for net liquid assets is relatively accurately estimated at 0.11, and the US estimate of 0.072 in Table 4, column 5 is only around one standard error away from this value, broadly consistent with the microeconomic evidence from Gross and Souleles (2002). However, the marginal propensity to consume for housing collateral in the UK in recent years is close to 0.032 while most values estimated for the US, as well as ours, are

[20] To check for possible short term asymmetries, the absolute value of the 4-quarter change in housing wealth/income interacted with the CCI proxy was included but found to be insignificant, t=0.3.

substantially higher. Our most plausible estimate of 0.055 has a substantial margin of uncertainty around it. One can ask whether it is plausible that the US figure could be twice or more that for housing in the UK. Transactions costs are broadly similar: the higher fees of US real estate agents being offset in the UK by higher transactions tax rates (Stamp Duty).

However, the US differs from the UK in three major ways. The first is that mortgage interest, even on second homes, is fully tax deductible in the US, while the UK, after heavily capping tax deductibility for many years, gradually eliminated it. This creates a relative incentive in the US for loading as much debt as possible on home equity. The second major difference is that the fixed rate mortgage system is highly effective in protecting US households from interest rate risk on both mortgage debt that finances housing and that used to finance non-housing consumption via "cash-out" mortgage refinancing. The U.S. system protects households from rising rates and gives them a low cost option to refinance when rates drop. The view that there was an implicit government guarantee underwriting Fannie Mae and Freddie Mac, and the existence of a deep financial system permitting prepayment risk to be hedged effectively through the government bond market and elsewhere, explain the low cost of this option, see Green and Wachter (2007). In contrast, as Miles (2004) makes clear, the high penalty charges for refinancing in the UK have discouraged demand for fixed rate mortgages. Thirdly, in most US states there is a 'walk away' option for households with negative housing equity: they can simply hand in the keys to their home to the mortgage lender and be free of further debt service obligations. In the UK, in contrast, borrowers can be pursued for seven years for any debt not covered by the sale of their repossessed home. Together, the tax and risk advantages of US mortgages make it plausible that the marginal propensity to consume for home equity should be significantly larger in the US than in the UK.

Our aim in future work is to sharpen the US estimate with more accurate estimates of a credit conditions index using the multiple indicator, common factor approach with a set of credit related equations. The policy relevance of such estimates is clearly high. Currently, both a contraction in credit availability and a fall in housing collateral are constraining US consumer spending. Our model, which also includes stock market wealth effects and the change in the unemployment rate throws a great deal of light on the short to medium term outlook for US consumer spending.

6. Concluding comments

Credit constraints can have important theoretical implications for consumer spending that are borne out in our empirical findings for the UK, U.S. and Japan. We find that the evolution of credit availability can differ across countries and may change over time within countries. In Japan, the consumption function has been very stable, reflecting a lack of financial liberalization since the 1970s. Consistent with theory, the absence of major changes easing the availability of finance to households, and differences in the tax code, likely accounts for why rising home prices tend to restrain consumption in Japan.

In contrast, there have been very large changes in the availability of credit in the UK and the U.S. over the last few decades. Greater access to credit has resulted in an upward shift of the consumption function in both countries, as reflected in consumption coefficients on credit conditions indexes for each nation. Furthermore, financial liberalization is linked to a positive impact of housing wealth on consumption that has become more enhanced in both countries over time and to changes in some other consumption coefficients.

As a result, the impact of large, recent declines in wealth, particularly housing equity, will likely have larger dampening effects on consumer spending in the UK and the U.S. for some time. Of course, some of the swings in housing wealth reflect the impact of changes in mortgage credit standards as shown in some of our other work in progress (Duca et al., 2008). These negative wealth effects are being compounded by a related and substantial tightening of consumer credit standards in the U.S., a combination not seen since 1974-75, when consumption was unusually weak. In both instances, the availability of mortgages fell sharply, reflecting the recent collapse of subprime mortgage lending and the disintermediation that deposit rate regulations induced in the mid-1970s. Although energy prices have recently declined in stark contrast to that earlier episode, consumer credit availability has plunged more in the recent episode. Moreover, much of the recent credit tightening reflects the effects of loan losses that will likely persist for some time, rather than the Regulation Q effects of the mid-1970s that quickly unwound when interest rates fell. In the UK, there are also indications that the availability of credit has been sharply curtailed in recent quarters amid very sizable declines in house and other asset prices.

Although Japanese consumer spending will likely be less *directly* affected by declines in housing wealth and credit availability in Japan, the impact of the global economic downturn, particularly in the U.S., is affecting Japanese household income via a slowdown in net exports. As others have demonstrated (Greenlaw, Hatzius, Kashyap, and Shin, 2008), the damage that loan losses are causing to financial institutions is also large enough to induce credit tightening outside of the U.S. Hence, the unwinding of the Anglo-American credit-fuelled consumption boom is impacting the world economy via lowering net exports and raising financial frictions. Nevertheless, it is important to make cross-country and time series distinctions regarding the nature of the current credit crisis and its impact on the world economy.

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Figure 1: The Channels of Transmission of the Mortgage and Housing Crisis

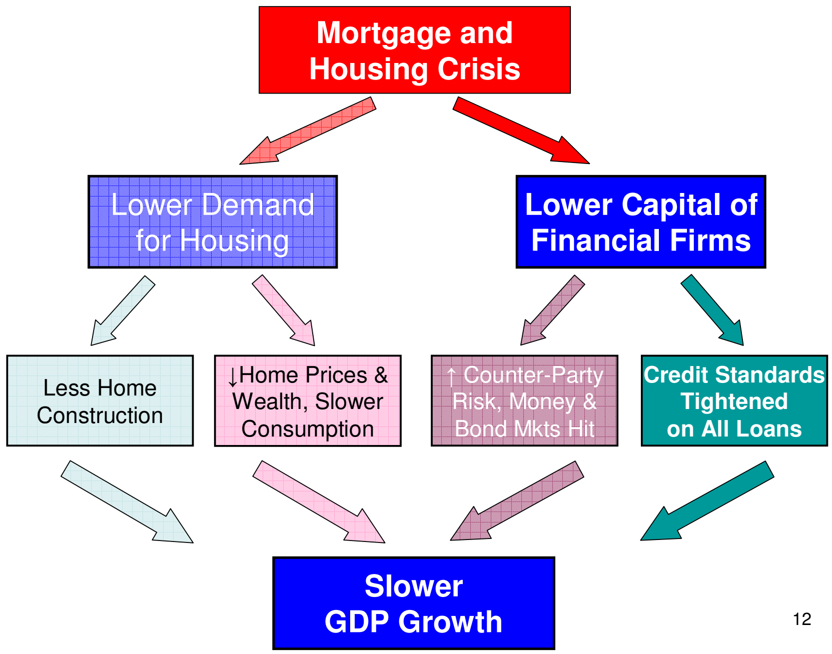
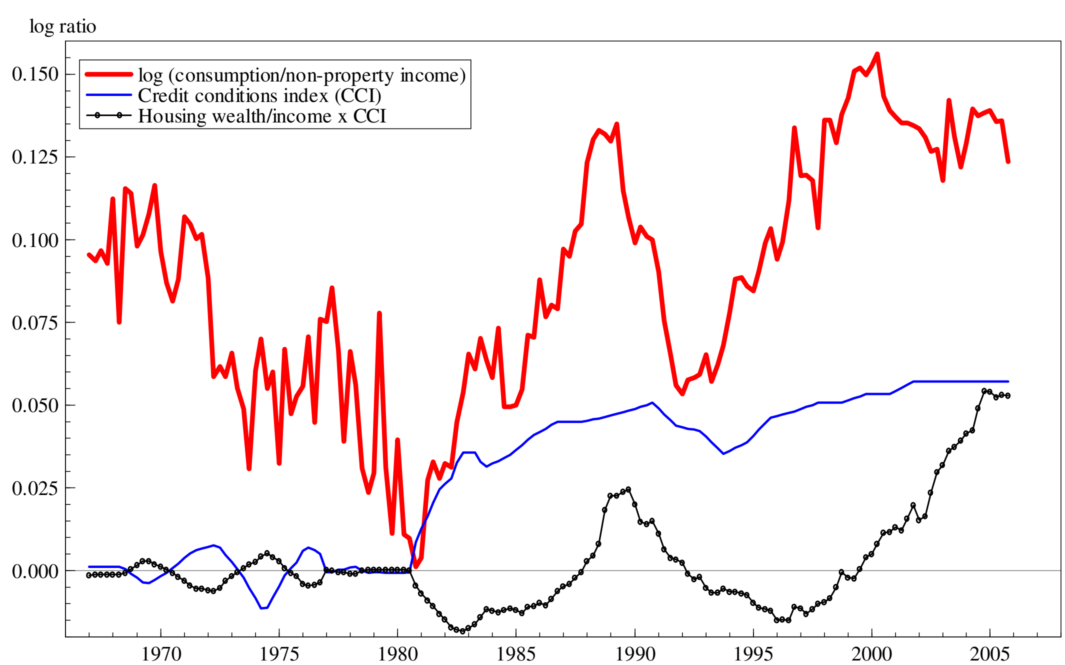
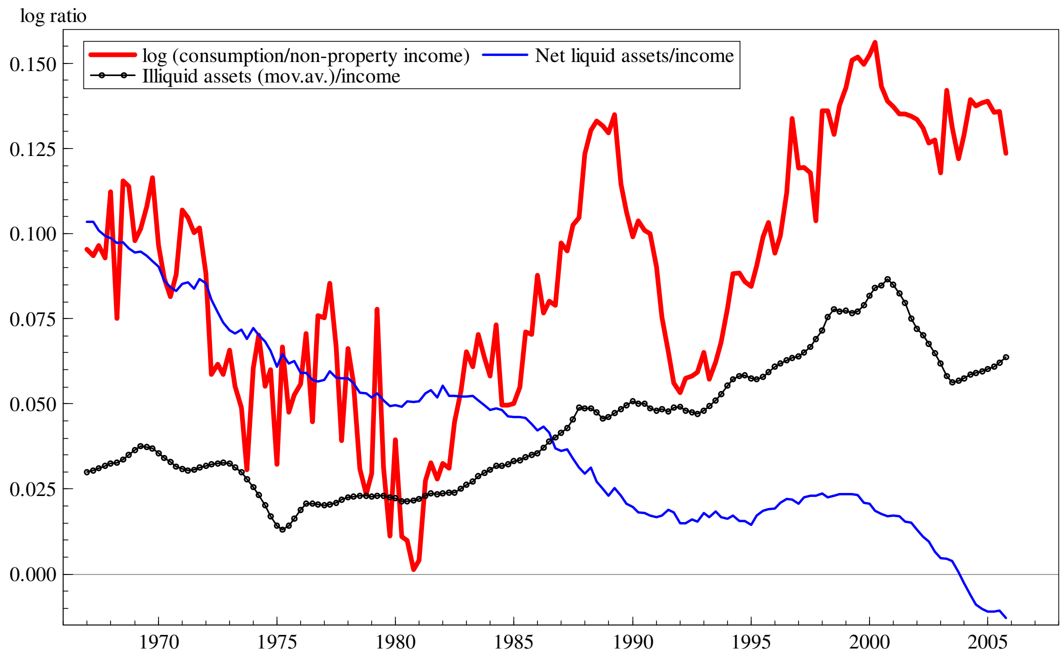


Figure 2: Estimated long-run contributions to log consumption/income of the credit conditions index and its interaction with housing wealth/income in the UK .



Note: asset to income ratios are defined as end of last quarter assets/4 (current quarterly non-property income). Explanatory variables are scaled by the estimated coefficients in the long-run solution. See footnote to Table 1 for the definition of the interaction between CCI and housing wealth/income.

Figure 3: Estimated long-run contributions to log consumption/income of net liquid assets/income and illiquid financial assets/income in the UK.



Note: asset to income ratios are defined as end of last quarter assets/4 (current quarterly non-property income). Explanatory variables are scaled by the estimated coefficients in the long-run solution.

Figure 4: Estimated long-run contribution to log consumption/income of net financial assets/income and log real land prices in Japan.

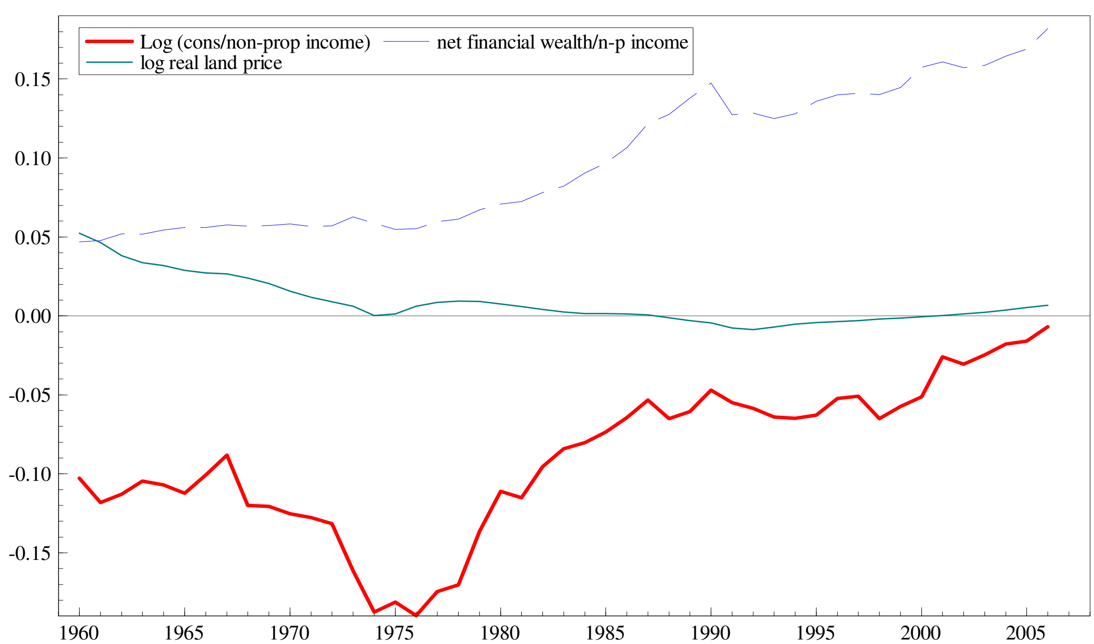


Figure 5: Estimated long-run contribution to log consumption/income of real interest rate and forecast income growth in Japan,

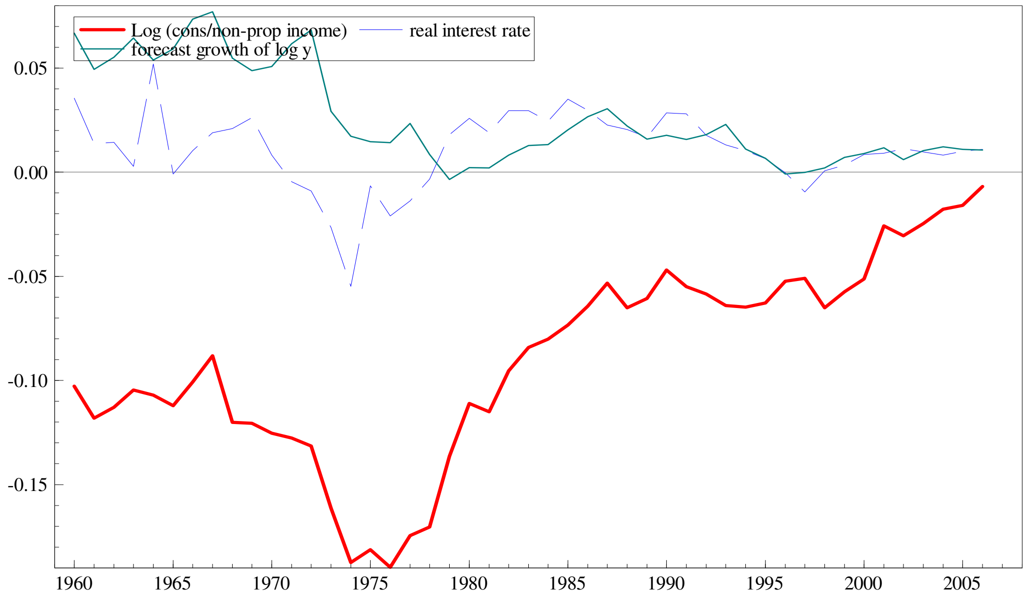


Figure 6: Fitted and actual DLYPERM for the US

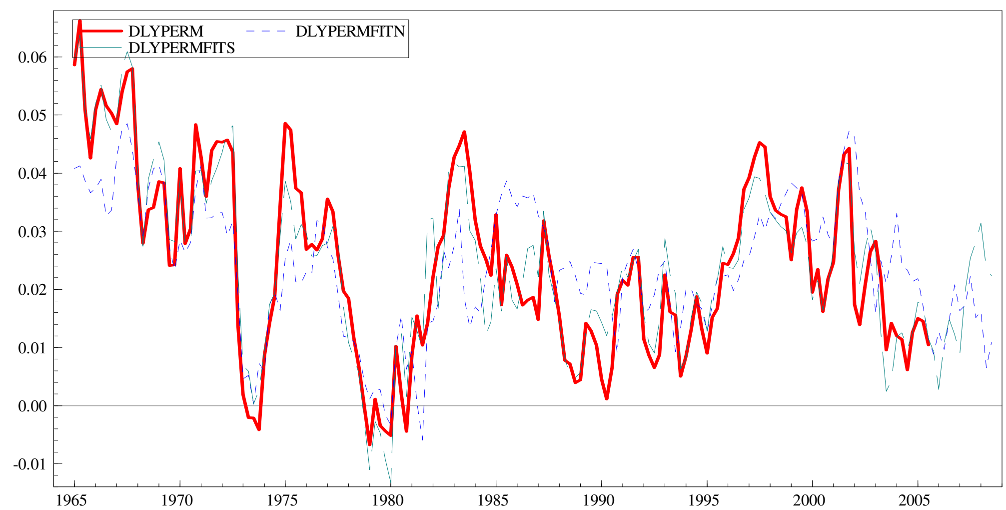


Figure 7: Credit Conditions Index for the US

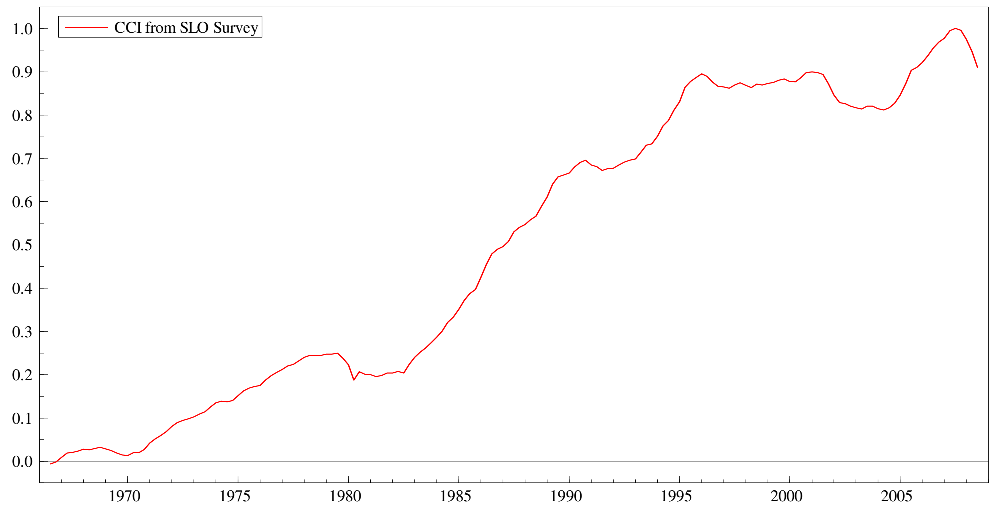


Figure 8: Fitted stochastic trend for the US

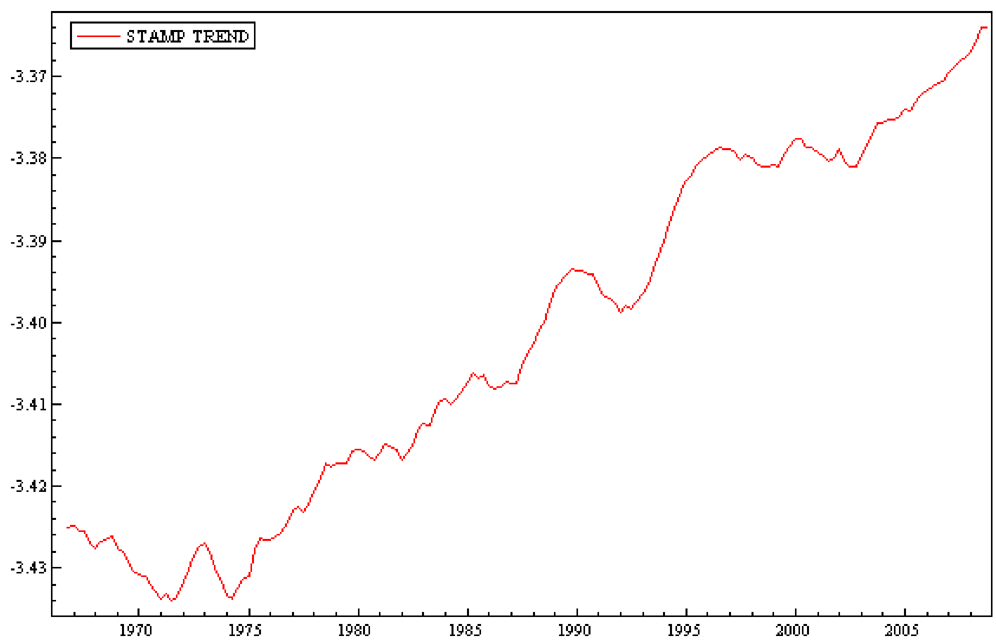


Table 1: Estimates of the UK Consumption Function for 1967 Q1 to 2005 Q4

Dependent variable = $\Delta \ln c$	*(1)*	*(2)*	*(3)*	*(4)*
$\ln y - \ln c_{-1}$	0.16 (4.9)	0.23 (6.3)	0.31 (8.5)	0.33 (8.9)
Credit conditions index *CCI*	-		-	0.020 (3.5)
Net liquid assets/income	0.0036 (4.8)	0.026 (4,3)	0.033 (5.6)	0.038 (5.6)
Illiquid financial assets/income	Ditto	0.0076 (5.9)	0.0071 (5.1)	0.0061 (4.7)
Housing assets/income	Ditto	Ditto	0.0111 (5.7)	-
Housing assets/income and *CCI* interaction	Ditto	-	-	0.0106 (6.3)
Expected income growth	0.21 (4.9)	0.18 (4.4)	0.22 (5.5)	0.10 (2.4)
Expected income growth and *CCI* interaction	-	-	-	0.15 (1.6)
Real mortgage interest rate	-	-	-0.03 (1.4)	-0.04 (1.5)
Change in unemployment rate $\Delta_4 ur$	-	-	-0.56 (7.8)	-0.64 (8.6)
Debt/income and Δ_4 nominal interest rate interaction	-	-	-0.0029 (3.8)	-0.0072 (3.1)
Debt/income, Δ_4 nom. interest rate and *CCI* interaction	-	-	-	0.0057 (1.9)
Diagnostics				
Standard Error * 100	0.76	0.73	0.61	0.59
R^2	0.58	0.62	0.73	0.76
D.W.	1.41	1.34	1.85	1.98
P-Value for No Structural Break	0.11	0.064	0.093	0.754

Notes: The interaction effect with housing assets/income, takes the form (housing assets/income - 3.08) * *CCI*, where 3.08 is the mean value of housing wealth/income for 1980-2005. The t-statistics are in parentheses.

Table 2: Estimates of the Forecasting Equation for $\Delta \ln y_{t+1}$ for Japan

Dependent Variable = $\Delta \ln y_{+1}$	*(1)* ***1959-2005***	*(2)* ***1959-2009***	*(3)* ***1959-1992***	*(4)* ***1975-2005***	*(5)* ***1959-2005***
Intercept	-3.814 (0.69)	-3.547 (0.62)	-4.121 (0.72)	-3.144 (0.64)	-4.440 (0.67)
Trend	0.028 (0.01)	0.024 (0.01)	0.025 (0.01)		0.026 (0.01)
Split Trend in 1993	-0.024 (0.01)	-0.022 (0.01)	-0.023 (0.01)		-0.024 (0.01)
$\ln y$	-0.458 (0.10)	-0.404 (0.07)	-0.446 (0.10)	-0.356 (0.07)	-0.473 (0.11)
3 Year Change in Nominal Call Rate $\Delta_3 nr$	-0.199 (0.06)	-0.212 (0.06)	-0.203 (0.08)	-0.159 (0.06)	-0.235 (0.06)
$\ln US\ GDP_{-1}$	0.174 (0.09)	0.178 (0.09)	0.223 (0.10)	0.205 (0.04)	0.246 (0.09)
$MA3\ (\text{Govt Bal/GDP})_{-1}$	0.538 (0.10)	0.628 (0.11)	0.593 (0.14)	0.612 (0.10)	
$(\text{Govt Debt/GDP})_{-1}$	-0.032 (0.03)				-0.165 (0.03)
$(\text{Govt Debt/GDP})_{-4}$					0.106 (0.02)
Diagnostics					
Standard Error * 100	1.20	1.20	1.36	0.91	1.18
Adjusted R^2	0.87	0.87	0.84	0.60	0.88
Durbin Watson	1.93	2.03	2.01	2.24	1.91
AR1/MA1 (p-value)	0.83	0.87	0.90	0.48	0.79
AR1/MA1 (p-value)	0.26	0.15	0.14	0.13	0.12
Hetersocedasticity (p-value)	0.001	0.003	0.035	0.29	0.002
Chow (p value)	0.85	0.67	0.51	0.71	0.86
Reset (p value)	0.82	0.26	0.64	0.50	0.75

Notes: Robust standard errors in parentheses.

Table 3: Japanese Consumption Function Estimates for 1961 -2006

Dependent variable = $\Delta \ln c$ ***:***	***(1)***	***(2)***	***(3)***	***(4)***
Intercept	-0.055	-0.057	-0.058	-0.063
	(0.017)	(0.016)	(0.015)	(0.015)
$\ln y - \ln c_{-1}$	0.356	0.345	0.347	0.359
	(0.067)	(0.064)	(0.063)	(0.064)
Income Growth $\Delta \ln y$	0.289	0.321	0.323	0.332
	(0.070)	(0.068)	(0.067)	(0.068)
Forecast Income Growth $\Delta \widehat{\ln} y$	0.367	0.347	0.348	0.350
	(0.083)	(0.079)	(0.078)	(0.079)
Income Growth Volatility	-0.225	-0.024		
	(0.094)	(0.128)		
Income Growth Volatility (using $\Delta \widehat{\ln} y$)		-5.666	-6.007	-5.648
		(2.574)	(1.782)	(1.796)
Change in Unemployment Rate	-0.008	-0.007	-0.007	
	(0.005)	(0.005)	(0.004)	
Real Interest Rate (Tax Adjusted)	0.346	0.346	0.350	0.367
	(0.062)	(0.059)	(0.054)	(0.054)
Net Financial Wealth$_{-1}$/Income	0.022	0.022	0.023	0.024
	(0.006)	(0.005)	(0.005)	(0.005)
Log Real Land Price$_{-1}$	-0.014	-0.015	-0.015	-0.016
	(0.004)	(0.004)	(0.004)	(0.004)
Standard Error * 100	0.681	0.648	0.640	0.650
Adjusted R Squared	0.941	0.947	0.948	0.946
Durbin Watson	2.14	2.20	2.20	2.22
AR1/MA1 (p-value)	0.621	0.386	0.417	0.396
AR2/MA2 (p-value)	0.742	0.711	0.717	0.726
Heteroscedasticity (p-value)	0.737	0.849	0.829	0.955
Chow (p-value)	0.255	0.191	0.298	0.635
RESET(p-value)	0.066	0.445	0.576	0.827

Notes: Robust standard errors in parentheses.

Table 4: US Consumption Function Estimates for 1966 Q3 to 2008 Q3

Dependent variable = $\Delta \ln c$	*(1)* **OLS**	*(2)* **OLS**	*(3)* **OLS**	*(4)* **OLS**	*(5)* **STAMP**
Intercept	-0.368** (2.4)	-0.493** (3.5)	-1.851** (5.8)	-1.752** (5.6)	n.a.
Credit conditions index *CCI*			0.020** (4.7)	0.021** (4.3)	n.a.
$\Delta \ln y$	0.290** (4.7)	0.188** (3.3)	0.134* (2.5)	0.203** (3.6)	-
$\ln y - \ln c_{-1}$	0.052* (2.5)	0.071** (3.5)	0.269** (5.9)	0.255** (5.7)	0.507** (9.4)
$E_t \Delta \log ym_{t+k}$ naive	0.135* (2.4)	0.101* (2.0)	0.137** (2.7)	0.100+ (1.8)	0.063 (0.9)
$E_t \Delta \log ym_{t+k}$ sophisticated	-0.030 (0.7)	-0.020 (0.5)	0.087+ (1.9)	0.106* (2.1)	0.139* (2.4)
Net wealth / income NW_{-1}/Y	0.0026+ (1.9)	0.0033** (2.8)	-	-	-
Net liquid assets / income NLA_{-1}/Y			0.005 (1.3)	0.006 (1.4)	0.037* (2.0)
Housing wealth / income $HsgW_{-1}/Y$			0.013** (3.5)	0.011* (2.0)	-
Demeaned $HsgW_{-1}/Y$ interacted with *CCI*				0.0016 (0.3)	0.028* (2.0)
Illiquid financial assets / income $(\frac{1}{4}\sum_1^4 IFA_{-s})/Y$			0.0044** (3.1)	0.0038** (2.7)	0.0090** (3.1)
$\Delta^2 (HsgW_{-1}/Y)$				0.035** (2.9)	0.016+ (1.6)
Demeaned $\Delta_4 (HsgW_{-1}/Y)$ interacted with *CCI*				0.015+ (1.6)	0.030+ (1.8)
Change in unemployment Δur		-0.767** (5.9)	-0.801** (6.5)	-0.712** (5.7)	-0.417** (2.8)
Real interest rate on autos r_{auto}		-0.013 (0.4)	-0.041 (1.3)	-0.035 (1.1)	-0.175** (3.1)
Change in nominal interest rate Δnr_{auto}		-0.437** (5.1)	-0.255** (2.9)	-0.263** (3.0)	-0.176* (2.0)
1980 Q2 Dummy	-0.024** (4.4)	-0.009+ (1.9)	-0.010* (2.0)	-0.011* (2.4)	-0.012* (2.6)
Diagnostics					
Standard Error * 100	0.529	0.454	0.427	0.416	0.430
R^2	0.38	0.55	0.61	0.63	0.73
DW	1.89	2.18	2.11	2.13	2.12
AR(2)/MA(2) (P Value)	0.02	0.28	0.44	0.64	-

Note: The STAMP specification uses an I(1) stochastic trend. The t-statistics are in parentheses. The symbols +, * and ** denote significance at the 90%, 95% and 99% confidence levels, respectively.

Appendix: Constructing a U.S. Credit Conditions Index

Before constructing a levels index from this relative change index, we first adjust it for identifiable effects of interest rates and the macroeconomic outlook by estimating an empirical model based on screening models. In such models (see Duca and Garrett, 1995; and the screening model of Stiglitz and Weiss, part IV, 1981), credit standards should be tightened when the real riskless rate rises and the macroeconomic outlook worsens. (Since the willingness to lend index is inversely related to credit standards, these expected signs are reversed in our empirical model of the diffusion index.) we track the former by including the t and t-1 lags of the first difference of the real federal funds rate (*ΔRFF*, the nominal funds rate minus the year-over-year percent change in the overall PCE deflator), and the latter by the two-quarter percent change in the index of leading economic indicators (*ΔLEI2*). To further adjust for factors affecting consumer loan quality, we include the time t year-over-year change in the delinquency rate on all consumer instalment loans at banks (*Δ4DEL*, American Bankers Association). Also included were two variables to control for the impact of regulation. One was a dummy equal to 1 in 1980:q2 when credit controls were imposed and equal to -1 when they were lifted in 1980:q3 (*DCON*). The second is a variable (*REGQ*) consistently measuring the degree to which Regulation Q ceilings impinged upon banks' ability to raise small time deposit rates (see Duca, 1996; and Duca and Wu, forthcoming) and thereby raised banks' shadow cost of raising loanable funds (Regulation Q was binding during an era before the loan sales and mortgage-backed securities markets were deep). After Reg Q was lifted, the interbank funding market increasingly became a marginal source of loanable funds, with the three-month LIBOR normally exceeding the expected 3-month average federal funds rate by about 10-12 basis points. However, as out-sized fears of bank portfolio losses and liquidity premiums soared during the current financial crisis, so did these spreads, which raised one marginal cost of funding loans. To control for this, we include the t and t-1 spreads between the 3-month LIBOR and 3-month OIS rates (*LIBOR3*).

Estimating the model with an AR(1) correction yielded the following estimates:

$$CR = \underset{(4.59)}{19.67} - \underset{(-6.05)}{4.79}*\Delta RFF_{t}^{**} - \underset{(-3.28)}{2.10}*\Delta RFF_{t-1}^{**} + \underset{(5.91)}{1.43}*\Delta LEI2_{t}^{**}$$

$$- \underset{(-3.20)}{14.11}*\Delta 4DEL_{t}^{**} - \underset{(-2.21)}{2.38}* REGQ^{*}$$

$$- \underset{(-10.26)}{47.16}*DCON_{t}^{**} - \underset{(-2.01)}{47.27}* LIBOR3_{t}^{*} - \underset{(-1.52)}{37.03}* LIBOR3_{t-1}$$

where t-statistics are in parentheses, $\bar{R}^2 = 0.798$, standard error = 9.06, LM(2) = 0.69, and Q(24) = 17.49. The coefficients all have the anticipated signs and are all significant, with the exception of the t-1 lag of *LIBOR3*, which, however, is jointly significant with its time t lag. We then subtracted the estimated impact of changes in the real federal funds rate, leading economic indicators, and delinquency rate to remove normal business cycle and interest rate effects, leaving the impact of regulations, unusual credit frictions in the LIBOR market, and unexplained movements in the adjusted diffusion index (*CRAdj*). We transformed *CRAdj* into a levels CCI based on the .01 positive and statistically significant correlation of the CR index with the growth rate of real per capita consumer loan extensions (available 1966-1982:q4). Setting 1966:q2 = 100 and using this correlation, we then allowed the CCI to multiplicatively evolve according to $CCI_{t} = CCI_{t-1}*(1+(.01)*CRAdj)$. The multiplicative aggregation is consistent with the senior loan officer question asking whether a bank's willingness to make consumer loans was much more, somewhat more, unchanged, somewhat less, or much less *compared* to three months earlier.

The resulting CCI rises greatly during the 1980s, and then rises over 2004-06, before reversing the gains of the early decade since 2006.