the review of income and wealth

Review of Income and Wealth
Series 58, Number 3, September 2012
DOI: 10.1111/j.1475-4991.2011.00466.x

CREDIT, HOUSING COLLATERAL, AND CONSUMPTION: EVIDENCE FROM JAPAN, THE U.K., AND THE U.S.

BY JANINE ARON

Department of Economics, and Institute for New Economic Thinking, Oxford Martin School, University of Oxford

JOHN V. DUCA

Federal Reserve Bank of Dallas & Southern Methodist University, USA

JOHN MUELLBAUER*

Nuffield College and Institute for New Economic Thinking, Oxford Martin School, University of Oxford

KEIKO MURATA

Tokyo Metropolitan University, Japan

AND

ANTHONY MURPHY

Federal Reserve Bank of Dallas, USA

The consumption behavior of U.K., 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 U.K. and U.S., but not Japanese, households has undergone large shifts since 1980. The average consumption-to-income ratio rose in the U.K. 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 U.K. In Japan, land prices (which proxy house prices) continue to negatively impact consumer spending. There are negative real interest rate effects on consumption in the U.K. and U.S. and positive effects in Japan. Overall, this implies important differences in the transmission of monetary and credit shocks in Japan versus the U.S., U.K., and other credit-liberalized economies.

JEL Codes: E21, E32, E44, E51

Keywords: consumption, credit conditions, housing collateral, housing wealth

Note: 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. Aron and Muellbauer's research was supported by grants from the Open Society Institute and the Oxford Martin School. 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.

*Correspondence to: John Muellbauer, Nuffield College, New Road, Oxford OX1 1NF, U.K. (john.muellbauer@nuffield.ox.ac.uk).

Published by Blackwell Publishing, 9600 Garsington Road, Oxford OX4 2DQ, UK and 350 Main St, Malden, MA, 02148, USA.

1. Introduction

The global economic crisis of 2007–09 had its origins in a credit crisis, at the heart of which is asymmetric information between lenders and borrowers, most simply the risk that lenders face concerning the ability and willingness of borrowers to service debt (see Stiglitz and Weiss, 1981; Duca *et al.*, 2010). 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 mainstream macroeconomists. Many dynamic stochastic general equilibrium (DSGE) modelers 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 modeling 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 on credit flows and changing household wealth portfolios from the flow of funds and balance sheets, now receiving far more attention from central banks (González-Páramo, 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 U.K., 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 use disaggregated household balance sheet data, and address important measurement issues both for income and for wealth.

Early work attributed much of the fall in the U.K. household saving rate to credit market liberalization, increases in house prices, and the greater "spendability" of housing wealth with liberalization (Muellbauer and Murphy, 1989, 1990; Miles, 1994).[1] Muellbauer and Lattimore (1995) reviewed further research and laid out the foundations of a solved-out consumption function, encompassing the classical life-cycle/permanent income hypothesis and a credit channel.[2]

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. Jappelli and Pagano (1994), Engelhardt (1996), and others demonstrated that mortgage down-payment constraints generate an economically significant motive to save. In countries with limited access to consumer and/or mortgage credit, such

[1]Elsewhere we have modeled the roles of credit supply and extrapolative expectations in explaining house prices (Muellbauer and Murphy, 1997; Cameron *et al.*, 2006; Duca *et al.*, 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 modeled expectations, the academic literature in macroeconomics has been dominated by approaches based on Euler equations with representative agents, as in many DSGE models.

as Italy and Japan, higher home prices can induce higher saving for down-payments, thereby generating negative housing collateral or "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 U.K. and U.S. First, credit liberalization lowers the down-payments required of first-time home buyers. Second, it provides those home owners 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 U.K., the most systematic estimates are in Fernandez-Corugedo and Muellbauer (2006), who jointly modeled 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 U.K. households.

In models of aggregate U.K. consumption, the *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 modeled using a simple but robust income forecasting equation. By interacting the *GCCI* with other variables such as housing wealth, the associated parameters are found to shift with credit market liberalization, in line with theory. Moreover, by including *GCCI* effects, the remaining model parameters become stable over the four decades from 1967 to 2005, with co-integration tests easily passed.

A similar approach is adopted when modeling aggregate U.S. consumption, except that actual household survey data are used to model income expectations. The long historical series 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 U.K. and Japan. For the U.S. consumption function, as for the U.K., 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 *et al.*, 2006). Our estimates suggest that it would be difficult to find co-integration of aggregate U.S. consumption, income, and wealth holdings for the last 40 years, 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, we find no evidence of

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

any parameter shifts in the consumption function 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 U.K. 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 U.K. 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 pertinent to the economic crisis of 2007–09. For example, our U.K. consumption function made it possible to predict by mid-2008 that the U.K. 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). An earlier and mistaken view from the Bank of England 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 incorporated powerful housing and stock market "wealth" effects. Along with a greater appreciation of the relevance of the financial accelerator amongst U.S. policymakers, this helped contribute to an early and decisive monetary policy response to the crisis.

Our results for Japan explain why the household component of the monetary transmission channel is far weaker there than in the U.K. or U.S. Had this been better understood in 2001–04, U.S. monetary policy may have been less concerned about the risk of a Japan-style "lost decade" over 2000–09. Leamer (2007) and Taylor (2007), amongst others, argue that, as a result, the federal funds rate was kept too low for too long. Concurrently, there was an unsustainable liberalization of the mortgage market that fuelled an unexpectedly strong credit, housing, and consumption boom, the collapse of which took years to play out.

Macroeconomic research using aggregate time-series and balance sheet data has been less fashionable in the last two decades, but recent events strongly 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 forges some of the links.

2. The Consumption Function

This section begins by demonstrating the weakness of the conventional housing wealth effect in the classical life-cycle model of consumption. We then

[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), now using an extended sample and a better, more forward-looking income forecasting model, *inter alia*.

discuss the housing credit channel for 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; Pagano, 1990). The lack of a strong positive housing wealth effect in standard frameworks can be shown using a stylized 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 maximizes life-cycle utility defined on the flows of non-housing 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:

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

where a 0 subscript denotes previous period levels, and $p_h(r + \delta)$ = real user cost of housing. We are interested in the effects of a change in p_h on a constant price index of total consumption, $c + p_h(r + \delta)H$, as found in the National Accounts. This index, of the Paasche type, since it uses current reference prices, includes the imputed rent on housing. Differentiating with respect to p_h, we find:

$$\partial c/\partial p_h + p_h(r+\delta)\partial H/\partial p_h = rH_0 - (r+\delta)H. \tag{2.2}$$

But with $H \approx H_0$, the initial housing stock, the right-hand side of equation (2.2) is negative. This point was overlooked in the classic work by Modigliani and Brumberg (1954, 1980), Friedman (1957, 1963), and Ando and Modigliani (1963), and hence they overstated the effect of increasing house prices on consumption.

In models with finite lives and transactions costs, the effect of a rise in house prices on total consumption is likely to be less negative than in equation (2.2) and depends on how well imputed rent $p_h(r + \delta)H$ is measured in the National Accounts. Yet even taking account of these factors, it would be difficult to generate a substantial positive aggregate housing wealth effect from classical life-cycle permanent income theory.[5] For non-housing consumption, a modest positive

[5]Buiter (2010) makes a similar point in the context of a Blanchard–Yaari overlapping generations model, assuming the utility function is additive in log non-housing consumption and log housing services. However, he agrees, as we argue in Section 2.2, that a change in house prices can more strongly affect aggregate consumption through "collateral or credit effects due to the collateralisability of housing wealth and the non-collateralisability of human wealth" (p. 26).

house price 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 Jappelli and Pagano, 1994; Engelhardt, 1996; Deaton, 1999).[6]

Now consider the impact on consumption of higher house prices via the operation of the down-payment constraint. With limited 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, financial deregulation and the 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 behavior is governed by the life-cycle model outlined above, will display a small negative consumption 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 the ability to borrow against housing wealth increases.

In countries like the U.K. 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

[6]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.

influenced by changes in the debt service burden, which can be well tracked by 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 weighted by the debt-to-income ratio should thus weaken with credit market liberalization, but become larger in a credit crunch. By contrast, greater access to unsecured credit should increase the role of *inter-temporal substitution*, enhancing the role of income growth expectations and making the overall 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:

$$c_{it} = \phi_{it} A_{it-1} + \omega_{it} y_{it}, \tag{2.3}$$

where ϕ_{it} and ω_{it} vary by age, and i is a household subscript. Hence average per capita consumption, where N is the number of households, is:

$$c_t = \frac{1}{N_t}\sum_i c_{it} = \left(\frac{\sum_i \phi_{it} A_{it-1}}{\sum_i A_{it-1}}\right)\frac{1}{N_t}\sum_i A_{it-1} + \left(\frac{\sum_i \omega_{it} y_{it}}{\sum_i y_{it}}\right)\frac{1}{N_t}\sum_i y_{it} \equiv \phi_t A_{t-1} + \omega_t y_t. \tag{2.4}$$

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

In practice, ϕ and ω evolve only slowly with life expectancy and 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 ϕ and ω are constant have very similar implications and quality of fit to models where they evolve according to the survey data. Furthermore, as households make long-run portfolio decisions, the level and composition of assets are likely to reflect the demographic evolution, implying that shifts in ϕ and ω due to demographic change have a less direct impact on consumption. Accordingly, in the next section, we simplify by assuming that ϕ and ω are constant.

2.4. *A Solved Out Consumption Function*

Ando–Modigliani–Brumberg and Friedman style consumption functions require an income forecasting model to generate permanent non-property income.

Unlike in the Euler equation (Hall, 1978), long-run information on income and assets is not ignored. As a result, the solved out consumption function has advantages for policy modeling and forecasting. When the real interest rate is constant, the basic aggregate life-cycle/permanent income consumption function has the form:

$$(2.5)\qquad c_t = \phi A_{t-1} + \omega y_t^P,$$

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 choose continuously and optimally between current and future consumption. Strong multi-country evidence against this fundamental prediction is found by Campbell and Mankiw (1989, 1991), amongst others. Equation (2.5) is less restrictive, since it is consistent with only a rudimentary comprehension of life-cycle budget constraints. Any household with some notion of wanting to sustain consumption will realize that not all of the 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, after a little manipulation, gives:

$$(2.6)\qquad \frac{c_t}{y_t} = \omega\left(\frac{\phi}{\omega}\frac{A_{t-1}}{y_t} + 1 + \frac{y_t^P - y_t}{y_t}\right).$$

The right-hand side of equation (2.6) has the form $1 + x$, where x is usually a fairly small number. We can then take logs, use the fact[7] that $\ln(1 + x) \approx x$ and the further approximation $(y^P - y)/y \approx \ln(y^p/y)$, to obtain:

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

where $\gamma = \phi/\omega$ and $\alpha_0 = \ln\omega$. 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. Demography can be proxied by variables such as the population proportions in different age groups.[8] The log ratio of permanent to current income reflects expected income growth and can be proxied using forecasted income and the approximation:

$$(2.8)\qquad E_t \ln\left(y_t^p/y_t\right) \approx \frac{E_t \sum_{s=1}^{K} \eta^{s-1} \ln\left(y_{t+s}/y_t\right)}{\sum_{s=1}^{K} \eta^{s-1}} \equiv E_t \ln yperm_t - \ln y_t,$$

[7]Where x is not "small," a second order approximation based on $\ln(1 + x) \approx x - {}^1\!/_2 x^2$ is easy to apply.

[8]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 for Japan.

where ln $yperm_t$ – ln y_t is a weighted moving average of forward-looking income growth rates over K periods, and η is a discount factor (see Campbell, 1987).

The static consumption function (2.7) can be made dynamic by introducing habits or adjustment costs, resulting in a partial adjustment version of equation (2.7), as derived by Muellbauer (1988), amongst others. Allowing for probabilistic income expectations suggests adding a measure of income uncertainty, θ_t, to equation (2.7), and including a risk premium in the discount factor η in the expected income growth term, E_t ln $yperm_t$ – ln y_t (equation 2.8). 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 that the real interest rate r_t also enters equation (2.7), with the usual interpretation of inter-temporal substitution and income effects.

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

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

where λ measures the speed of adjustment. In a more complicated model, the parameters α_3 and γ should depend upon the real interest rate r_t.[9] They may also depend on θ_t, since discount rates applied to expected future incomes will increase with income uncertainty, as Skinner (1988), Zeldes (1989), Kimball (1990), and Carroll (1997, 2001) have emphasized. For simplicity, we will suppress these complication and the associated potential non-linearities.

Returning to equation (2.8), there are several practical reasons why the income growth expectations embodied in E_t ln $yperm_t$ – ln y_t are likely to reflect a limited horizon for K and a discount factor, η, that is substantially less than one. With aggregate data, it is difficult to forecast income beyond three or so years. 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 and trend. Furthermore, short horizons are suggested if households anticipate future credit constraints according to the buffer-stock theory of saving (see Deaton, 1991, 1992). Buffer-stock savings can be generated by other mechanisms. For instance, Carroll (2001) argues that precautionary behavior with uncertain "worst case scenarios" also generates buffer-stock saving. Then plausible calibrations of micro-behavior can give a practical income forecasting horizon as short as three years, as Friedman (1957, 1963) also suggested. In practice, we assume a discount rate of 5 percent per quarter, so that $\eta = 0.95$, in the empirical work below.

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

[9]Note that the definition of permanent income in equation (2.8) excludes current income. If it were included, equation (2.9) would be identical but for a somewhat higher value of α_3.

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

where 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 $\Delta nr_t(DB_{t-1}/y_t)$, 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 inclusion of the term in the log change of income can be rationalized by aggregating over credit constrained and unconstrained households. The change in the income uncertainty proxy θ_t is included with the short-term variables. Note that shifts in the availability of credit over time imply that many of the parameters in equation (2.10) are time-varying.

The credit channel enters the consumption function through the different MPCs 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 (Aron *et al.*, 2006). In principle, credit market liberalization should: (i) raise the intercept α_0, implying a higher level of $\ln(c/y)$; (ii) lower the real interest rate coefficient, α_1; (iii) raise α_3 by increasing the impact of expected income growth; and (iv) increase the MPC for housing collateral, γ_3. Credit market liberalization should also lower the current income growth effect, β_1, and the cash flow impact of changes in the nominal interest rate, making β_2 less negative. We handle these credit induced shifts in the U.K. consumption function by writing each of the time-varying parameters as a linear function of the index of general credit conditions, *GCCI*. Thus *GCCI* enters the model both as an intercept shift and interacted with several economic variables.

3. The Estimated U.K. Consumption Equation[10]

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, and includes durables and the imputed rent on owner occupied housing. Income is real per capita disposable non-property income.[11] The net worth to income ratio A/y is defined as liquid assets minus mortgage and other consumer debt plus net illiquid financial assets and housing wealth (using end of previous quarter asset levels), relative to current disposable non-property income.

In Table 1, Column 1 shows the textbook REPIH model with habits. This is essentially equation (2.9), but with the income uncertainty and real interest rate

[10]A data appendix with sources, statistics, and unit root tests for the U.K., U.S., and Japanese variables is available on request.

[11]This is defined as personal disposable income minus tax-adjusted property income. Permanent income growth, as defined in equation (2.8), is forecast using a time trend, current and lagged four quarter changes in log real income and its level, average log real share prices, reflecting their anticipatory role (see Poterba, 2000), and, from 1980 when credit liberalization began, average log real house prices. Details are available on request.

TABLE 1

U.K. Consumption Function Estimates for 1967 Q1 to 2005 Q4

Dependent Variable = $\Delta \ln c_t$	Symbol	(1)	(2)	(3)	(4)	(5)
Speed of adjustment	λ	0.078** (2.4)	0.122*** (3.6)	0.291*** (5.9)	0.371*** (7.1)	0.369*** (9.0)
Long-run coefficients						
Constant	α_0	–0.080** (2.1)	0.016 (0.4)	–0.166*** (7.0)	–0.041*** (3.0)	–0.041*** (3.0)
GCCI credit conditions index	α_{01}	–	–	–	0.050*** (3.6)	0.050*** (3.6)
Real mortgage rate (4 quarter average)	α_1	–	0.0007 (0.4)	–0.0017** (2.4)	–0.0017** (2.3)	–0.0017** (2.3)
Forecast future income growth	α_3	1.061*** (2.7)	0.597*** (3.5)	0.485*** (6.5)	0.201*** (2.9)	0.201*** (2.9)
Forecast future income growth × *GCCI*	α_{31}	–	–	–	0.252 (1.6)	0.254 (1.6)
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)	0.114*** (8.0)
Illiquid financial assets$_{t-1}$ (4 quarter average)/income	γ_2	–	–	0.026*** (7.2)	0.022*** (7.9)	0.022*** (8.0)
Housing wealth$_{t-1}$/income	γ_3	–	–	0.047*** (8.6)	–	–
Housing wealth$_{t-1}$/income × *GCCI*	γ_{31}	–	–	–	0.043*** (10.3)	0.043*** (10.3)
Short-run coefficients						
Current income growth	β_1	0.250*** (5.5)	0.175*** (4.0)	0.093** (2.1)	–0.003 (0.1)	–
Change in debt weighted nominal interest rate	β_2	–	–0.0020*** (3.3)	–0.0030*** (4.8)	–0.0061*** (3.8)	–0.0061*** (3.9)
Change in debt weighted nominal interest rate × *GCCI*	β_{21}	–	–	–	0.0041** (2.1)	0.0041** (2.1)
Change in unemployment rate	β_3	–	–0.0043*** (5.6)	–0.0058*** (7.2)	–0.0071*** (8.4)	–0.0071*** (9.6)
Dummies						
Change in credit controls	β_4	–0.0007*** (2.5)	–0.0010*** (3.7)	–0.0012*** (4.7)	–0.0009*** (3.6)	–0.0009*** (3.7)
Change in lagged working days lost to strikes	β_5	–0.0002*** (3.1)	–0.0002*** (3.7)	–0.0002*** (3.5)	–0.0002*** (3.4)	–0.0002*** (3.4)
1968 preannounced tax increase dummy	β_6	0.033*** (6.3)	0.033*** (7.0)	0.031*** (7.1)	0.030*** (7.1)	0.030*** (7.1)
1973 VAT introduction dummy	β_7	0.012*** (2.3)	0.012*** (2.4)	0.011*** (2.5)	0.011** (2.4)	0.011** (2.4)
1979 VAT rise dummy	β_8	0.045*** (8.6)	0.044*** (9.3)	0.040*** (8.9)	0.039*** (9.1)	0.039*** (9.2)
Q4 seasonal dummy	β_9	–0.022 (1.1)	–0.015 (1.3)	–0.008* (1.8)	–0.006** (2.1)	–0.0065** (2.1)
Diagnostics						
Standard error × 100		0.72	0.66	0.62	0.58	0.58
Adjusted R^2		0.60	0.67	0.71	0.74	0.74
Durbin Watson		1.78	1.98	1.91	1.96	1.96
AR1/MA1 (p-value)		0.16	0.92	0.62	0.79	0.80
AR4/MA4 (p-value)		0.06	0.76	0.20	0.10	0.10
Heteroscedasticity (p-value)		0.52	0.38	0.87	0.05	0.05
Chow (1985 Q1 break, p-value)		0.55	0.70	0.05	0.89	0.91
RESET (p-value)		0.50	0.82	0.29	0.13	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:

$$\begin{aligned}\Delta \ln c_t = \lambda\Big\{&\alpha_0 + \alpha_{01}GCCI_t + \left(\ln y_t - \ln c_{t-1}\right) + \alpha_1\overline{r}_t + \alpha_3\left(\widehat{\ln y^p_{t+1}} - \ln y_t\right) + \alpha_{31}GCCI_t \times \text{demeaned}\left(\widehat{\ln y^p_{t+1}} - \ln y_t\right)\\ &+ \gamma_1 NLA_{t-1}/y_t + \gamma_2\overline{IFA}_{t-1}/y_t + \gamma_3 HA_{t-1}/y_t + \gamma_{31}GCCI_t \times \text{demeaned } HA_{t-1}/y_t\Big\}\\ &+ \beta_1\Delta\ln y_t + \beta_2\Delta_4 nr_t \times DB_{t-1}/y_t + \beta_{21}GCCI_t \times \Delta_4 nr_t \times DB_{t-1}/y_t + \beta_3\Delta_4 ur_t + \beta_4\Delta Credit\ Control_t\\ &+ \beta_5\Delta WDL_{t-3} + \beta_6 d1968q1_t + \beta_7 d1973q1_t + \beta_8 d1979q2_t + \beta_9 q4_t + u_t\end{aligned}$$

where $\overline{r}_t$ and $\overline{IFA}_{t-1}$ are four quarter averages. $\Delta_4 nr_t$ is a weighted average of the four quarter changes in the nominal mortgage and unsecured borrowing rates, using the lagged shares of secured and unsecured borrowing as weights.

terms omitted. We obtain highly significant estimates of net worth and income growth expectations effects, as well as a low speed of adjustment of 0.08 per quarter.[12] The estimated long-run MPC out of net worth is approximately 2.5 percent (0.026). Column 2 adds the real interest rate, the debt-weighted change in nominal interest rates on debt, and the change in the unemployment rate, a proxy for income insecurity. The real interest rate is insignificant, but the changes in the unemployment rate and the nominal interest rate are both negative and significant. Column 3 relaxes the textbook 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 MPC out of net liquid assets equaling 0.126, far larger 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 an F test strongly rejects the textbook model in Column 1.

Finally, Columns 4 and 5 allow the relevant parameters of equation (2.10) to shift with *GCCI*, the general credit conditions index from Fernandez-Corugedo and Muellbauer (2006). The expected shifts in parameters occur (although some are insignificant) and there is a large improvement in fit over Column 3. 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.[13] The current income growth term is insignificant in Column 4, so it is dropped in Column 5. 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. The estimated housing "wealth" effect is lower than generally found in the literature. Most illiquid financial assets are in pension funds, so the model plausibly reflects the slow adaptation of contribution and pay-out rates to changes in asset values by using a four-quarter moving average of observations on illiquid financial assets. This fits a little better than the end of previous quarter value, consistent with Lettau and Ludvigson (2004).

The real interest rate effect in Column 5 is negative and significant. According to point estimates, not shown, this effect strengthens (becomes more negative) as *GCCI* rises. The effect of the changes in the debt-weighted nominal interest rate, also negative, weakens as *GCCI* rises. With easier access to credit, inter-temporal substitution should assume a bigger role, which can explain these two results and also 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. Its interaction effect with *GCCI* is positive, but insignificant, suggesting that higher debt levels approximately neutralize the reduced impact of income uncertainty on consumption caused by easier

[12]The specifications in Table 1 include an intercept, dummies for temporary consumption shifts due to VAT (sales tax) anticipations, and measures of the change in consumer credit controls for durables purchases and of working days lost in labor disputes. The details are in the data appendix, which is available on request.

[13]This interaction effect takes the form: (housing wealth/income minus its 1980–2005 mean) multiplied by *GCCI*.

access to credit. The quarterly speed of adjustment is 0.37, with over 80 percent of the adjustment of consumption to equilibrium occurring within four quarters.

The parameters of the equation in Column 5 are remarkably stable in charts of recursive estimates. The model can be interpreted in terms of co-integrated variables (treating the interaction between *GCCI* and the housing wealth-to-income ratio as one variable). The unique cointegrating vector consists of the log ratio of consumption to non-property income, the three asset-to-income ratios, and *GCCI*, which effectively shifts the intercept over time. Since the real interest rate is arguably I(0), and plays only a marginal role, it is not included in the cointegrating analysis. *GCCI* is treated as an exogenous shift dummy. I(0) variables such as income growth, forecast income growth, and the change in the unemployment rate are part of the system. Impulse dummies are also included, but outside the co-integration space. There is only one co-integrating relationship and this is close to the long-run solution implied by the Column 5 estimates.[14] For the U.K., therefore, the pessimism expressed by Lettau and Ludvigson (2004) and Carroll *et al.* (2011) regarding the existence of a cointegrating relationship between consumption, income, and assets appears to be misplaced, at least once the effect of the changing credit conditions, proxied by *GCCI*, is accounted for and assets are disaggregated.

Figures 1 and 2 show the long-run contribution to the log consumption-to-income ratio of the three asset-to-income ratios, the general credit conditions index, and forecast income growth (the empirical counterpart of equation (2.8)), weighting each by its estimated long-run coefficient. These are not general equilibrium effects, but are nevertheless useful in understanding the estimation results. 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 help account for relative rises in consumption from 1984 to 1989 and from 1995 to 2005. The role of income growth expectations (see Figure 2) appears far smaller, having a negative effect from 1984 to 1989 and a small positive role from 1995 to 2005.

Figure 2 further suggests that the upward trend in the value of illiquid wealth holdings relative to income helps explain the similar trend in consumption relative to income. However, rising debt, reflected in the fall of the net liquid assets-to-income ratio, 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 of other assets is critical. Conventional discussions of wealth effects tend to focus on net worth and so miss the special role of liquidity and debt. U.K. consumption levels are quite vulnerable to downturns in asset prices, given high levels of debt and the difficulty of reducing debt in the short-run.

[14]This analysis treats current income growth, the forecast of future growth and the unemployment rate as weakly exogenous variables. Evidence for weak exogeneity is found in models for these I(0) variables in which the lagged equilibrium correction term implied by the co-integration vector is insignificant. While income is likely to be endogenous for consumption, on the U.K. data, current quarter growth of real income appears to be weakly exogenous for the log consumption to income *ratio*.

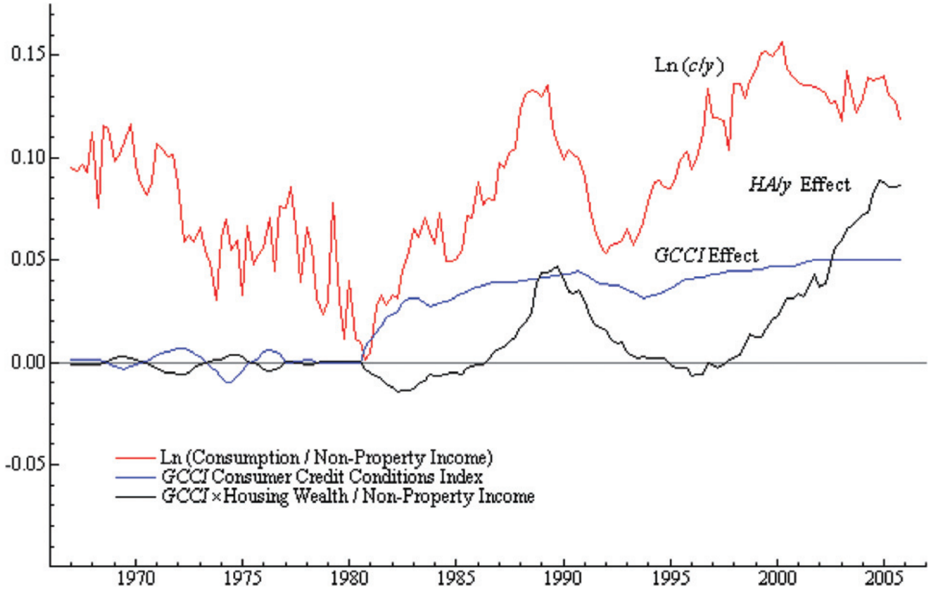

Figure 1. Estimated Long-Run Contributions to Log Consumption-to-Income Ratio of the General Consumer Credit Conditions Index (GCCI) and its Interaction with the Housing Wealth-to-Income Ratio in the U.K.

Note: The U.K. 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. Footnote 13 gives the definition of the interaction between GCCI and the housing wealth 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 U.K. series of Fernandez-Corugedo and Muellbauer (2006), the closest U.S. data source for tracking mortgage LTV and loan-to-income ratios is the American Housing Survey. However, the sample is far smaller than in the U.K. survey of mortgage lenders and the LTV data are only usable from 1979. As neither ratio rose much from 1979 to 1998, it does suggest that the easing of mortgage credit conditions for U.S. first-time home-buyers may have been less dramatic in this period than for the U.K. (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. *Unsecured Consumer Credit Index for the U.S.*

One data advantage the U.S. has over the U.K. is the Federal Reserve's long running quarterly Senior Loan Officer Opinion Survey. Using this survey, Duca *et al.* (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 installment loans. This index is

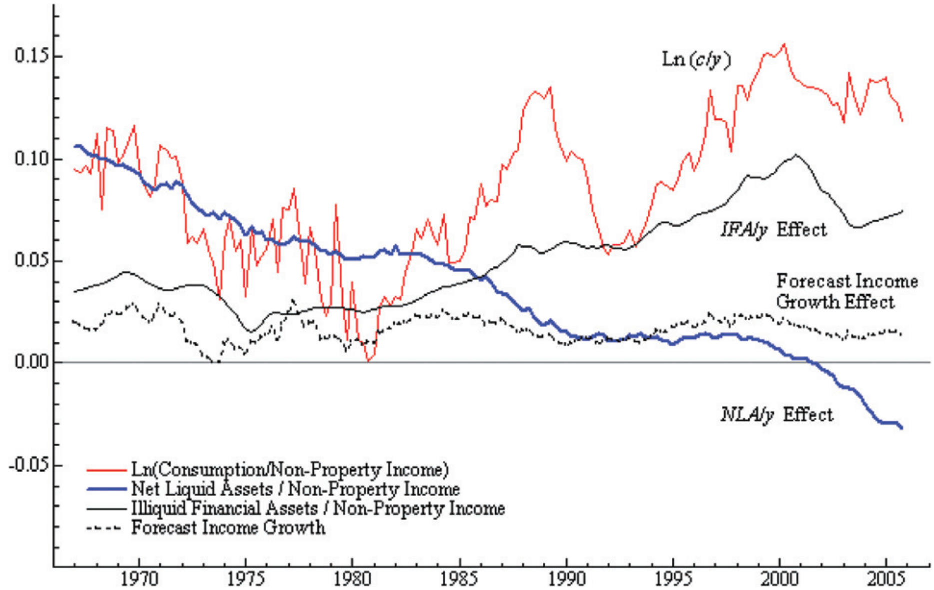

Figure 2. Estimated Long-Run Contributions to Log Consumption-to-Income Ratio of Net Liquid Assets (NLA) and Illiquid Financial Assets (IFA) to Income Ratios and Forecast Income Growth in the U.K.

negatively and strongly 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 the change index, *CR* was adjusted for the effects of changes in interest rates and in the macroeconomic outlook using a regression based on screening models. The adjusted *CR* index was then chained into a levels index, based on the correlation of the *CR* index with the growth rate of real consumer loan extensions at banks. The resulting *CCI* rises greatly during the 1980s, rises again during the height of the subprime mortgage boom in 2004–06, before reversing these gains since 2006 (see Figure 3).

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 U.K. The consumption measure includes durables and imputed housing services, the same as in the U.K. Similar results were obtained using consumption excluding housing services (Duca *et al.*, 2011c). The estimation 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, 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

TABLE 2

U.S. CONSUMPTION FUNCTION ESTIMATES FOR 1973 Q1 TO 2010 Q3

Dependent Variable = $\Delta \ln c_t$	Symbol	(1)	(2)	(3)	(4)	(5)
Speed of adjustment	λ	0.093***	0.091***	0.075***	0.292***	0.292***
		(4.3)	(3.7)	(3.5)	(7.7)	(7.7)
Long-run coefficients						
Constant	α_0	0.137	0.037	0.101	0.042	0.042
		(1.3)	(0.3)	(0.9)	(1.6)	(1.6)
Unsecured consumer credit index, *CCI*	α_{01}	–	–	–	0.145***	0.146***
					(12.7)	(15.2)
Real interest rate$_{t-1}$	α_1	–0.0069**	–0.0080**	–0.0068*	–0.0035***	–0.0035***
		(2.1)	(2.4)	(1.8)	(3.3)	(3.3)
Forecast future income growth	α_3	0.961***	0.948***	0.710**	0.588***	0.588***
		(3.3)	(2.9)	(2.1)	(5.7)	(5.7)
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**	0.086**	0.153***	0.153***
			(3.0)	(2.3)	(11.0)	(12.0)
Illiquid financial assets$_{t-1}$/income	γ_2	–	0.051***	0.049***	0.011***	0.011***
			(4.0)	(3.9)	(3.8)	(3.8)
Housing wealth$_{t-1}$/income	γ_3	–	0.069***	0.044**	0.0001	–
			(3.5)	(2.0)	(0.0)	
Housing wealth$_{t-1}$/income × housing liquidity index *HLI*	γ_{31}	–	–	–	0.084***	0.084***
					(8.8)	(10.9)
Short-run coefficients						
Income growth	β_1	0.208***	0.209***	0.150***	0.068	0.068
		(3.5)	(3.5)	(3.0)	(1.6)	(1.6)
Change in nominal interest rate	β_2	–	–	–0.0047***	–0.0036***	–0.0036***
				(6.1)	(5.5)	(5.5)
Change in unemployment rate	β_3	–	–	–0.0062***	–0.0036***	–0.0036***
				(5.5)	(3.7)	(3.7)
Dummies						
Oil shocks dummy	β_4	–0.012***	–0.012***	–0.012***	–0.0084***	–0.0084***
		(4.5)	(4.6)	(5.3)	(4.6)	(4.6)
1974 proposed tax increases dummy	β_5	–0.023***	–0.022***	–0.015***	–0.016***	–0.016***
		(4.5)	(4.4)	(3.4)	(4.4)	(4.5)
1978 coal strike dummy	β_6	–0.0070**	–0.0070**	–0.014**	–0.015***	–0.015***
		(2.0)	(2.1)	(3.0)	(4.0)	(4.0)
1980 Carter credit controls dummy	β_7	–0.028***	–0.028***	–0.0065**	0.0071***	0.0071***
		(5.5)	(5.5)	(2.2)	(2.9)	(3.0)
1987 Tax Reform Act dummy	β_8	–0.0082**	–0.0081**	–0.0074**	–0.0067***	–0.0067***
		(2.3)	(2.3)	(2.6)	(2.8)	(2.8)
Diagnostics						
Standard error × 100		0.50	0.48	0.40	0.33	0.33
Adjusted R^2		0.45	0.47	0.63	0.75	0.75
Durbin Watson		1.62	1.38	1.85	2.21	2.21
AR1/MA1 (p-value)		0.02	0.07	0.38	0.10	0.10
AR4/MA4 (p-value)		0.00	0.00	0.19	0.08	0.08
Heteroscedasticity (p-value)		0.20	0.24	0.68	0.82	0.82
Chow (1985 Q1 break, p-value)		0.00	0.00	0.12	0.91	0.88
RESET (p-value)		0.12	0.10	0.64	0.82	0.82

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:

$$\begin{aligned}\Delta \ln c_t = {} & \lambda\Big\{\alpha_0 + \alpha_{01}CCI_t + (\ln y_t - \ln c_{t-1}) + \alpha_1 r_{t-1} + \alpha_3\Big(\widehat{\ln y^p_{t+1}} - \ln y_t\Big) + \gamma_1 NLA_{t-1}/y_t + \gamma_2 IFA_{t-1}/y_t \\ & + \gamma_3 HA_{t-1}/y_t + \gamma_{31} HLI_t \times HA_{t-1}/y_t\Big\} + \beta_1 \Delta \ln y_t + \beta_2 \Delta nr_t + \beta_3 \Delta ur_t + \beta_4 Oil\ Shock_t \\ & + \beta_5 d1974q4_t + \beta_6 Coal\ Strike_t + \beta_7 d1980q2_t + \beta_8 Tax\ Reform\ Act_t + u_t\end{aligned}$$

Middle East disturbances, a proposed but later aborted tax hike in 1974 Q4, a major coal strike in the first half of 1978, as well as the temporary effect of tax shifting surrounding the major 1987 Tax Reform Act.[15]

Permanent income is forecast using a simple model based on reversion to a split trend (with a slow-down in growth from 1968 on and a small pickup in 1988 which reverses in 1999) and 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 a Thomson Reuters/University of Michigan survey measure of consumer expectations.

In the basic model in Column 1, the estimated long-run MPC for net worth is 4 percent (0.04), near conventional estimates in the literature. 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 to the key coefficients in Column 1. The estimated net liquid assets MPC jumps to over 12 percent. The estimated MPCs of illiquid financial assets and housing wealth are approximately 5 and 7 percent, respectively.

In Column 3, the level of the real interest rate and changes in the unemployment rate and the nominal interest rate are added. The interest rate is the auto-finance rate adjusted for depreciation, which reflects special offers sometimes available to borrowers. The changes in the nominal interest rate and the unemployment rate are both highly significant, as in the U.K. Surprisingly, the estimated net liquid asset, illiquid financial asset, and housing wealth MPCs are fairly similar. This is implausible since liquid assets such as cash should be more spendable than illiquid financial assets and housing assets (even if used as collateral for a loan rather than sold outright, these can entail large transactions costs).

A plausible explanation is that the ability to use housing wealth as collateral has varied over time. Ignoring the rise in the liquidity of housing wealth imparts a downward bias to the estimated 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 1980s. In part, this decline reflects easier mortgage credit conditions, which tend to boost consumption.[16] A simple check on the model in Column 3 was performed by adding 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 unsecured consumer credit conditions index, *CCI*, and interacting housing wealth with an index of housing liquidity, *HLI*. Duca *et al.* (2011c) estimate *HLI* as a common latent factor or spline in a three-equation model of U.S. consumption, mortgage

[15]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.

[16]The costs of refinancing fixed interest rate U.S. mortgages fell in the 1990s, as shown by Bennett *et al.* (2001) and discussed by Green and Wachter (2007).

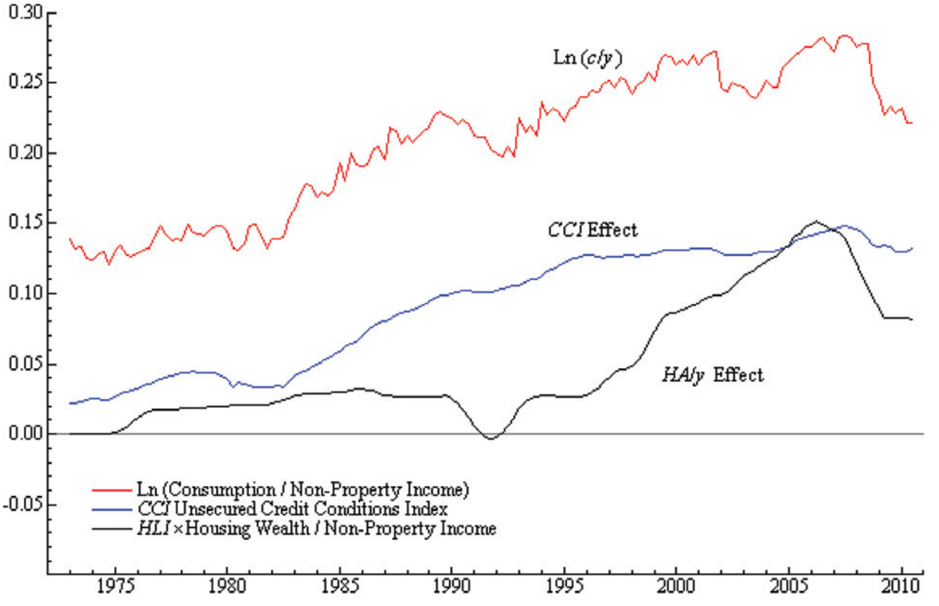

Figure 3. Estimated Long-Run Contributions to Log Consumption-to-Income Ratio of the Unsecured Credit Conditions Index (*CCI*) and the Housing Wealth-to-Income Ratio Adjusted for Housing Liquidity in the U.S.

equity withdrawal, and mortgage refinancing.[17] Similar estimates of *HLI* were obtained from a non-linear, two equation state space model for consumption and mortgage refinancing. It turns out that both the unsecured credit conditions index, *CCI*, and the housing liquidity–housing wealth interaction, $HLI \times HA/y$, are highly significant. Their inclusion in Column 4 lowers the equation standard error by nearly 20 percent and raises the speed of adjustment from 0.10 to 0.29, implying a better specified long-run solution.[18] The significance of housing wealth interacted with *HLI* and the insignificance of the housing wealth on its own supports the collateral view of housing wealth in the consumption function. This result accords both with theory and the micro evidence that observed housing wealth effects are really housing collateral effects (Browning *et al*., 2009).

Finally, Column 5 reports the same regression, but without the insignificant level of housing wealth and current income growth terms in Column 4. The estimated long-run contribution to the log consumption-to-income ratio of the three asset-to-income ratios, the unsecured credit conditions index, *CCI*, and forecast income growth are shown in Figures 3 and 4. The much higher MPC of net liquid assets (15 percent) than from illiquid financial assets (1 percent) is

[17]The estimated *HLI* is zero before 1975. It rises to about 0.13 in 1976. It rises again to about 0.19 between 1983 and 1986 when it reverts to its previous value. In the early 1990s it plunges to zero before recovering in the mid 1990s and then climbing steeply to about 0.5 in late 1999. It rises again to about 0.55 between 2003 and 2007 before falling back to about 0.47 at the end of sample.

[18]In other runs, current income was instrumented using a simple forecasting model and very similar results were obtained, with the *CCI* and *HLI* housing wealth interaction terms still highly significant. This suggests that the endogeneity of current income is not a major issue in these regressions.

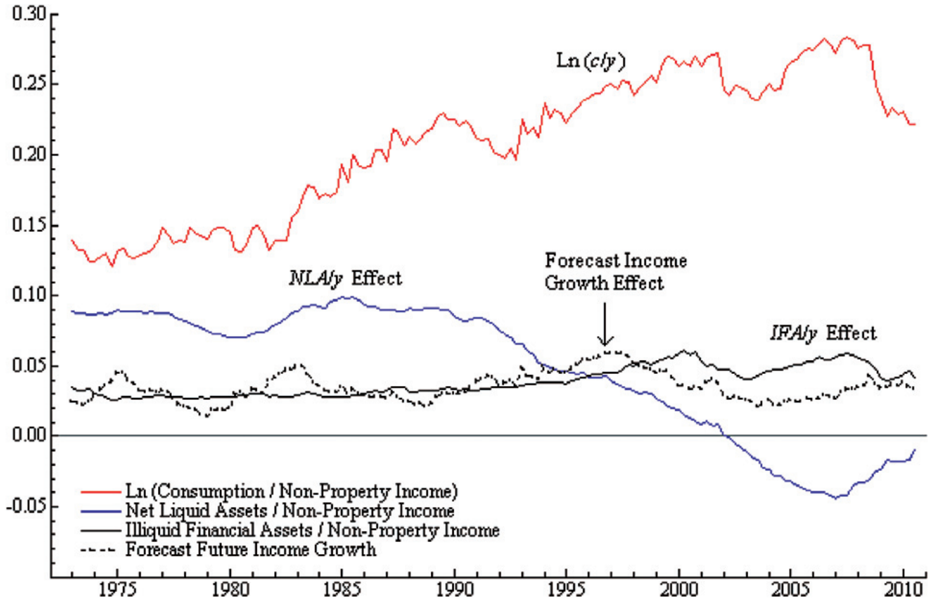

Figure 4. Estimated Long-Run Contributions to Log Consumption-to-Income Ratio of the Net Liquid Assets (NLA) and Illiquid Financial Assets (IFA) to Income Ratios and Forecast Income Growth in the U.S.

broadly consistent with microeconomic evidence in Gross and Souleles (2002). Likewise, the higher MPC of housing wealth (given more liberal mortgage credit) compared to that of illiquid financial wealth supports the findings of Benjamin *et al*. (2004), Case *et al*. (2005), and Carroll *et al*. (2011).

The similarities in the estimated wealth MPCs for the U.K. and U.S. are noteworthy (Table 3). The MPC of net liquid assets is estimated to be 11 percent in the U.K. and 15 percent in the U.S. The estimated MPC of illiquid assets is 2 percent in the U.K. and 1 percent in the U.S. The peak MPC of housing wealth is estimated to be between 4 and 5 percent in both countries, despite several structural differences in the two housing markets.[19] Muellbauer and Williams (2011) obtain very similar estimated wealth MPCs for Australia, 0.159 for net liquid assets, 0.022 for illiquid financial assets, and a peak value of 0.049 for housing assets.

Other researchers obtain much higher housing wealth MPCs for the U.S. For example, 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 to ours. His evidence suggests that institutional differences between countries have large effects on the MPC of housing wealth, in particular

[19]On the one hand, transactions fees (estate agents fees and taxes etc.) are much larger in the U.S. than in the U.K., implying greater housing liquidity in the U.K. 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), many homeowners are not subject to deficiency judgments, so lenders have recourse only to the house collateralizing the mortgage, which may make housing more attractive to borrow against.

TABLE 3

Estimated Marginal Propensities to Consume for Various Components of Net Worth in the U.K. and the U.S.

	Net Liquid Assets MPC	Illiquid Financial Assets MPC	Maximum Housing Wealth MPC
U.K.	0.114	0.022	0.043
U.S.	0.153	0.011	0.047

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

that it is larger in countries with more liberal mortgage markets. This concurs with our evidence for the U.K., U.S., and Japan (see below). His evidence is also consistent with an upward drift over time in the housing "wealth" MPC, linked with greater credit market liberalization. However, Slacalek's estimates of the housing "wealth" MPC in both the U.K. and the U.S. 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

We estimate variants of equation (2.10) for Japan using annual data from 1961 to 2008. Consumption includes durables expenditure as well as imputed housing, the same as in the U.K. and U.S. Permanent income is constructed using a five year horizon and a 20 percent per annum discount rate, equivalent to the 5 percent per quarter rate used for the U.K. and U.S. Our permanent income forecasting model includes a trend, split trends from 1973 and 1991—reflecting a large slowdown in Japanese growth from 1973 and a smaller slowdown in 1991—and 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; as well as the two-year change in the nominal interest rate (which has a negative effect). The parameters of the income forecasting equation are stable when estimated over alternative samples.

The estimates of the Japanese consumption function are set out in Table 4. In the basic REPIH model in Column 1, only current income growth is significant. *Inter alia*, the real interest rate, forecast future income growth (defined by equation (2.8)), the net wealth to income ratio, and the speed of adjustment, 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. A demographic variable, 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, the disaggregation of net wealth produces a sharp improvement in fit, though physical assets including housing are never significant. However, real land prices are significant at the 10% level and negatively signed.

TABLE 4
Japanese Consumption Function Estimates for 1961 to 2008

Dependent Variable = $\Delta \ln c_t$	Symbol	(1)	(2)	(3)	(4)
Speed of adjustment	λ	0.056	0.105***	0.489***	0.461***
		(1.4)	(3.0)	(6.3)	(6.8)
Long-run coefficients					
Constant	α_0	0.318	0.188	–0.169***	–0.165***
		(0.7)	(1.3)	(6.9)	(15.9)
Real interest rate	α_1	0.059	0.027***	0.0073***	0.0083***
		(1.4)	(2.3)	(4.8)	(6.3)
Forecast future income growth	α_3	1.025	0.770**	0.471***	0.460***
		(1.2)	(2.3)	(7.6)	(7.7)
Forecast income growth × income growth volatility	α_{32}	–	–	–	–3.688**
					(2.2)
Log real land prices	α_4	–	–	–0.024*	–0.021**
				(1.7)	(2.2)
Net worth$_{t-1}$/income	$\gamma_1 = \gamma_2 = \gamma_3$	–0.033	–0.012	–	–
		(0.6)	(0.7)		
Net financial assets excluding shares$_{t-1}$/income	γ_1	–	–	0.064***	0.063***
				(11.9)	(19.0)
Shares$_{t-1}$/income	γ_2	–	–	0.039	
				(0.7)	
Physical assets$_{t-1}$/income	γ_3	–	–	0.0034	–
				(0.6)	
Short-run coefficients					
Income growth	β_1	0.547***	0.448***	0.236***	0.272***
		(6.6)	(6.1)	(3.6)	(4.4)
Change in unemployment rate	β_3	–	–0.017***	–0.015***	–0.013***
			(3.3)	(3.1)	(3.4)
Income volatility	β_4	–	–0.310***	–0.169**	–
			(2.8)	(2.0)	
Acceleration in ratio of population aged under 20 to working age population	β_5	–	0.871**	0.779**	0.696**
			(2.2)	(2.5)	(2.4)
Diagnostics					
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 are shown in parentheses. Statistical significance at the 10%, 5%, and 1% levels is denoted by *, **, and *** respectively. The general model is:

$$\begin{aligned}\Delta \ln c_t = \lambda \Big\{ & \alpha_0 + (\ln y_t - \ln c_{t-1}) + \alpha_1 r_t + \alpha_3 \left(\widehat{\ln y^p_{t+1}} - \ln y_t\right) + \alpha_{32} \left(\widehat{\ln y^p_{t+1}} - \ln y_t\right) \times \sigma_{\Delta \ln y, t} + \alpha_4 \ln pland_{t-1} \\ & + \gamma_1 (NFA_{t-1} - SHARES_{t-1})/y_t + \gamma_2 SHARES_{t-1}/y_t + \gamma_3 PHY_{t-1}/y_t \Big\} + \beta_1 \Delta \ln y_t + \beta_3 \Delta ur_t + \beta_4 \sigma_{\Delta \ln y, t} \\ & + \beta_5 \Delta^2 (POP_{<20,t}/POP_{20-65,t}) + u_t\end{aligned}$$

The real interest rate, speed of adjustment, 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 in the U.K. and U.S. analysis. Notably, the positive real interest rate effect is not a disguised inflation effect as the inflation rate is insignificant when added, while the real rate remains significant.

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

There is no evidence of a shift in the consumption function due to financial liberalization, in contrast to the results for the U.K. and U.S. This equation is stable when estimated over samples ranging from 1961 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 (proxied by land prices) on Japanese consumption.

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. Lower income growth and the uncertainty indicators explain some of the dramatic decline in the consumption-to-income ratio in the 1970s. Interestingly, net financial assets relative to income shows little cyclical variation, as the pension fund component is not very sensitive to the stock market, unlike in the U.K. The decline in net financial assets in the early 1990s did, however, contribute to the drop in the consumption-to-income ratio at that time.

6. Concluding Comments

Consistent with theory, our empirical findings for the U.K., U.S., and Japan demonstrate the importance of credit constraints for consumer spending. The evolution of credit availability differs over time within countries, as well as between them. The large changes in the availability of credit to U.K. and U.S. households in recent decades have shifted the consumption function in both countries. Financial liberalization has enhanced the positive impact of housing wealth on consumption in the U.K. and U.S. and also the role of expected income growth on consumption in the U.K. By contrast, the Japanese consumption function has been stable since the 1970s, reflecting the lack of household credit liberalization. This fact, together with differences in the tax code which favor inheritance via

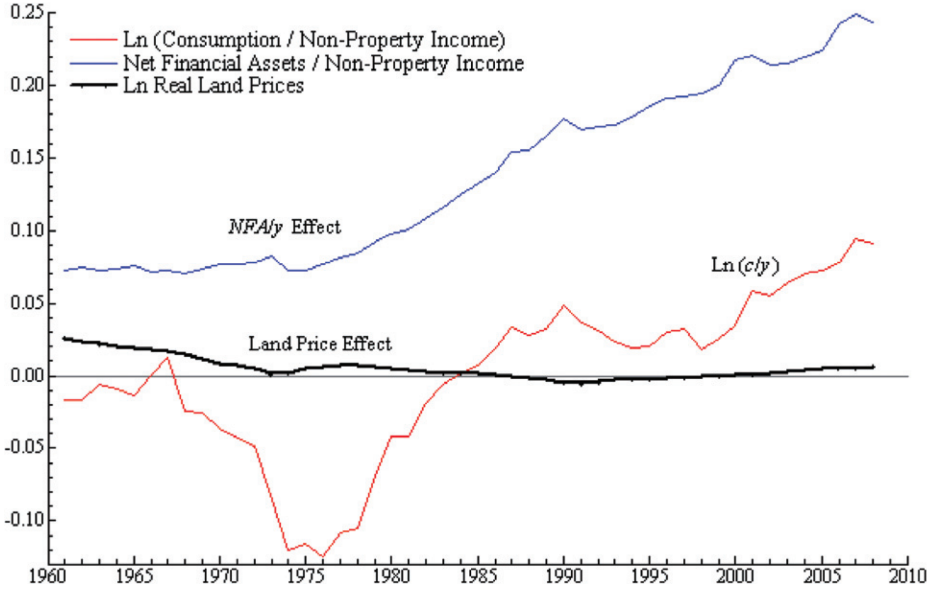

Figure 5. Estimated Long-Run Contribution to Log Consumption-to-Income Ratio of the Net Financial Assets (*NFA*) to Income Ratio and Log Real Land Prices in Japan

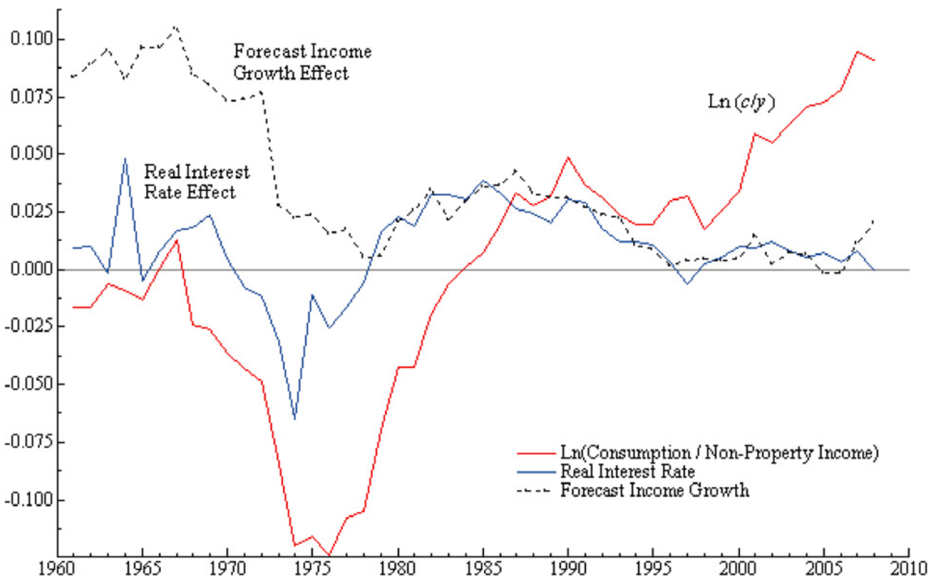

Figure 6. Estimated Long-Run Contribution to Log Consumption-to-Income Ratio of Real Interest Rate and Forecast Income Growth in Japan

housing, likely accounts for the restrained response of consumption in Japan to rising home prices.

These differences in the consumption functions suggest that the transmission of monetary policy via the household sector is far less powerful in Japan than in the U.K. or U.S. Large household liquid asset holdings relative to debt imply that households in Japan, particularly older households, feel poorer when short-term interest rates fall and so reduce spending. In the U.K. and U.S., 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 in Japan, as renters are likely to save more in anticipation of higher future rents or of higher mortgage down-payments. In the U.K. and U.S., in contrast, greater possibilities for housing equity withdrawal combine 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 large declines in wealth between 2007 and 2009, particularly in housing equity, will have strong and persistent dampening effects on consumer spending in the U.K. 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.*, 2010). In both episodes, mortgage availability declined sharply. In the U.K., tighter credit standards since 2007 sharply reduced loan-to-value and loan-to-income ratios for first-time homebuyers, contributing to substantial 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. Nevertheless, the global recession, particularly in the U.S., has proved detrimental to Japanese household income, due to declining 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 U.K. and U.S., and their contrast with Japan, reflects the importance of institutional differences between countries. This underlines the contribution of a modernized Ando–Modigliani–Brumberg consumption function, which incorporates credit frictions, uncertainty, and income expectations. Household balance sheets, though 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 can neither be properly modeled nor well understood.

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