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## **How Financial Innovations and Accelerators Drive Booms and Busts in U.S. Consumption\***

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*Abstract:* The recent US consumption boom, bust and recovery have been attributed to fluctuations in financial and housing wealth but these do not explain the longer term record. Coherent explanations highlight changes in the supply of credit and the liquidity of housing wealth, factors which are not directly observed. Our indexes of unsecured consumer credit availability and the liquidity of housing wealth address this gap, and yield a consumption function with far superior parameter stability and ability to account for U.S. consumption than models that omit these factors. The liquidity of housing wealth is estimated as a common unobservable state in a jointly estimated, non-linear state space model of consumption and mortgage refinancing. The resulting credit-augmented, life cycle model of consumption shows that financial innovations and frictions play critical roles in the booms and busts in U.S. consumption.

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## 1. Introduction

The record decline in the ratio of U.S. consumption to income from 2007 to 2010 and the preceding rise and subsequent recovery have been attributed to fluctuations in illiquid financial wealth (mainly stocks) and housing wealth. However, as Figures 1 and 2 suggest, the rise in the consumption-to-income ratio in the second half of the 1970s and in the 1980s is not easily explained by these factors. There is controversy over the nature and size of housing wealth or housing collateral effects on consumption (Cooper and Dynan, 2014). Some studies find that the sensitivity of U.S. consumption to housing wealth rose between the mid-1990s and mid-2000s (Carroll *et al.*, 2011; Case and Shiller, 2008; while Slacalek, 2009 documents similar cases in other countries).

Basic life-cycle consumption theory implies that, since housing is a consumption good as well as an asset, there cannot be a large aggregate housing wealth effect on consumption. Large estimated housing wealth effects likely reflect an increase in the liquidity of housing wealth (Carroll and Kimball, 2005; Aron *et al.*, 2012) that enables otherwise credit constrained households to borrow against housing equity. Consistent with this collateral view of “housing wealth effects,” several cross-section studies have found that consumption is much more sensitive to housing wealth among those most apt to be credit constrained (Browning, et al., 2013; Disney and Gathergood, 2011; Hurst and Stafford, 2004; Mian and Sufi, 2011a, b, and Windsor, et al, 2015). Increased access to mortgage credit will then increase the liquidity of housing wealth, consistent with evidence that a greater ability of families to tap housing wealth via mortgage-equity withdrawals (MEW) contributed to the early 1990s boom and bust in UK consumption (Muellbauer and Murphy, 1990; Miles, 1992). Macroeconomic forecasters have found MEW series useful in gauging U.S. consumer spending (e.g., Duca, 2006; Greenspan and

Kennedy, 2008). However, the main reason may be that MEW tracks the evolution of credit availability and the sensitivity of consumption to housing collateral. Exogenous, hard to predict, changes in credit supply and the liquidity of housing wealth can make MEW, and thereby consumption, prone to large shifts. Indeed, prior to the 1986 tax changes which made home equity loan interest tax deductible, the ratio of MEW to income barely changed in response to large swings in real home price appreciation (Figure 3). But after 1986, the two series become more positively correlated. The advent of new mortgage products, especially those increasing access to and lowering the costs of refinancing mortgages, enabled households to withdraw more housing equity via cash-out mortgage refinancing. Through this channel, MEW became more sensitive to housing wealth, having implications for cash flow and leverage effects on consumer spending.

We explore these effects using an econometric models based on solved-out consumption specifications, which allow for short- and long-run effects of credit availability and wealth. In this framework, the non-income impact of the crisis on consumption can be gauged through two channels: the availability of unsecured consumer credit and the availability of mortgages for accessing housing collateral. The former helps explain the rise of the consumption-to-income ratio in the 1970s and 1980s. Innovations in mortgage and other related products may increase the liquidity of housing wealth, thereby raising the m.p.c. (marginal propensity to consume) of housing wealth and helping to account for the rise of the consumption-to-income ratio between the late 1990s and mid-2000s. We track the first by an index of consumer credit availability derived from the Federal Reserve's Senior Loan Officer Survey, that is adjusted for cyclical and interest rate effects improving upon Muellbauer (2007) and Duca and Garrett (1995).

To model the second channel, we track the evolution of major wealth components and the liquidity of housing wealth, estimated as a common unobservable factor or state, in a jointly estimated, non-linear state space model of consumption and mortgage refinancing. *Inter alia*, the consumption equation accounts for income, expectations, and different types of wealth, allowing the impact of housing wealth to depend on the latent housing liquidity index (*HLI*). *HLI* only enters the consumption function interacted with housing wealth, so it may be interpreted as the evolving m.p.c. of housing wealth. *HLI* enters the refinancing equation both as its own and interacted with interest rate variables that measure the incentive to refinance. Duca and Muellbauer (2014) give the name Latent Interactive Variables Equation System (LIVES) to such a system where the latent variable interacts with observed variables.

*HLI* is inversely related to the unobserved pecuniary and other costs of refinancing, which have fallen over time, lowering the barriers to and costs of withdrawing housing equity via cash-out mortgage refinancing (e.g. Bennett, *et al.*, 2001). MEW activity reflects a combination of mortgage refinancing, home equity borrowing, and the roll-over of capital gains when homeowners change homes. Greenspan and Kennedy (2008) show that the principal component of active MEW is cash-out refinancing, so the two series move closely together. Our paper estimates the liquidity of housing wealth (*HLI*) from a two-equation model of consumption and mortgage refinancing, rather than a three-equation model including MEW, because the determinants of refinancing are more directly observable and MEW is harder to measure than refinancing.

Our credit augmented consumption function allows for innovations in consumer credit and the liquidity of housing wealth. As a result this model has a better fit, more stable coefficients, and more plausible short and long-run properties than other consumption functions.

We find that financial innovations have altered the housing collateral (wealth) and unsecured consumer credit channels. In the boom, consumption was boosted by easier consumer credit standards and by an increased liquidity of housing wealth. In the bust these developments partially unwound, which combined with the large falls in house and equity prices, induced a sharp drop in consumption-to-income ratio and uptick in the personal saving rate.

The next section outlines our credit augmented consumption function. Section 3 describes our measure of unsecured consumer credit availability. Section 4 presents the refinancing equation and shows how the index of housing liquidity (*HLI*) is estimated in our two equation model of consumption and mortgage refinancing. The refinancing results and our estimated *HLI* are discussed in Section 5. Section 6 reviews our consumption function results, while Section 7 and then assesses the gains from estimating housing liquidity from refinancing activity, along with implications for how shifts in wealth, credit availability, and the liquidity of housing wealth contributed to the recent uptick in the U.S. personal saving rate that followed a long downtrend over 1980-2007. Section 8 concludes.

## **2. Credit Constraints, Housing Liquidity and Consumption – The Linkages**

This section reviews the implications of the evolution of consumer credit availability and housing wealth liquidity in a solved-out consumption function. The Euler equation approach has the attraction of simply specifying consumption with first difference terms that do not appear, to require tracking structural factors, but, in fact, omits important long-run relationships involving wealth and credit frictions. As Campbell and Mankiw (1989) and Muellbauer (2010) *inter alia* show, empirical aggregate Euler equations violate the martingale condition implied by simple theory. In contrast, our modernized Ando-Modigliani style consumption specification

encompasses the rational expectations permanent income hypothesis but incorporates wealth and credit channels, passes a number of diagnostic tests and yields sensible coefficients and results.

## 2.1. A Consumption Function with Wealth and Credit Channels

The perfect capital markets version of the basic life-cycle and permanent income hypotheses (LCH/PIH) implies that real per capita consumption  $c$  is given by:

$$c_t = \phi A_{t-1} + \omega y_t^p \quad (2.1)$$

where  $y^p$  is permanent real non-property income and  $A$  is the real net wealth, both in per capita terms. Letting  $y$  be current real income and using the approximation  $(y^p - y)/y \approx \ln(y^p/y)$  and some algebra yields:

$$\ln c_t = \alpha_0 + \ln y_t + \gamma A_{t-1}/y_t + \ln(y_t^p/y_t) \quad (2.2)$$

where  $\gamma = \phi/\omega$  and  $\alpha_0 = \ln \omega$ . The log difference between permanent and actual income reduces to a discount-weighted moving average of forward income growth rates (Campbell, 1997):

$$E_t \ln(y_t^p/y_t) \approx E_t \left( \sum_{s=1}^K \eta^{s-1} \ln y_{t+s} \right) / \left( \sum_{s=1}^K \eta^{s-1} \right) - \ln y_t \equiv E_t \ln yperm_t - \ln y_t \quad (2.3)$$

where  $K$  is the horizon and  $\eta$  is a discount factor.

As Deaton (1991) showed, the consequence of income uncertainty and of liquidity constraints, especially the possibility that some households may not be able to obtain credit when they need to borrow, invalidates this text-book model. An implication of the research by Deaton and related papers by Carroll, e.g., Carroll (2001), is that households are heterogeneous with respect to the discount rates and horizons that they apply to expected but uncertain future income. Heterogeneity across households does not, however, imply that aggregate consumption cannot be modeled.<sup>1</sup> *At the very least*, an aggregate model needs to relax the assumption of a

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<sup>1</sup> This is just one reason why there is no ‘representative consumer’ following rational calculus through whom one can explain the consumption decisions of an entire economy. Deaton (1992) marshals theory and evidence in favour

coefficient of unity on  $\ln y^p/y$  in equation 2.2 and accept that, on average, the discount factor  $\eta$  is likely to be substantially below  $1/(1+r)$ , where  $r$  is a real risk-free interest rate.

Making these changes and rearranging yields an expression for a modified REPIH model:

$$\ln c_t = \alpha_0 + \ln y_t - \alpha_1 r_t - \alpha_2 \theta_t + \alpha_3 (E_t \ln y^{perm}_t - \ln y_t) + \gamma A_{t-1}/Y_t + \varepsilon_t \quad (2.4)$$

More realism can be added by adjusting for habits (or rational inattention) and uncertainty. The wealth-to-income ratio can be disaggregated into ratios to income for liquid assets less debt ( $NLA/Y$ ), illiquid financial assets ( $IFA/Y$ ), and gross housing assets ( $HSG/Y$ ). If structural factors raise the liquidity of housing wealth ( $HLI$ ), this could bolster consumption in several ways.  $HLI$  could enter as an intercept to track a higher average propensity to consume out of income, or enter interacted with permanent income growth (as an enhanced collateral role for housing allows more borrowing in anticipation of future income), or  $HLI$  could enter interacted with  $HSG/Y$ , reflecting a larger housing collateral effect. Finally, a consumer unsecured (non-mortgage) credit conditions index ( $CCI$ ) may also affect consumption.

All of this implies the following equilibrium-correction model for consumption:

$$\begin{aligned} \Delta \ln c_t \sim \lambda (\alpha_0 + (\ln y_{t-1} - \ln c_{t-1}) + \alpha_1 r_{t-1} + \alpha_2 (\ln y^{perm} - \ln y) + \alpha_3 CCI_{t-1} + \gamma_1 (NLA/Y)_{t-1} \\ + \gamma_2 (IFA/Y)_{t-1} + \gamma_3 (HSG/Y)_{t-1} \times HLI_{t-1} + \varepsilon_t \end{aligned} \quad (2.5)$$

where the term in brackets is equilibrium minus actual consumption,  $\lambda$  is the speed of adjustment toward long-run equilibrium and the  $\gamma$ 's are the m.p.c.'s of the wealth components.

The m.p.c of housing wealth varies with the liquidity of housing wealth.

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of incorporating liquidity constraints and precautionary saving in the face of uncertainty to understand consumption. He also notes that consumers' balance sheets include illiquid assets such as pensions, stocks and bonds, and houses: "the presence of these illiquid and sometimes high-yielding assets needs to be integrated into the model of credit-constrained consumers." This has implications not only for reformulating the consumption function, but also for jointly modelling the behaviour of consumption and household portfolios.



The m.p.c.'s should differ by asset type. The m.p.c. out of net liquid assets should be higher than out of illiquid financial assets or housing wealth, since cash-like assets are more spendable and borrowers face penalties for not meeting debt obligations (see Mian and Sufi, 2011a,b; Mishkin, 1976, 1978; and Muellbauer and Lattimore, 1995). There are good theoretical reasons for why the m.p.c.'s for illiquid financial assets and housing assets should differ. Most importantly, housing gives direct utility in the form of services implying that there are income and substitution effects not present for financial assets. The  $\gamma_3$  coefficient reflects how the evolution of housing wealth liquidity alters the m.p.c. of housing collateral or wealth.<sup>2</sup> Consumption is tracked by total real consumption expenditures and income is measured by non-property (labor plus transfer) income, which omits dividends and interest earned on wealth that are embodied in asset prices. As Blinder and Deaton (1985) show, temporary tax changes induce larger deviations in income than in consumption, reflecting the small impact of temporary taxes on permanent income. Similarly, we adjust non-property income for temporary tax changes using BEA estimates of their impact on disposable income.<sup>3</sup> We track income uncertainty using the four-quarter contemporaneous change in the unemployment rate ( $\Delta 4ur$ ). For expectations of the deviation of permanent from current income, we use a simple model based on reversion to a split trend (with a slow-down in growth after 1968) with two drivers (see Appendix A and Figure 4). These are the 4-quarter change in the 3-month Treasury bill yield to track monetary policy and the Michigan index of consumer expectations of future economic conditions. Permanent income was constructed with three alternative quarterly discount rates, 0.025, 0.05 and 0.1. As

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<sup>2</sup> Down-payment constraints were relaxed in the mid-2000s (Duca, Muellbauer and Murphy, 2011a, b, 2016).

<sup>3</sup> These include the tax surcharges during the Vietnam War, temporary tax cuts in 1975, 2001, 2005, and 2008; but not Blinder and Deaton's estimates for the phase-in of the tax cuts of the early 1980s; details available upon request.

there is little difference in fit between the last two, a discount rate of 0.05 ( $\eta = 0.95$ ) was chosen.<sup>4</sup>

The real interest rate ( $r$ ) is the Federal Reserve Board's user cost of capital for autos (the real interest rate on finance company auto loans plus auto depreciation). To track short run credit effects such as large inter-temporal shifts in auto sales induced by changes in auto interest rate incentives, we include the change in the real interest rate ( $\Delta r$ ). Using Flow of Funds data, liquid assets ( $NLA$ ) are the sum of deposits and credit market instrument minus consumer ( $CDEBT$ ) and mortgage ( $MDEBT$ ) debt. Housing assets ( $HA$ ) are gross housing assets, while illiquid financial assets ( $IFA$ ) equal all other household assets. The last two variables are the credit conditions ( $CCI$ ) and housing liquidity ( $HLI$ ) indexes, discussed below.

### 3. The Unsecured Consumer Credit Conditions Index ( $CCI$ )

We construct a levels index of unsecured consumer credit conditions ( $CCI$ ) index using data from the Federal Reserve's Senior Loan Officer Opinion Survey of 60 large banks which report on how their willingness to make consumer installment loans has changed relative to three months prior. This index, which is used in Aron, *et al.* (2012), is negatively correlated with 1994-2015 survey data on changes in credit standards on non-credit card consumer loans. Appendix 1 sets out details of index construction.

$CCI$  has several notable shifts, as shown in Figure 5. It dips below 0 in the credit crunch of 1966, before rising in a series of shifts to its peak of 1 in 2015 q2.  $CCI$  rises during the 1970s, punctuated by declines or pauses that coincide with Reg Q-induced disintermediation in 1970, 1973-74 and in the late 1970s and early 1980s. The index rose much following deposit deregulation through the imposition of tougher capital standards under Basel 1 in 1990. During

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<sup>4</sup> Hausman (1979) finds that, for cash-flow timing choices, households use quarterly discount rates of around 5%.

this time there were large rises in installment credit, typically used to purchase autos, home furniture and large appliances.

Other signs confirm a general increase in the ability of households to borrow. The timing and shape of the rise in the *CCI* also reflects those of the share of U.S. families owning bank credit cards—cards which do not require full monthly payments of outstanding balances and partly serve as a means of incurring debt (Figure 6). The relationship is less tight using a broader definition covering credit cards without this debt feature or which are usable at a particular retailer. For example, in 1970, 51% of families had cards using the broader definition, but only 16% had cards with general debt features. By 2001, this gap had disappeared. In this sense, the *CCI* picks up the distinction between the impact of credit card technology on transaction and debt services, that latter of which has far more important implications for consumption.

*CCI* drops during banks' transition to meeting tougher capital standards under Basel I. The index then rises moderately until the mid-1990s, by which time the scope for the securitization to alleviate the burden of capital standards had largely been used. The index was relatively flat from the mid-1990s to mid-2000s, an era when financial liberalization affecting households occurred mainly in mortgages, first enhancing the ability to withdraw housing equity from price appreciation and then to buy homes under weaker credit standards. In the mid-2000s the index rose notably, coinciding with the peaking of structured finance that funded much nonprime lending. The index, however, then fell to an extent similar to that seen in the credit crunch of the early 1980s, when consumer durable spending also had fallen sharply.

## 4. The Index of Housing Liquidity (*HLI*) and Mortgage Refinancing

### 4.1 Estimation Strategy

A housing liquidity index should track the not-directly-observable extent to which financial innovations have made it easier and less expensive for Americans to refinance their mortgage at a lower rate and/or borrow against the equity in their homes. Such latent effects allow for ease of mortgage equity withdrawals, enhancing the impact of housing wealth on consumption. We estimate these latent effects by estimating mortgage refinancing activity with controls for observable interest rate incentives to refinance and by estimating a housing liquidity index (*HLI*) in a system of equations with a latent variable – the liquidity of housing wealth, *HLI* - that is interacted with other variables.<sup>5</sup> The systems approach is used because it uses more information, and permits more precise estimation of *HLI*. The *HLI* interactions capture parameter variation over time in a parsimonious and economically meaningful way.

We employ the Kalman Filter to estimate the latent *HLI* in a non-linear state space model of system (Table 1) consisting of consumption (eq. 2.5) and mortgage refinancing equations (eq. 2.7 and 2.8 below). An appendix (available upon request) presents qualitatively similar estimates using a spline function based on annual smooth-transition dummies as an alternative way of estimating *HLI*.

#### 4.1 Mortgage Refinancing Data

We track mortgage refinancing with estimates of the percent of earlier mortgage debt packaged into Fannie Mae, Freddie Mac, or Ginnie Mae mortgage-backed securities (MBS) that was refinanced in quarter  $t$ . When a mortgage is refinanced without a cash-out withdrawal of homeowner equity, the outstanding stock of mortgage debt is unchanged, but originations rise. If the mortgages involved are securitized a refinancing raises the volume of newly issued MBS

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<sup>5</sup> We give this approach the acronym, LIVES, as an abbreviation for latent interactive variable equation system.

relative to the change in the stock of outstanding MBS. Based on this effect, Anderson and Duca (2016) calculate a time series of refinanced securitized mortgages as the gap between MBS originations and the change in the stock of MBS outstanding adjusted for principal payments.

Specifically, they calculate monthly unscheduled prepayments as equal to the volume of issuance of new MBS in month  $t$  minus the change in outstanding MBS from month  $t-1$  to month  $t$  minus scheduled payments of principal. This volume of refinancing in month  $t$  can be divided by the stock of outstanding mortgages at the end of month  $t-1$  to calculate what percent of mortgages in  $t-1$  were refinanced in month  $t$ .<sup>6</sup> By counting the amount of first liens refinanced and not any additional homeowner equity withdrawn in cash-out mortgage refinancings and the refinancing of home equity loans and other second liens, this method more closely tracks the percent of mortgages in month  $t-1$  were refinanced in month  $t$ . The benefits of refinancing arise mainly from the interest rate savings from refinancing old mortgage balances, but also for some from cashing-out equity by borrowing more than the original balance. Consistently tracking the former incentive is arguably more precisely estimated by empirically modeling the share of old mortgages that were refinanced, leaving the impact of the latter latent incentive and latent costs of refinancing to be gauged with state-space techniques in a system of consumption and refinancing equations.<sup>7</sup> By covering 18 more years than alternative (and less accurate) time series measures,<sup>8</sup> the Anderson-Duca data provide more interest rate and business cycles to estimate the coefficients of consumption and mortgage refinancing specifications.

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<sup>6</sup> For example, if a \$120,000 mortgage from month  $t-1$  were refinanced in month  $t$ , MBS issuance rises by \$120,000, with no net change in the outstanding stock of mortgages other than the scheduled principal payments (about .0015 percent on a monthly basis for mortgages with an average age of 4 years). If the \$120,000 mortgage were replaced by a \$140,000 mortgage (\$20,000 cashout), unscheduled principal prepayments equal the \$140,000 in new issuance minus the \$20,000 rise in net mortgage debt from the cash-out refinancing, which also equals \$120,000.

<sup>7</sup> Because MBSs have a trust structure, the guarantors make mortgages whole for investors against foreclosure losses, leaving MBS balances and prepayment calculations unchanged when mortgage foreclosures occur.

<sup>8</sup> Anderson and Duca (2015) show that this approach also has advantages over using the Mortgage Bankers Association (MBA) index of mortgage refinancing applications. First, MBA data starts later (1990) than the

These estimates of Fannie and Freddie mortgage refinancings are highly correlated with the dollar volume of new conventional mortgages from refinancings (Mortgage Bankers' Association (MBA) data) minus cashouts of conventional mortgages (Freddie Mac data) over 1993-2007. This period precedes large changes in FHA lending programs that shifted market share away from conventional mortgages. The Government National Mortgage Association (GNMA) securitizes FHA and VA mortgages and by including GNMA refinancings, Anderson and Duca avoid undercounting mortgage refinancings by accounting for FHA and VA mortgage refinancings. The resulting time series consistently tracks the annual volume of all mortgages refinanced (securitized and non-securitized in HMDA data) in contrast to the MBA series which tracks just securitized conventional mortgages. In addition, by tracking the stock of mortgages from which refinancings arise in their sample, Anderson and Duca can scale refinancings (by prior end-of-period outstandings) to calculate the mortgage refinancing rate shown in Figure 6.

#### 4.2 Empirical Specification for Modeling Mortgage Refinancings

The specification of the refinancing equation takes the basic form:

$$refi_t = rr_1 refi_{t-1} + rr_2 HLI_t + z'_t \delta + rr_2 (HLI_t \times z'_t \delta) + v_t \quad (2.7)$$

where  $HLI$  = the common factor housing liquidity index and  $z'_t \delta$  contains a constant and economic factors affecting the incentives to refinance. Since the entire function of variables is shifted by  $HLI$ , it has both level and interaction effects. The function  $z'_t \delta$  is given by:

$$z'_t = \delta_0 + \delta_1 PosGap_{t-1} + \delta_2 PosGap_{t-2} + \delta_3 Payback_t + \delta_4 Low_t + \delta_5 \Delta MortFore_{t-1} \quad (2.8)$$

The vector  $z$  includes the t-1 to t-2 lags of  $PosGap$ , which equals the maximum of 0 and the gap between the average interest rate on outstanding (existing) mortgages minus the average

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Anderson-Duca series (1973), thus omitting business cycles that help identify consumption coefficients. Second, rejection rates and delays in processing applications vary, whereas MBS liquidations more consistently reflect mortgage closings. Third, perhaps reflecting shifts in the composition of mortgage bankers and higher rejections rates, the MBA application index overstates total refinancings unlike the Anderson-Duca series in recent years.

interest rate on new mortgages used to purchase existing homes, with the gap scaled by the level of the average interest rate outstanding. The scaling reflects that a given rate gap has a larger percentage effect on house payments when existing rates start out lower. The variable *PosGap* is positive when there is a rate incentive to refinance, and should have positive coefficients apart from some dynamic unwinding effects (discussed below). The prevalence of fixed rate mortgages also implies that a given positive value of *PosGap* may not fully account for the possibility that new mortgage rates may appear to be at a low, when there is an additional incentive to lock in a low interest rate. To control for this effect, we include a dummy, *Low*, which equals 1 if the prevailing average new mortgage rate is at a 30-quarter low.

To further control for strong payback effects and a tendency for refinancing booms to abruptly end, we also include the *Payback*, equal to the product of a 0/1 dummy for the quarter following a mortgage rate low and the number of mortgage rate lows in the two years up to that quarter. The bigger the second element, the more households have refinanced in the two years leading up to the end of a down-cycle in mortgage interest rates, and the more likely is the payback effect to be more abrupt if mortgage rates rise off a low, as suggested by the sharper falls in refinancing following the two longest refinancing waves of 1992-3 and 2002-3.<sup>9</sup> We also include the lagged change in the overall mortgage foreclosure rate ( $\Delta MortFore_{t-1}$ , the rate at which mortgages are entering the process of foreclosure) to track the downside risk that housing collateral could lose value in the future when repossessions or short-sales occur and have lagging effects on house prices. Using a full set of variables allows us to strip out from refinancing

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<sup>9</sup> In other specifications we tried controlling for the advent of adjustable-rate mortgages, but found no evidence of significant effects on the mortgage refinancing rate. This may reflect that the absence of prepayment penalties on conventional FRMs and FHA mortgages has reduced the benefits of ARMs in the event of interest rate declines. Also, because ARMS were introduced in the early 1980s, the experience of interest rate volatility in the late-1970s and early 1980s may have deterred households from taking on the interest rate risk of ARMS.

activity all the effects not associated with financial innovations and to avoid contaminating estimates of *HLI* with endogenous factors.

## **5. Refinancing Results and *HLI* Estimates**

Before proceeding to the consumption results, it is instructive to review what the joint estimation model implies for mortgage refinancing behavior and the estimated *HLI* series.

### ***5.1. Refinancing Equation Results***

Table 2 reports results for jointly estimating the refinancing model with three slightly different versions of the consumption models reported in Table 3. Because these are very similar, we comment on results from the system using the preferred consumption model (#5 in Table 3). This refinancing model has a good fit and the residuals are relatively clean.<sup>10</sup> In order of the variables, after the lagged dependent variable, are two interest rate incentive terms. The t-1 lag of the asymmetric mortgage interest rate gap is positive and highly significant. The t-2 lag of this variable is negative and significant picking up the tendency for refinancing activity to decline two quarters after surging. The size of its coefficient roughly equals 90-95 percent of the magnitude of the positive coefficients on the t-1 lag, suggesting that the t-2 coefficient reflects the unwinding of incentives to refinance earlier. Also reflective of strong payback effects and a tendency for refinancing booms to abruptly end is the highly significant, negative coefficient on the term inter-acting the end of a mortgage rate low with the number of mortgage rate lows in the last two years. Also significant is the fifth interest rate incentive term, which is the time t dummy for mortgage interest rates being at their lowest level over the prior 30 quarters. Finally, as expected, the change in the foreclosure rate negatively affects refinancing. This is consistent

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<sup>10</sup> There is some evidence of heteroscedasticity, likely reflecting big Iraq war-related outliers in 2003.



with higher downside risk to collateral having negative loan supply effects and a decline in risk-adjusted housing wealth lowering the demand for mortgages.

### ***5.2 The Estimated Housing Liquidity Index (HLI)***

The *HLI* series is estimated using the nonlinear, two-equation state space model set out in Table 1. The state variable is only identified up to scale so the normalization means that *HLI* may be interpreted as the m.p.c. of housing wealth. Results based on a flexible spline function, augmented by one economic variable give a very similar profile for *HLI* (appendix available upon request).<sup>11</sup>

*HLI* has contours that are consistent with developments that likely affected the liquidity of housing (Figure 5). The *HLI* falls in the 1973-74 credit crunch, and again in the 1981-82 credit crunch, when binding Regulation Q ceilings hurt the ability of intermediaries to fund consumer and mortgage credit. The *HLI* rises a little in the late 1970s, coinciding with steps taken to deregulate bank deposits at a time when the mortgage-backed securities market was underdeveloped (Duca, 1996). The timing also coincides with the rise of second mortgages (Seiders, 1979). Afterward, apart from the temporary 1981-2 dip, *HLI* is flat for several years, before dipping a little in the late 1980s, perhaps because of the Savings and Loans crisis.<sup>12</sup> *HLI* plunges in the early 1990s credit crunch, when Basel 1 imposed higher capital requirements on mortgage loans held in portfolio than on securitized mortgages. This distinction was important because the market for securitized flexible interest rate mortgages and home equity loans was small.<sup>13</sup> *HLI*

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<sup>11</sup> The spline function uses annual smooth-transition dummies which rise in an S-shape from zero to 1 over 8 quarters. The dummy takes the values 0.05, 0.15, 0.3, 0.5, 0.7, 0.85, 0.95, 1 over an 8-quarter interval. One can also test down from the general form of the *HLI* function by sequentially setting to zero, the most insignificant of the annual smooth-transition dummies. These permit the shape of *HLI* to be smoothly non-linear, except for the effects of the economic variable, which is the quarterly change in the spread between the Aaa corporate bond and the 10-year Treasury yields to track changes in risk appetite and temporary spill-over effects from credit market tightening.

<sup>12</sup> Though at the same time financial sector productivity rose which lowered the costs of financial intermediation (Duca, 2005), probably moderating the effects of the S&L crisis.

<sup>13</sup> The smaller size of home equity loans relative to home purchase mortgages provided an additional (cost) hurdle.

begins to recover in 1993 near when Congress pressured Fannie Mae and Freddie Mac to expand mortgage lending. Increased mortgage securitization also occurred via home equity loans and cash-out mortgage refinancing. *HLI* surges between the late 1990s and mid-2000s, consistent with: declines in mortgage refinancing costs (Bennett, *et al.*, 2001); findings that proceeds from cash-out mortgage refinancings partially funded consumer spending (Canner *et al.*, 2002); and cross-section consumption results (Hurst and Stafford, 2004). *HLI* recedes a little after the 2003-4 peaks, drops somewhat as the housing market cools in 2005-6, then plunges in 2008-9, and recovers a little in 2010, before falling back again, with signs of only a partial recovery in 2014. The *HLI* implies a smaller mid-2000s peak in the m.p.c. out of housing wealth than estimated in other, related studies. For example, Carroll, *et al.* (2011) find that the housing wealth m.p.c. rose to about 9 percent in the late 1990s. However, their estimate may convolute the roles of unsecured consumer credit constraints and housing collateral as suggested by the high significance of our CCI in our estimated consumption function.<sup>14</sup> That said, the general pre-housing bust contours of their estimates are somewhat similar.

Nevertheless, our approach provides a more rigorous method for identifying shifts in the liquidity of housing and identifies a recent and large drop in the m.p.c. out of housing wealth. The recent fallback of *HLI* to levels of the mid-1990s is consistent with the imposition of large extra costs for—and tighter quantitative limits on—withdrawing equity when refinancing mortgages, which Fannie Mae and Freddie Mac imposed since late 2008. (FHA also tightened its criteria for cash-out mortgage refinancings). The recent bust in *HLI* also implies that the recent, partial recovery of U.S. home prices has not bolster consumption as much as suggested by the experience of the mid-1990 to mid-2000s.

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<sup>14</sup> Carroll *et al.*'s (2011) estimates also predate upward revisions to housing wealth in the Flow of Funds accounts.

## 6. Consumption Function Results

Estimates from seven consumption models are presented in Table 3. The first four show the advantages of using a version of the baseline specification in eq. (2.5) in which *HLI* is set to zero ( $\gamma_4 = 0$ ). Models 1-4 illustrate the benefits from progressively modifying the canonical LCH/PIH model of consumption by disaggregating wealth (Models 1 versus 2), then adding a control for labor uncertainty (Models 2 versus 3), and then also including an index for consumer credit conditions (Models 3 vs. 4). Each modification progressively improves model fit, with model 4 having a corrected  $R^2$  of 0.42. Permanent income growth, real interest rates, net liquid assets, net housing wealth, and the four quarter change in the unemployment rate are all significant with the expected sign. On the other hand, the speed of adjustment while respectable (29 percent per quarter) is not high, the residuals suffer from significant serial correlation, and illiquid wealth has an insignificant, albeit, positive m.p.c.

The next three models allow for a time-varying m.p.c. out of housing that is jointly estimated with a baseline mortgage refinancing equation. Of these, Model 5 is most comparable to Model 4 in the disaggregation of wealth and the inclusion of credit and uncertainty variables. Model 5 outperforms model 4 in terms of having a notably better model fit ( $R^2$  of 0.50 vs 0.42; S.S.E. of 0.47 vs. 0.50), a much faster speed of adjustment (44% versus 29% per quarter), clean residuals, and a significant m.p.c. out of illiquid assets. Of the wealth ratios, net liquid assets have the strongest impact with an estimated m.p.c. of around 0.08, somewhat below the UK estimates in Aron *et al.* (2012) (see Table 5 and that for Australia found by Muellbauer and Williams (2011). Illiquid financial assets including pension and stock market wealth have an estimated m.p.c. of 0.014, close to those found for the UK and Australia, but smaller than common estimates of 0.03 to 0.05 implied by consumption functions conditional on net worth.

Part of the reason is that standard models lack controls for income growth expectations and shifts in credit unlike the three papers listed above. In particular for the U.S., the University of Michigan index of expected economic conditions is strongly correlated with stock prices. Our findings accord with Poterba's (2000) point that stock market wealth effects partly embed growth expectations as well as a classical wealth effect. Another—and likely related—notable difference between Models 4 and 5 is that the average of the latter's time-varying m.p.c. out of housing (the *HLI*) is about one-half of the non-time varying housing wealth effect estimated in Model 4: 0.019 vs. 0.041.

The key results from Model 5 are generally robust to the exclusion of the *CCI* (Model 6) or the proxy for labor market uncertainty ( $\Delta u$ , Model 7). One exception is that permanent income growth is insignificant in model 7 if the significant four-quarter change in unemployment is dropped.<sup>15</sup> Another difference is that when the consumer credit conditions index (*CCI*) is omitted in Model 6, the time-varying m.p.c. out of housing wealth has an average (0.041) that is twice that of Model 5 (0.019), with a notably higher peak (0.051 versus 0.034). This likely reflects the impact of omitted factors as discussed earlier.

Using the consumption function estimates from model 5 in Table 3 that includes *CCI* and *HLI*, we decompose how much the equilibrium consumption-to-income ratio fell in response to credit and wealth effects. As a pre-crisis benchmark, we use 2007:q3, the quarter in which the financial crisis started to disrupt the Labor markets. Between 2007:q3 and 2009:q4, the ratio of consumption to non-property income fell 7.2 percentage points. The long-run equilibrium ratio implied by the two-equation system tracks this ratio remarkably well as shown in Figure 9.

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<sup>15</sup> A plausible explanation is that the omission of tracking uncertainty with the year-over-year change in the unemployment rate makes it difficult to identify the impact of changes in permanent income, especially in economic recessions and recoveries.

Based on the long-run coefficient estimates in Table 3, the model implies that the long-run equilibrium consumption-to-non-property income ratio fell by 6.8 percent. Of this, 1.7 percentage points was attributable to the fall in *CCI* and 5 percentage points to the combination of declines in housing wealth and housing liquidity. The latter is partially offset by about a 2 percentage point rise in the equilibrium consumption-to-income ratio associated with declines in mortgage debt. Some of the fall-back in mortgage debt stems from voluntary repayment of debt or not taking on new debt; but some will arise from the writing off of bad debts. Further deleveraging by households, coupled with recoveries in house prices, stock prices and consumer credit availability have induced a significant recovery in consumption.

Figure 10 plots the consumption-to-income ratio and its key long-run drivers: the fitted long-run components due to net liquid assets/income, the consumer credit index and housing wealth/income scaled by the housing liquidity index. The last two account for a major part of the secular rise in the consumption-to-income ratio, as well as its recent sharp fall. However, there is a major offset from the accumulation of debt, a consequence of credit market liberalisation, which pulls down net liquid assets/income. Since the m.p.c. out of net liquid assets is far larger than out of illiquid assets, this offset is substantial. Although higher income growth expectations help explain some phases of the rise in consumption relative to income, such as in the early 1980s and the mid-1990s, they cannot account for the rise after 1997. Also, the scale of variation implies that one cannot base much of a long-run story on this source. These cast doubt on the contention that the rise in U.S. consumption of the 2000-09 decade owed to large increases in expected growth income—if anything, income growth expectations appear to have down-shifted from the 1990s. Another ‘long-run’ fitted component reveals that the upward trend in illiquid financial wealth accounts for some of the upward drift and cyclical fluctuations in the

consumption-to-income ratio. The impact of the real interest rate on auto loans also has little long-run effect, although changes in it help explain short-run dynamics of consumer spending.

It is instructive to compare the above estimates of the consumption function including the two credit-friction indicators to estimates excluding them. Table 4 shows estimates for samples up to 2007:q3 and to 2015:q2 for consumption equations omitting the credit conditions index, *CCI* and using the housing wealth to income ratio, not weighted by *HLI*. Columns 1 and 2 omit both, while columns 3 and 4 include *CCI*. The speeds of adjustment are low, particularly when *CCI* is excluded. Results from estimating the credit-augmented model 5 from Table 3 over the full and pre-Great Recession sample (up to 2007:3)—shown in columns 5 and 6 respectively—yield similar coefficients, with one exception. That difference is the end-of-sample, time-varying m.p.c. out of housing wealth, which is about half as large in 2015:q2 as in 2007:q3. The patterns seen across columns 5 and 6 also generally hold when comparing results from these samples in corresponding models (columns 7 and 8) that use net housing wealth instead of gross housing wealth, and which use net nonmortgage liquid assets instead of net liquid assets. The only exceptions are that the effect of the credit conditions index is about 40 percent smaller in the net housing wealth models and these models do not fit the data as well.

We can interpret these findings with our full credit-augmented consumption function. The omission of the interaction of *HLI* with housing wealth relative to income results in a downward bias on the coefficient of net liquid assets relative to income, as increasing housing liquidity drove up debt, so that net liquid assets turned negative in the mid-2000s. The credit-augmented consumption function approach, combined with our disaggregation of net wealth components, has important implications for the downswings that follow consumption booms fueled by rising house prices and mortgage borrowing, such as those of the late 1990s and mid-

2000s. As a result of the increases in the liquidity of housing wealth during the late 1990s, the moderate increases in house prices then and the sharper rises of the mid-2000s induced greater mortgage borrowing that at first boosted consumption. During the early phases of such consumption booms, the positive impact of rising housing wealth overwhelms any drag from higher debt. Later, when real house prices stopped rising, the drag from previously built up debt predominates, giving way to reduced consumption and deleveraging. A fallback in housing liquidity has exacerbated the negative payback effect of the house price boom of the mid-2000s. The negative payback or deleveraging phase arises in our model, which disaggregates net wealth, because net liquid assets have a higher estimated m.p.c. (15%) than gross housing wealth (6% at the peak). This feature of our framework, combined with slow recoveries in consumer credit availability in these episodes, helps account for why consumption was slow to recover early in the recovery of the early 2000s before house prices surged. These features are much less notable in the conventional consumption function and would have been absent from models estimated up to the late 1990s.

## **7. Conclusion**

Assuming that capital markets are perfect under certainty equivalence yields the canonical type of saving function based on the permanent income-life cycle hypothesis. We find that imposing market completeness and certainty equivalence can render consumption models, much as with asset price models, less useful for understanding and tracking cycles and disequilibria. The existence of credit constraints and major shifts in credit availability can imply departures from those highly stylized models, and may explain why traditional models have generally failed to track the recent decline in consumption and the boom that had preceded it,

along with the longer historical record. In addition, by explicitly modeling the factors driving the long run evolution of the consumption-to-income ratio, our LIVES approach accounts for important parameter shifts in the basic responses of consumption to wealth, credit, and income shocks. Consequently, by not ignoring long-run information by detrending and linearizing, our approach avoids the parameter instability that often plagues conventional linear VARs, especially during the recent recession.

Consistent with our credit-augmented life-cycle/permanent income approach, we find that indexes tracking changes in the availability of consumer credit and the liquidity of housing wealth greatly improve empirical models of consumer spending. These indexes indicate that consumer credit markets became more complete during the 1980s, while the liquidity of housing wealth rose in the late 1990s. Our results imply that differences in the timing of these innovations are statistically and economically important. In addition, adding these channels enables us to gauge the impact of the financial crisis on consumption, via both its short-run effect on some types of financial frictions (e.g., the LIBOR-OIS spread) and by other elements that may have longer-term effects on credit availability and the ability of homeowners to tap housing equity. Overall, our findings imply that it is important to carefully account for financial liberalization and innovation when modeling consumption.

One particular contribution from this study is its construction of a levels index for the availability of consumer credit. This index is constructed by removing short-run cyclical influences from a diffusion index of the change in bank lending and then scaling the resulting diffusion index using its common sample growth rate versus that of consumer loan extensions relative to income over 1966-82. Including this index notably improves model fit and characteristics (e.g., increase the speed of adjustment). Removing short-term cyclical influences



from the index improves on the original version of the index used in Muellbauer (2007), adopted in the President's *Economic Report 2010* to model long-run variations in the U.S. saving rate.

Another data contribution of this study is its construction of a time series for the level of housing liquidity. We specify a model for mortgage refinancing activity that includes many plausible economic control variables, including financial incentives to refinance such as lower interest rates, and changing interest rate expectations. Using our two-equation system, we extract a common latent index whose trends are consistent with other evidence of major declines in the pecuniary and non-pecuniary costs of refinancing mortgages. We show how gleaning information from refinancing behavior yields more plausible and less noisy estimates of the m.p.c. of housing wealth. In addition, movements in this index coincide with major shifts in business practices and regulations. In this way, our estimated *HLI*, in conjunction with other information, sheds light on the changing sensitivity of mortgage refinancing activity to interest rate incentives to replace old mortgages and to swings in house price appreciation. As a result of underlying financial innovations and incentives from the Tax Reform Act of 1986, the collateral role of housing became enhanced over the years leading up to the recent housing bust, as had the effects of mortgage rate and house price swings on MEW.

The combination of large declines in wealth and substantial tightening of mortgage and consumer credit standards in 2007 to 2010 has not been seen since the recession of 1974-75, when U.S. consumption was also unusually weak. Our estimates and calibrations indicate that the equilibrium ratio of consumption to non-property income fell by 6.8% from mid-2007 to year-end 2009, in line with actual data.<sup>16</sup> Estimates imply that about one-quarter of the rise in the personal saving rate during that time stemmed from tighter credit standards and, about three quarters, from wealth effects. The latter not only reflected prior increases in the impact of

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<sup>16</sup> Because of partial adjustment, the equilibrium ratio falls by somewhat more than the actual ratio over this interval.

housing liquidity, but also asset price declines associated with declines in credit and mortgage availability, the latter of which also reflected tighter credit standards on mortgages for home purchases as shown by Duca, Muellbauer, and Murphy (2011a). Declines in 2007-2010 in consumer credit standards partly owe to shifts in LIBOR spreads that have affected the inter-bank lending market which helps banks fund loans. In this way, our *CCI* index is affected by financial frictions that are associated with the broader financial and credit crisis of 2007-09.

Peculiarly American institutions and history underlie the empirical modeling in this paper. There are interesting parallels with the UK (Aron, *et. al* (2013) and Australia (Muellbauer and Williams, 2011) where there have also been substantial housing collateral effects on consumption, though the shifts in credit market architecture and decline in lending standards in the 2000s were far less pronounced than in the U.S. The contrasts with Canada are far more pronounced. Muellbauer, St. Amant and Williams (2015) find no evidence of a housing collateral effect on consumption. Indeed for much of the period since 1980, higher house prices relative to income had a negative effect on consumption, via the down-payment constraint, though credit liberalization seems to have more or less eliminated this mechanism by 2008. These differences are attributable to important institutional differences. In contrast to the U.S., in Canada there is no tax deduction on mortgage interest, mortgages are mainly full-recourse, inducing individuals to be cautious, and a conservative, highly regulated banking system limits risky mortgage practices—including mortgage equity withdrawals.<sup>17</sup> Institutional differences across countries, as well as the evolution of institutions, are critical to understanding the cross-country and time-varying patterns of aggregate household saving and consumption.

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<sup>17</sup> For example, there is a system of compulsory, well-supervised insurance for mortgages with loan-to-value ratios above 80%.

## Appendix 1: Construction of the unsecured credit conditions index

We adjust the willingness to lend index for the 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.) We track the former by including the first difference of the real federal funds rate ( $\Delta r_{ff}$ , 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 ( $\Delta_2 LEI$ ). To further adjust for factors affecting consumer loan quality, we include the time  $t$  year-over-year change in the delinquency rate on bank consumer installment loans ( $\Delta_4 Del$ , American Bankers Association).

We include three regulatory variables. One is a dummy equal to 1 in 1980 q2 when credit controls were imposed and equal to -1 when they were lifted in 1980 q3 (*CrControl*). Another (*RegQ*) measures the degree to which Regulation Q interest rate ceilings reduced banks' ability to attract deposits (Duca, 1996; Duca and Wu, 2009) and thereby raised banks' shadow cost of loanable funds in an era before the loan sales and mortgage-backed securities markets became deep. The third regulatory variable (*MMDA*) is a dummy equal to one in 1982 q4 and 1983 q1 to control for the re-intermediation effects of allowing banks to offer variable interest money market deposit accounts, which boosted deposits (Duca, 2000).

After Reg Q was lifted, the interbank funding market increasingly became a marginal source of loanable funds, with the 3-month LIBOR normally exceeding the expected 3-month average federal funds rate by about 10 basis points. To control for this, we include the spread

between the 3-month LIBOR and 3-month Treasury Bill rates (*Libor Spread*). We also include a dummy (*Lehman* = 1 in 2008:q4) for the failure of Lehman (which was after the 2008:q3 Fed survey). Estimating the model from 1966 q3 to 2015 q2 with an AR(1) correction yields:

$$\begin{aligned}
 CR = & 15.18 - 2.42 \Delta rff_t^{**} + 0.77 \Delta_2 LEI_t^{**} - 8.82 \Delta_4 Del_t^+ + 26.22 MMDA_t^{**} \\
 & (4.56) \quad (3.56) \quad (4.11) \quad (1.95) \quad (7.64) \\
 & - 7.90 RegQ_t^{**} - 50.83 CrControl_t^{**} - 4.93 Libor Spread_t^* - 20.49 Lehman_t^{**} \quad (2.6) \\
 & (3.01) \quad (6.62) \quad (2.37) \quad (3.12)
 \end{aligned}$$

where t-statistics are in parentheses,  $R^2 = 0.78$ ,  $AR(1) = 0.75^{**}$  (15.24), equation standard error = 9.09,  $LM(2) = 0.43$  and  $Q(24) = 24.34$ . The coefficients are significant with the expected signs. Reassuringly, coefficients hardly change in samples before the financial crisis started in August 2007 and its peak effects on interbank lending in late 2008. We subtract the estimated impact of changes in the real federal funds rate, leading economic indicators, and the delinquency rate to remove cyclical and interest rate effects, leaving the impact of regulations, Lehman's fall, unusual credit (*Libor Spread*) frictions, and unexplained variation in the adjusted diffusion index (*CRAdj*). The adjusted *CR* index was then chained into a levels index, based on its correlations with the growth rate of real consumer loan extensions at banks, and normalized (see Figure 2).

## Appendix 2: Modeling Income Expectations

Estimating equation (2.5) requires measuring income growth expectations. We choose a subjective discount rate of 5% per quarter as noted above and construct  $E_t \ln(y_t^p / y_t)$  defined by equation (2.3) taking a horizon of 40 quarters. This is more forward-looking than Friedman's (1963) three-year horizon but less forward-looking than is usually assumed in DSGE models. After 2009 we assume that the historical growth rate resumes from 2010 q1, building in a permanent component of the 'Great Recession.'

$\ln(y_t^p / y_t)$  is regressed for 1961 to 2009 on a constant, trend, a 1968 split trend for the productivity slowdown,  $\log y$ ,  $\Delta_4$  T-bill yield, and the University of Michigan index of consumer expectations of future economic conditions. Estimating the same equation for 1961 to 2006 results in almost identical coefficients and fit, suggesting the assumptions made about income after 2009 q4 are consistent with the estimated equation. Figure 6 shows the fitted value against the actual value of  $\ln(y_t^p / y_t)$ , given post-2009 assumptions on income. Since 1970, the fitted value has remained in the range 0.02 to 0.1, with a low in 1979 and a high in the late 1990s.

The joint estimation results correspond very well with theoretical priors. An initial general specification was estimated in which the housing liquidity index enters both as an intercept and in interaction with demeaned income growth expectations and housing wealth to income ratio and similarly in the MEW equation. This is compared with a restricted specification in which there is no intercept role for *HLI* in either equation but only interaction effects with income growth expectations and the housing wealth-to-income ratio, *not* demeaned, and the level effect of the housing wealth-to-income ratio is zero. The difference in twice log likelihood between the two specifications is 4.48 and is asymptotically chi-squared. With four restrictions the 5% critical value is 9.49 so that the restricted specification passes easily.

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Figure 1: Net Worth-to-Income Ratio Alone Cannot Account for Saving Rate Trends

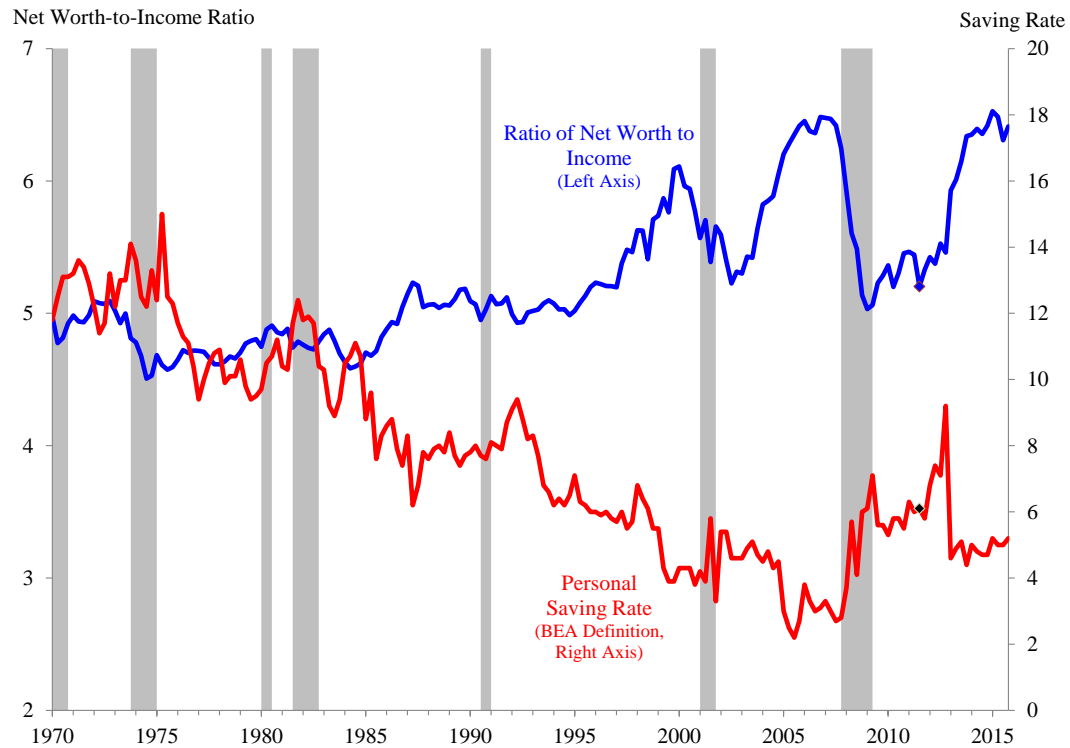


Figure 2: Log Ratio of Consumption to Non-property Income and Scaled Illiquid Financial assets and Housing Wealth Relative to Non-property Income

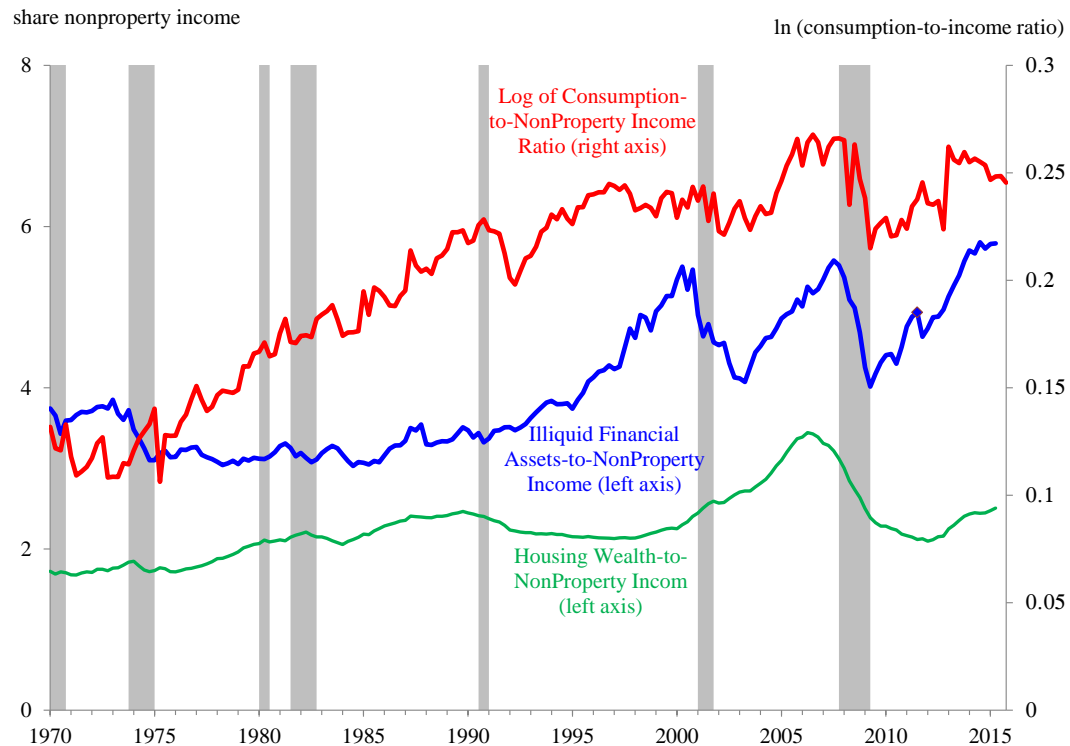


Figure 3: Financial Innovation Linked to Changes in MEW Sensitivity to Swings in Real House Price Appreciation

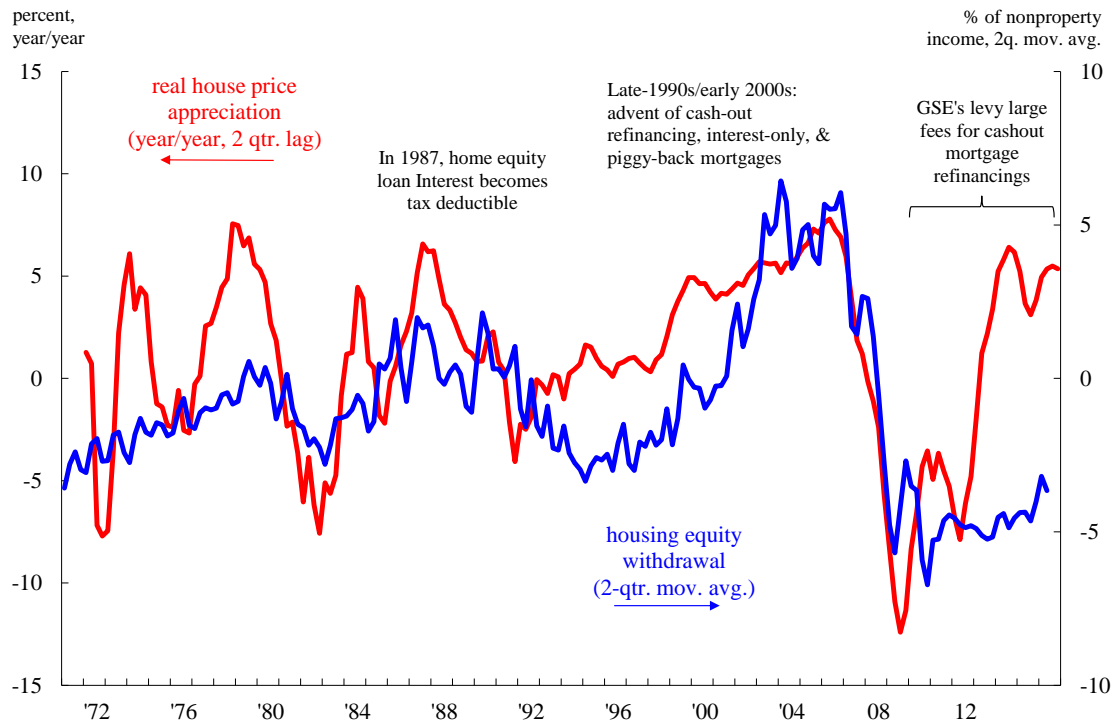


Figure 4: Fitted values of log permanent income/actual income

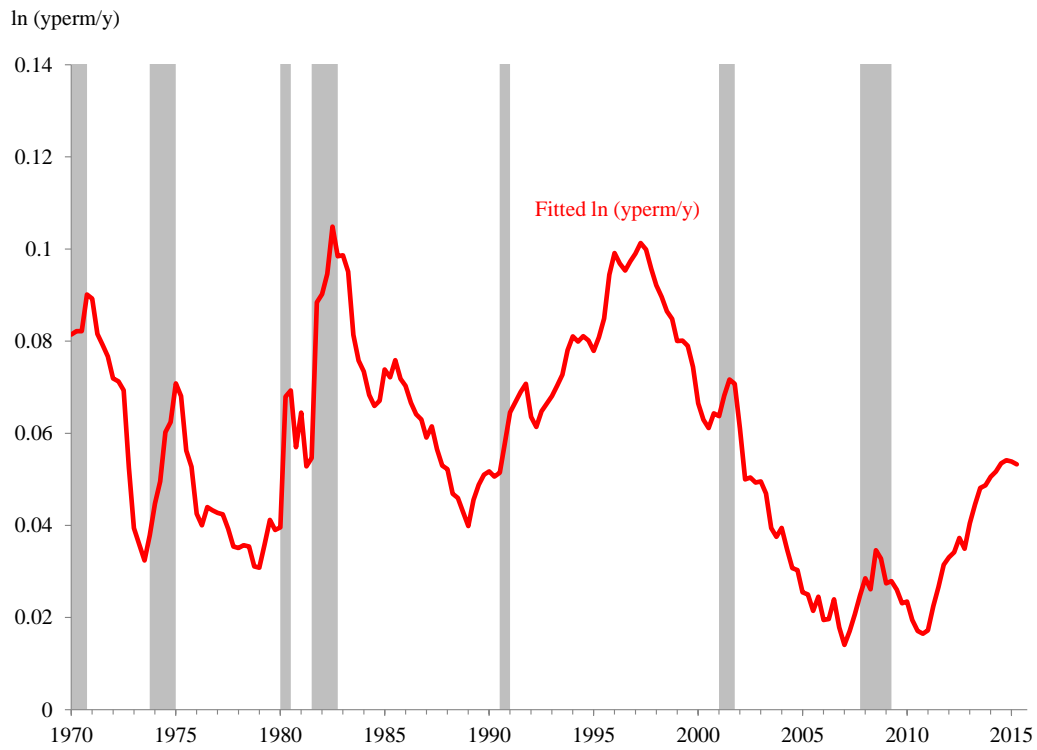


Figure 5: Consumer Credit Availability Rises Much from 1970 to the Mid-1990s, Falls During Recent Bust and Then Recovers

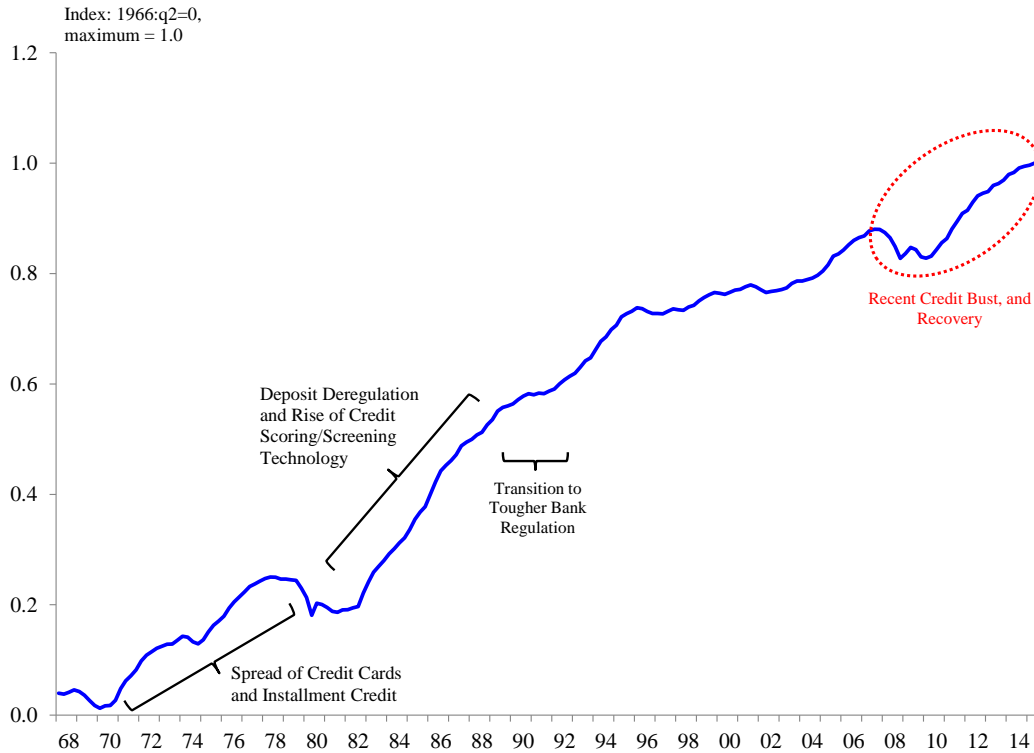
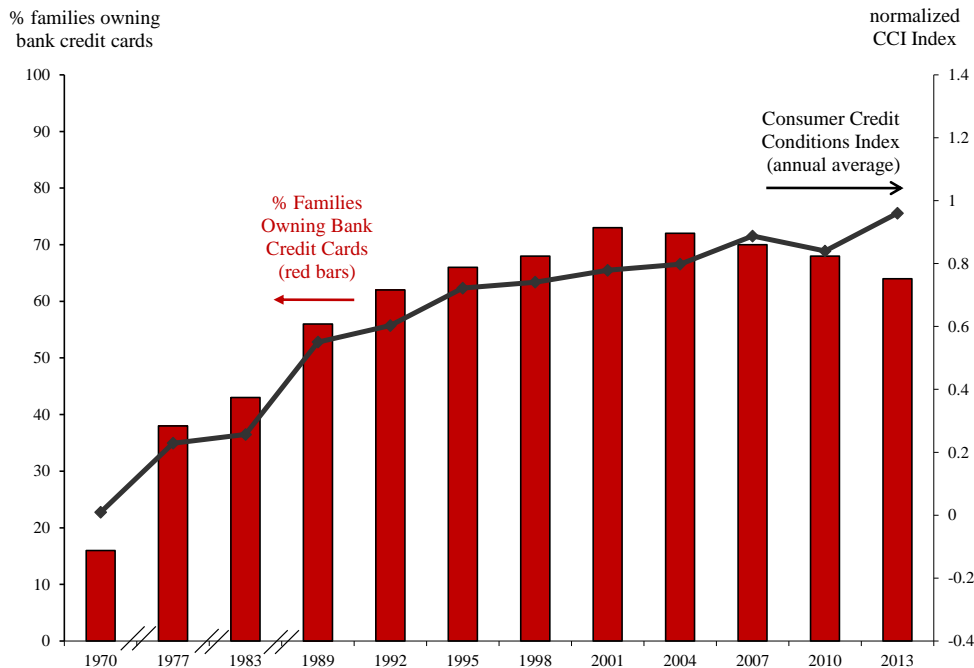


Figure 6: The Consumer Credit Conditions Index (CCI), Tracks the Rise of Bank Credit Card Ownership Rates



Notes: All credit cards generally excludes cards limited to only one particular retailer. Bank cards are those on which households can carry-over balances. Sources: Durkin (2000), Bertaut and Haliassios (2006) for 1992 data, Bucks, et al., (2007, 2009) for 2001-07, and authors' calculations using Bucks, et al. (2009) figures for bank card ownership in 2004 and 2007.

Figure 7: U.S. Financial Innovations Five Rise to Changes in the Sensitivity of the Share of Mortgages Refinanced to Mortgage Interest Rate Differentials

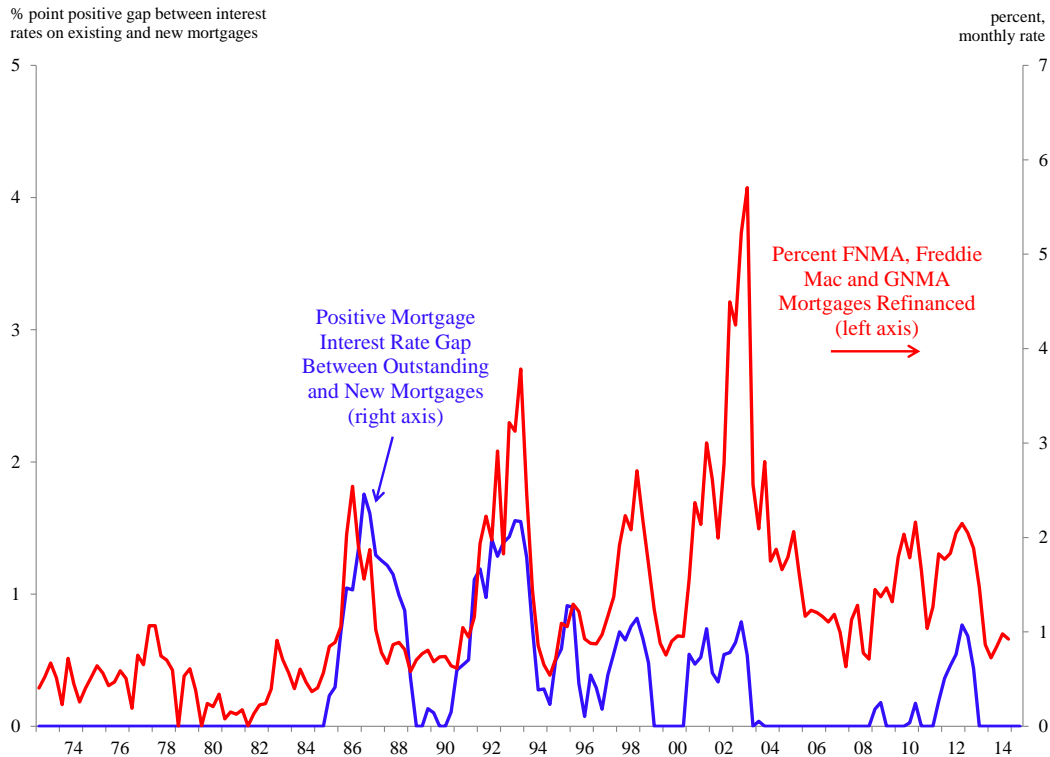
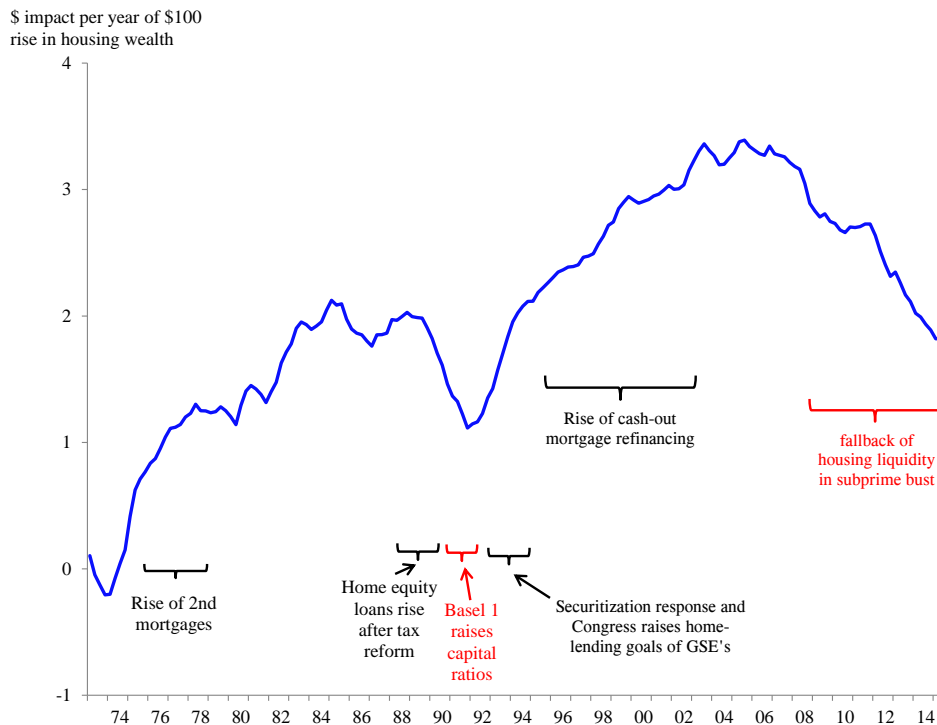
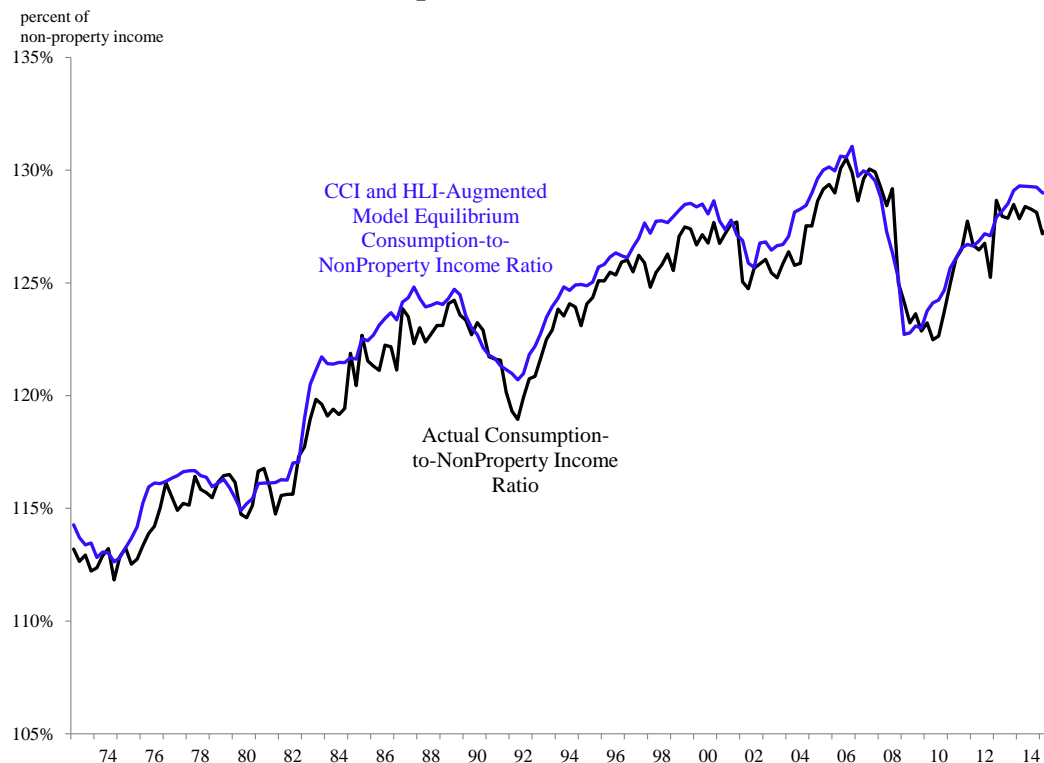


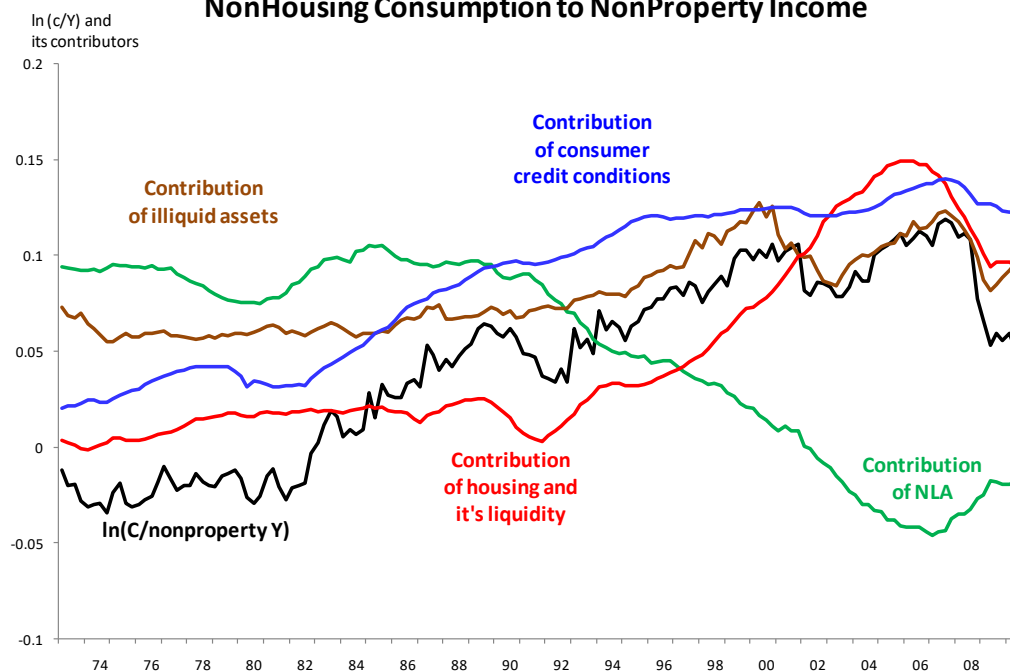
Figure 8: Housing Liquidity and Marginal Propensity to Consume out of Housing Wealth Rises in Late-1990s, Retreats Some During the Subprime Bust



**Figure 9: Long-run Equilibrium Relationship in Credit-Augmented Model Tracks the Fall in the Consumption-to-Income Ratio Since the Financial Crisis**



**Figure 8: Estimated Equilibrium Components of Log Ratio of NonHousing Consumption to NonProperty Income**



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**Table 1: The Two-Equation State Space Model**

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**1. Consumption Function:**

$$\begin{aligned}\Delta \ln c_t = & \lambda \left\{ \alpha_0 + (\ln y_t - \ln c_{t-1}) + \alpha_1 r_{t-1} + \alpha_2 CCI_{t-1} + \alpha_3 (\ln y_{t+1}^p - \ln y_t) \right. \\ & + \gamma_1 NLA_{t-1}/Y_t + \gamma_2 IFA_{t-1}/Y_t + \gamma_3 HSG_{t-1}/Y_t + \gamma_4 HLI_t \times HSG_{t-1}/Y_t \left. \right\} \\ & + \beta_1 \Delta \ln y_t + \beta_2 \Delta ur_t + \beta_3 \Delta nr_t + \dots + u_t\end{aligned}$$

**2. Refinancing Equation:**

$$\begin{aligned}ref_t = & rr_1 ref_{t-1} + rr_2 HLI_t + z_t' \delta + rr_2 (HLI_t \times z_t' \delta) + v_t \\ z_t' = & \delta_0 + \delta_1 PosGap_{t-1} + \delta_2 PosGap_{t-2} + \delta_3 Payback_t + \delta_4 Low_t + \delta_5 \Delta MortFore_{t-1}\end{aligned}$$

**3. State Equation:**

$$HLI_t = HLI_{t-1} + \varepsilon_t$$


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Notes: The random error terms  $u_t$ ,  $v_t$  and  $\varepsilon_t$  are independent, mean zero normal random errors and the normalization  $\gamma_4 = 1$  is used.  $\ln y_{t+1}^p - \ln y_t$  is the OLS fitted value of  $\left( \sum_{s=1}^K \eta^{s-1} \ln(y_{t+s}/y_t) \right) / \left( \sum_{s=1}^K \eta^{s-1} \right)$ , with  $K = 40$  and  $\eta = 0.95$ , in an OLS regression 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 two other explanatory variables - the four-quarter change in the three-month Treasury bill yield and the Thomson Reuters, University of Michigan survey measure of consumer expectations.



**Table 2: Two-Equation State Space Model Estimates of the Pace (Rate) of Mortgage Refinancing, 1972 Q1 to 2015 Q2**

Dependent variable = $Refi$ percent of mortgages in quarter t-1 refinanced in quarter t	Time-Varying Housing Liquidity Models from Table 3		
	Table 3, Consumption Model 5 Including Consumer Credit Conditions and 4 qtr change in unemployment rate	Table 3, Consumption Model 6 Omitting $CCI$ , the Consumer Credit Conditions Index	Table 3, Consumption Model 7 Omitting uncertainty as tracked by the 4 quarter change in the unemployment rate ( $\Delta u$ )
$z'_t$ Part of the Refinancing Equation			
$PosGap_{t-1}$ , Positive Interest Rate Gap <sub>t-1</sub> x 100	1.047** (5.7)	1.030** (5.8)	1.038** (5.6)
$PosGap_{t-2}$ , Positive Interest Rate Gap <sub>t-2</sub> x 100	-0.983** (5.6)	-0.998** (5.9)	-0.991** (5.6)
$Payback_t$ , Unwinding of prior refi surge x100	-0.166** (5.6)	-0.168** (5.9)	-0.166** (5.7)
$Low_t$ , Mortgage interest rate low x100	0.474** (5.8)	0.439** (5.7)	0.482** (5.9)
$\Delta MortFore_{t-1}$ , Change mortgage foreclosure rate x 100	-1.067** (2.6)	-0.923* (2.4)	-1.042* (2.5)
$Refi_{t-1}$ , Lagged mortgage refinancing rate	0.754** (18.7)	0.773** (22.7)	0.758** (18.3)
$HLL_t$	0.116* (2.5)	0.067** (3.0)	0.109* (2.4)
Log Likelihood	1427.4	1422.4	1422.2
AIC	-16.2	-16.1	-16.1
SIC	-15.8	-15.8	-15.8

Notes: t-statistics are in parentheses. <sup>\*(\*\*,\*)</sup> denotes significance at the 95% (99%, 90%) confidence levels. The refinancing equation is

$$refi_t = rr_1 refi_{t-1} + rr_2 HLL_t + z'_t \delta + rr_2 (HLL_t \times z'_t \delta) + v_t \text{ with: } z'_t = \delta_0 + \delta_1 PosGap_{t-1} + \delta_2 PosGap_{t-2} + \delta_3 Payback_t + \delta_4 Low_t + \delta_5 \Delta MortFore_{t-1}$$

**Table 3: U.S. Consumption Function OLS Estimates for 1972 Q1 to 2017 Q2**

Dependent variable = $\Delta \ln c$	Time Invariant Wealth MPC Models				Time-Varying Housing Liquidity		
	(1) Basic Model	(2) Split Wealth	(3) Split Wealth uncertainty	(4) Add CCI	(5) <i>HLI</i> + uncertain -ty term	(6) =(5) No CCI	(7) = (5) ex. uncertainty term ( $\Delta Au$ )
Speed of Adjustment	0.074** (2.7)	0.085** (3.5)	0.105** (4.6)	0.165** (4.2)	0.378** (9.0)	0.401** (9.3)	0.384** (8.5)
<i>Long-Run Effects:</i>							
Intercept	0.207 (1.5)	0.192 (1.6)	0.189* (2.1)	0.160** (2.8)	0.039 (1.2)	0.083* (2.0)	0.057 (1.2)
Credit Conditions Index, $CCI_t$				0.063** (2.9)	0.061* (2.0)		0.050 (1.4)
Real interest rate on autos, $r \times 100$	-0.913+ (1.9)	-1.086** (2.7)	-0.780** (2.7)	-0.559** (2.8)	-0.139 (1.2)	-0.196+ (1.8)	-0.192 (1.6)
Forecast income growth, $\Delta \ln y^{perm}$	1.054* (2.5)	0.892** (3.2)	0.930** (4.3)	0.709** (4.7)	0.269* (2.3)	0.242* (2.1)	0.079 (0.8)
Net worth / income, $(NW/Y)_{t-1}$	0.034* (2.3)						
Net liquid assets/income, $(NLA/Y)_{t-1}$		0.091* (2.1)	0.021 (0.7)	0.020 (1.0)	0.075** (2.8)	0.064* (2.2)	0.089** (3.0)
Illiquid financial assets / income, $(IFA/Y)_{t-1}$		0.034** (3.4)	0.018* (2.3)	0.052 (0.8)	0.013* (2.0)	0.018* (2.4)	0.022** (3.3)
Housing wealth / income $(HSG/Y)_{t-1}$		0.050** (2.8)	0.050** (3.7)	0.045** (5.3)			
Housing wealth x housing liquidity, $(HSG/Y)_{t-1} \times HLI_{t-1}$ (final state)					0.023** (6.5)	0.038** (10.2)	0.023** (5.4)
Peak value of time-varying housing liquidity index					0.034	0.051	0.036
<i>Short Run Effects:</i>							
Change in real user cost cars, $\Delta r \times 100$	-0.195** (4.1)	0.205** (4.3)	-0.231** (5.1)	-0.219** (5.0)	-0.159** (4.0)	-0.181** (4.2)	-0.154** (3.7)
Change in unemployment rate, $\Delta Au \times 100$			-0.194** (5.0)	-0.215** (5.5)	-0.196** (3.9)	-0.183** (3.5)	
1980:q2 Credit Controls x 100				-1.088** (3.1)	-0.196** (3.9)		-0.929** (4.7)
Adjusted R <sup>2</sup>	0.31	0.31	0.39	0.44	0.50	0.47	0.47
DW	1.64	1.64	1.81	1.74	2.06	2.10	2.12
LM AR/MA(4) – P Value	27.52**	24.02**	10.08*	20.65**	0.27	0.46	0.15
LB Q Stat(24) – P value	98.55**	107.45**	57.03**	58.21*	0.30	0.12	0.25

Notes: t-statistics are in parentheses. <sup>\*(\*\*,\*)</sup> denotes significance at the 95% (99%, 90%) confidence levels. The equation SE's, adjusted R<sup>2</sup>'s and regression diagnostics from the state space models are from OLS regressions, treating the estimated *HLI*'s as given. The general model is:  $\Delta \ln c_t \sim \lambda (\alpha_0 + (\ln y_{t-1} - \ln c_{t-1}) + \alpha_1 r_{t-1} + \alpha_2 (\ln y^{perm} - \ln y) + \alpha_3 CCI_{t-1} + \gamma_1 (NLA/Y)_{t-1} + \gamma_2 (IFA/Y)_{t-1} + \gamma_3 (HSG/Y)_{t-1} \times HLI_{t-1} + \varepsilon_t$ .

**Table 4: U.S. Consumption Function OLS Estimates for 1972 Q1 to 2015 Q2 and 1972 Q1 to 2007 Q3**

Dependent variable = $\Delta \ln c$	Non-Time-Varying Wealth MPC Models			Time-Varying Housing Liquidity Models			
	(1) Basic Model	(2) Split Wealth	(3) Split Wealth + $CCI, \Delta u$	(5) $HLI \times GrossHsg$ 72:1-15:2	(6) $HLI \times GrossHsg$ 72:1-07:3	(7) $HLI \times NetHsg$ 72:1-15:2	(8) $HLI \times NetHsg$ 72:1-07:3
Speed of Adjustment	0.147** (4.5)	0.239** (5.7)	0.294** (4.0)	0.440** (10.5)	0.545** (10.5)	0.436** (10.3)	0.563** (9.69)
<i>Long-Run Effects:</i>							
Intercept	0.031** (3.1)	0.027* (2.5)	0.033* (3.3)	0.062 (1.5)	0.046 (1.1)	0.020 (0.4)	0.016 (0.3)
Credit Conditions Index, $CCI_{t-1}$			0.096* (5.0)	0.088** (2.8)	0.093* (2.4)	0.055* (2.4)	0.053* (1.7)
Real interest rate on autos, $r \times 100$	-0.112** (3.3)	-0.148** (4.4)	-0.121** (3.9)	-0.177* (1.7)	-0.164* (1.7)	-0.173* (1.7)	-0.158* (1.7)
Forecast income growth, $\Delta \ln y_{perm}$	0.110** (4.9)	0.111** (4.6)	0.143** (6.3)	0.224* (2.3)	0.193* (2.1)	0.301** (3.1)	0.234* (2.4)
Net worth / income, $(NW/Y)_{t-1}$	0.033** (3.0)						
Net liquid assets/income, $(NLA/Y)_{t-1}$ (ex. mortgage debt in models (7) & (8))		0.070** (4.2)	0.028* (2.1)	0.081** (3.0)	0.105** (3.5)	0.085** (2.9)	0.092** (3.1)
Illiquid financial assets / income, $(IFA/Y)_{t-1}$		0.016** (3.2)	0.005 (1.4)	0.014* (2.2)	0.013* (1.9)	0.015* (2.5)	0.012* (1.9)
Housing wealth / income $(HSG/Y)_{t-1}$		0.045** (4.4)	0.041** (5.3)				
Housing wealth $\times$ housing liquidity, $(HSG/Y)_{t-1} \times HLI_{t-1}$ end of sample, (ex. mortgage debt in models (7) & (8))				0.019** (7.3)	0.040** (18.1)	0.019** (4.6)	0.036** (10.6)
Peak value of time-varying housing liquidity index				0.034	0.040	0.040	0.044
<i>Short Run Effects:</i>							
Change in real user cost cars, $\Delta r \times 100$	-0.215** (4.4)	-0.233** (4.9)	-0.247** (5.7)	-0.207** (5.4)	-0.207** (4.5)	-0.205** (5.3)	-0.205** (4.5)
Change in unemployment rate, $\Delta u \times 100$			-0.244** (6.1)	-0.189** (3.8)	-0.174** (2.8a)	-0.248** (4.9)	-0.218** (3.2)
Standard Error $\times 100$	0.58	0.56	0.50	0.47	0.48	0.47	0.49
Adjusted R <sup>2</sup>	0.24	0.30	0.42	0.50	0.48	0.49	0.47
DW	1.33	1.31	1.51	2.06	2.10	2.07	2.10
LM AR/MA(4) – P Value	45.03**	16.66**	25.75**	0.27	0.39	0.40	0.32
LB Q Stat(24) – P value	135.71**	167.53**	68.01**	0.30	0.27	0.26	0.22

**Table 5: Estimated Wealth Effects\***

	MPC out of net liquid assets	MPC out of illiquid financial assets	Peak MPC out of housing wealth
U.S. - consumption excluding housing services	0.081	0.014	0.034
UK - total consumption	0.114	0.022	0.043
Australia - total consumption	0.159	0.022	0.049

\*Estimated m.p.c.'s from the preferred models for the UK (column 4 in table 1) from Aron, et. al (2011), U.S. (2 equation state space) from Duca, Muellbauer, and Murphy (this paper), and Australia from Muellbauer and Williams (2011).