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How Important is Wealth for Explaining Household Consumption Over the Recent Crisis? An Empirical Study for the United States, Japan and the Euro Area

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ABSTRACT/RÉSUMÉ

How Important is Wealth for Explaining Household Consumption Over the Recent Crisis? An Empirical Study for the United States, Japan and the euro area

This paper provides new empirical results linking financial and housing wealth to household consumption for the United States, Japan and the euro area. The results suggest that there are important cross-country differences in how wealth, especially housing wealth, affects consumption. They further demonstrate that it can be important to take into account wealth effects on consumption in short-term forecasting exercises, a point which is particularly well illustrated in relation to the recent economic crisis. In addition, conditional projections underline the importance of asset price developments and wealth in determining US savings over the medium term.

JEL classification: D12; C22; E21; E44
Keywords: Housing wealth; financial wealth; wealth effects; consumption equation

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Quelle est l'importance du patrimoine pour expliquer la consommation des ménages au cours de la crise récente ? Une étude empirique appliquée aux États-Unis, au Japon et à la zone euro

Cette étude fournit de nouveaux résultats empiriques établissant un lien entre la consommation des ménages et leur patrimoine financier et immobilier aux États-Unis, au Japon et pour la zone euro. Les résultats suggèrent que l'importance des effets-richesse dans la consommation varie sensiblement entre les pays, en particulier en ce qui concerne la richesse immobilière. Ils montrent également qu'il peut être important de prendre en compte ces effets-richesse dans les exercices de prévision de la consommation à court terme, ce que la période de crise récente permet d'illustrer. De plus, des projections conditionnelles soulignent l'importance de l'évolution des prix d'actifs et de la richesse dans la détermination de l'épargne aux Etats-Unis sur le moyen terme.

Classification JEL : D12 ; C22 ; E21 ; E44
Mots-Clés : Patrimoine immobilier ; patrimoine financier ; effets richesse ; Équation de consommation

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HOW IMPORTANT IS WEALTH FOR EXPLAINING HOUSEHOLD CONSUMPTION OVER THE RECENT CRISIS? AN EMPIRICAL STUDY FOR THE UNITED STATES, JAPAN AND THE EURO AREA

by

Clovis Kerdrain[1]

1. Introduction and summary of the study

1.1 Scope and objectives of the study

1. Consumption equations are central in most models used for macroeconomic analysis, simulations or forecasts. Recent economic developments have underscored the importance of stocks of wealth in consumption behaviour and considerably increased the statistical information embedded in the data. In this light, this paper proposes new empirical results on consumption behaviour using aggregate time-series evidence for the United States, Japan and the euro area at an aggregated level. These equations allow easy inference of the effects of an increase in wealth stocks and could constitute a basis for further improvement of macroeconometric tools.

2. While theory suggests that wealth should be taken into account to derive consumption, some models use a simplified form, while some others ignore it entirely.[2] The OECD Global Model (Hervé *et al.*, 2010) has a fairly rich specification of wealth, but the link to consumption is through total net worth, and although housing and financial wealth are separately identified, they are assumed to have the same effect on consumption. This paper provides updated equations which allow for more disaggregated effects from different components of wealth. They were estimated over the longest available sample, and changes in financial conditions that households are facing were controlled for when possible with proxy measures for the tightness of credit supply.

3. The estimated effects on consumption of a change in both financial and non-financial wealth are derived from these estimates and impulse response functions are plotted. They allow wealth effects between the three major economic zones within the OECD to be compared, both in terms of the wealth

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1. OECD, Economics Department, Paris, and Institut National de la Statistique et des Études Économiques (INSEE), Paris. The author is grateful to Ane Kathrine Christensen, Jean-Luc Schneider, David Turner, Philip Bagnoli and Niels-Jakob Harbo Hansen for most useful help, comments and discussions on this work, and to Diane Scott for assistance in preparing the document. All views expressed herein represent those of the author and are not necessarily shared by the OECD or member countries.

2. For example, the MESANGE model of the French economy described by Klein and Simon (2010) does not include wealth effects. Indeed, the estimated wealth effects were not significant in their work.

elasticity of consumption and in terms of the marginal propensity to consume out of wealth. The importance of considering wealth in forecast exercises is also addressed by comparing the forecasting errors over the period starting in 2006Q1 with an alternative forecast for which the wealth-income ratio is kept constant. Particular attention is paid to the case of the United States, where the consumption equation is used to derive the possible path of US savings over the medium term, conditional on the path of asset prices. This exercise allows the size of wealth effects on medium-term savings to be gauged.

1.2 Principal findings and contributions

4. This study builds on previous ones, in particular using recent economic developments to better identify the size of wealth effects on consumption. Using the appropriate control variables (especially credit supply when possible), functional form, and keeping the results cross-country comparable, it sheds new light on these effects. The estimation of short-term dynamics of consumption allows reasonable inference on the size of financial and real estate wealth effects, both in the short and in the long-run. Such empirical analysis on the euro area as a whole completes the few previous studies on these issues in the euro area. Significant effects are found for Japan, the United States and the euro area both for housing and financial wealth. Specifically, cointegration between wealth, consumption, income and some necessary controls is found. Complete consumption equations are estimated in order to allow fine inference on the short-term dynamics of consumption, and prove rather robust, noticeably in the recent crisis. The main findings detailed in the text are the following:

- First, the magnitude of the long-run marginal propensity to consume (MPC) out of financial wealth is very similar for the United States, Japan, and the euro area, namely about 5 to 6 cents from each dollar increase in wealth.

- Second, housing wealth effects are much larger in the United States than elsewhere at about 5 cents to the dollar, while they seem to be as low as 1 to 1.5 cents to the dollar in the euro area and Japan.

- Third, mainly because of the differential effects from housing wealth, overall wealth effects, as measured by MPCs, are higher in the United States (about 5 cents to the dollar) than in the euro area and Japan (3 to 4 cents to the dollar).

5. These effects are important in understanding the evolution of consumption over the recent crisis period. Had the wealth-income ratio been kept constant from 2006 onwards, important forecast errors would have occurred, especially in the United States. With the equations presented here, predicted consumption tracks actual consumption well throughout the crisis. This importance is also illustrated for longer term exercises by gauging the impact of asset prices on US savings. Changing assumed real asset price growth rates in the United States from 0% to 2% shifts the US saving ratio downwards by about 1.5 percentage points. The impact of credit conditions is also assessed.

6. Together with these main findings, it is worth mentioning some less central results yielded by the present empirical analysis as regards non-wealth determinants of consumption as well as its short-run dynamics. First, the unemployment rate is not found to have significant effects in the long-run. Significant effects in the short-run are found only for the United States, where a rise in unemployment reduces the growth rate of consumption, possibly reflecting precautionary savings. Second, the effects of inflation on consumption differ across countries. In the euro area, inflation seems to reduce consumption in the long run, but has no clear effect in the short run. Conversely, inflation is not found to have a significant long-run effect on consumption in the United States and in Japan, but appears to influence short-term dynamics of consumption in both countries. However, it does not seem to matter anymore for Japanese consumption when it is actually *deflation*. Third, the adjustments towards long-run equilibrium could be as much as two times faster in Japan than in the euro area and the United States. Fourth, in all countries at study, the

adjustment towards long-run equilibrium of consumption shows some signs of asymmetry. It might be faster when consumption is above equilibrium, than when it is below equilibrium. Theoretical understanding of such features is, however, beyond the scope of this paper.

7. The rest of the paper is organised as follows: this section next presents the main issues, together with a short literature review. Section two, three and four then present the empirical results for the United States, Japan and the euro area respectively. Section five derives the impulse response functions of consumption to an increase in wealth for each of them in terms of elasticities and MPCs for financial, housing and total wealth. Section six compares the forecasting capabilities of the estimated consumption equations with and without taking wealth developments into account after 2006, and finds that the contribution from changes in wealth is large. Section seven focuses on longer run wealth effects-related issues, where the estimated equations are used to project the future outcome of US savings given alternative assumptions on asset prices. All the notations used in this paper are defined in Table A1 in the annex.

1.3 Some elements of consumption behaviour theory

8. Life-cycle consumption theory generally relies on intertemporal optimisation. Blanchard (1985), for instance, uses a “perpetual youth” model to derive the links between consumption, income and wealth. He shows that aggregate consumption should be roughly proportional to the sum of current wealth and expected future non-property income.[3] The involved factor of proportionality is the Marginal Propensity to Consume (MPC) out of wealth. Thus, a 1% increase in both wealth and non-property income is expected to raise consumption by 1%. This adjustment is not immediate given numerous sources of short-term dynamics (*e.g.* perceived permanence of wealth change, consumer habits, *etc.*). Basically, the consumer uses all information at his disposal to plan his consumption and smooth it throughout his entire life. The ability to save and borrow is essential for the achievement of this inter-temporal programme. Where households face uninsurable income risks, liquidity constraints or hold illiquid assets, smoothing consumption would be impeded -- particularly in the face of adverse events. Since these constraints will vary over time for a given cohort, different kinds of agents[4] will have different MPCs, as noted by Blanchard (1985). As they are likely to have different kinds of wealth as well, marginal propensities to consume out of each wealth stock at the aggregated level should be different. This effect is due to the fact that financial assets, housing wealth and debt are not in the same hands, thus making further improvements of consumption theory difficult. Empirical analysis must be involved to go further.

9. Current real disposable (non-property) income should generally be expected to remain a constant share of expected future (non-property) income, as noted by Davis and Palumbo (2001) or Davis (2010), so that a 1% increase in current non-property income is equivalent to a 1% increase in expected future non-property income. While this is not true for a given cohort of agents, this result should be expected at the macro-level if the age-distribution of aggregate wealth in the economy is stable. In the presented models, (log) disposable income, possibly augmented by survey data on expectations, is then expected to capture the (log) expected future income effect on consumption.

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3. More precisely, “expected future non-property income” refers to the flow of current and discounted expected future non-property income. This is sometimes referred to as “human wealth” as opposed to asset-based wealth -non-human wealth. The sum of human and non-human wealth is generally thought to be the economically meaningful measure of the “wealth” and the purchasing power of agents in standard life-cycle consumption theory.

4. Differences in income, wealth, culture or education, risk aversion, access to credit etc. are likely to matter. Age or, more precisely, life-expectancy is also widely thought to affect the MPCs, as shown in Blanchard (1985).

10. Many time-series empirical studies estimate aggregate wealth effects on US consumption. However, several fail to constrain properly the model so that if both the wealth and the non-property disposable income of the agents increase by 1%, then consumption also increases by 1% in the long-run *ceteris paribus*, as was noted earlier. The summary of Muellbauer (2007) underlines the importance of such constraint and proposes an appropriate functional form based on Taylor approximations detailed in Muellbauer and Lattimore (1995). This form is used throughout the present paper. It basically links log-consumption as a share of non-property income to the un-logged ratios of wealth to non-property income. The theoretical advantages of this form are twofold: first, it satisfies the constraint detailed above and, second, one can expect it to better fit the behaviour of households holding small stocks of wealth, whereas for these households a logarithmic form would imply that a small change in wealth could generate a large percentage change in their consumption.

Box 1. Estimated coefficients, elasticities and marginal propensities to consume

It is not always straightforward to assess the magnitude of the elasticity of consumption to wealth or the marginal propensity to consume (MPC) out of wealth from the estimation output. The interpretation of the estimated coefficients depends on the specification of the equation. When log consumption per capita (*c/pop*) is linked to log wealth (*w*) and non-property income (*y*), say, by $\ln(\frac{c}{pop}) = \hat{\theta}\ln(\frac{w}{y}) + \ldots \ln\left(\frac{c}{pop}\right) = \hat{\theta}\ln\left(\frac{w}{y}\right) + \cdots$ (*log* specification) the estimated parameter $\hat{\theta}$ is an elasticity because $\frac{dc}{c} = \hat{\theta}\frac{dw}{w}\,\frac{dc}{c} = \hat{\theta}\frac{dw}{w}$. The effect on consumption of one dollar increase in wealth (the value of *dc* when *dw*=1), which is the MPC, is not directly measured by the estimator, since $\frac{dc}{c} = \hat{\theta}\frac{dw}{w} \Rightarrow dc = \hat{\theta}\frac{c}{w}$. In this case one would use a reasonable average of the consumption-wealth ratio *c/w* to derive the direct effect *dc* on consumption, for instance, by using the last years of the sample.

When log consumption is linked to the wealth-income ratio, say $\ln(\frac{c}{pop}) = \hat{\theta}\frac{w}{y} + \ldots \ln\left(\frac{c}{pop}\right) = \hat{\theta}\frac{w}{y} + \cdots$ (*semi-log* specification) then the estimated parameter is much closer to the MPC. Indeed, as $\frac{dc}{c} = \hat{\theta}\frac{dw}{y} \Rightarrow dc = \hat{\theta}\frac{c}{y}$ the correction term *c/y* to get the MPC is never far from one. In this case one would also use a reasonable average of the consumption-income ratio to derive the MPC *dc* from $\hat{\theta}$.

11. Muellbauer (2007) relies on a derivation where property income is excluded from the analysis -- since it is related to the stock of wealth, it would lead to double counting of some wealth effects. One could also note that this allows easy inference to be made on the effects of an increase in wealth on consumption if the stocks of wealth alone are present in the equation.[5] One key feature of the work of Muellbauer (2007) is that is stresses that expectations and financial conditions, especially credit supply, ought to be controlled when running regressions over a long time horizon. This point is particularly

5. Indeed, many studies on the effects of wealth control for total disposable income without subtracting property income from it. In such a setup, wealth effects come through both income (through property income) and wealth stocks, rendering the estimation of *overall* wealth effects particularly difficult. In this case, estimated coefficients on wealth regressors are not directly comparable to those of the present paper. As Benjamin *et al.* (2004) put it, "including labour income and property income together can possibly confuse the propensities to consume out of human and property wealth".

relevant for the United States, where the mortgage credit market has grown considerably since the years 1970s. Housing wealth has indeed become increasingly "liquid" in recent decades, allowing for house owners to increase their consumption, thanks to "second-mortgage" loans and, more recently, home-equity credit lines which represent potential borrowing until the homeowner draws on the line (see Kindleberger, 2005). Such collateralised credits give homeowners easy and adaptable access to cheaper money, as the availability of a collateral tends to drive the cost of borrowing down. Accordingly, these developments should be expected to increase the "spendability" of US housing wealth, both over time and relative to other countries.[6]

1.4 Some elements of recent literature review on wealth effects on consumption

12. There have been numerous studies on the links between wealth and consumption, using various methods and datasets. Most were done on US data or for a panel of OECD countries, but only a few estimate wealth effects in the euro area as a whole, probably due to lack of data. However, many fail to control properly for key factors (interest rates or financial conditions for instance) and/or use an inappropriate functional form or income aggregate, including property income. This may explain the large range of wealth effects estimates that can be found in the literature.

13. Boone *et al.* (2001) also separate housing and non-housing wealth for several countries. They use an error-correction model (ECM) approach similar to that employed here. They also use the same functional form for the equation, with consumption and income coming in logs and wealth-income ratios in level. Their focus is on financial deregulation, so that they share the motivation underlying Muellbauer (2007) that financial conditions are important for consumption and wealth effects. However, their empirical results are not in line with what could be expected *à priori*, as the United States shows lower financial and housing wealth effects than most other G7 countries, especially Japan and Canada. Catte *et al.* (2004) also undertake a cross-country study, and underscore the heterogeneity of wealth effects across countries. The methodology is again comparable to the one used in the present paper, although they use the Stock-Watson (1993) Dynamic OLS (DOLS) method (which is tested here and reported in Tables A2 to A4 in the annex) and put all variables in logs. They also control for potential long-term effects of inflation, unemployment and interest rates, and argue for the use of non-property income in the model, as in the estimation here, but still do not control for financial conditions. Table 2 of Catte *et al.* (2004) reports long-run MPCs of 1 and 5 cents for housing wealth in Japan and the United States respectively, and of 7 and 3 cents for financial wealth. Their empirical results appear relatively robust and still seem to rank financial wealth effects significantly higher in Japan than in the United States. Their results for the euro area countries appear less robust, probably due to data quality issues. Belsky and Prakken (2004) use US macro data and an ECM specification where variables enter only in levels as a ratio to labour income. Though their specification lacks many controls and differs from the one of the present paper in several ways, their results nonetheless appear in line with those presented here. Specifically, they provide a couple of impulse response functions on their Chart 13 for US wealth effects that are comparable to those reported here in Figures 12 and 14 in terms of magnitude. Benjamin *et al.* (2004) used US macro data and, unlike the results below, found no cointegration between variables, defined relative to total disposable income -- possibly due to lack of relevant controls used here. Accordingly, they estimated a model in differences and report consumption effects of a one dollar increase in wealth of 8 cents for real estate and 2 cents for non-tangibles.

6. A higher consumption is also expected from first-time home buyers when credit conditions ease and when there is large supply for mortgage credit. As stressed in Muellbauer (2007), a large credit supply drives downpayments down. Young people then anticipate less interest payments and much smaller initial need of (saved) money. In that perspective, they will save less and, again, "smooth" more their consumption as far as their future-income expectations are correct, raising their consumption relative to income.

14. More recently, Slacalek (2006) used time series for 16 countries to study wealth effects. Underscoring strong cross-country heterogeneity, he noted weaker wealth effects in the euro area. He also reports lower housing wealth effects than financial wealth effects in most countries, but not in the United States and the United Kingdom. Table 5 reports numerous estimated long-run MPCs. For the United States and Japan, they are 7 and 6.3 cents, respectively, in a dollar of housing wealth and 5.3 and 9.5 cents, respectively, in the dollar of financial wealth. His results for the euro area countries are rather mixed: financial wealth effects are estimated three times higher in Germany than in the United States, while housing wealth effects in Italy are estimated negative. Interestingly, he uses pooled estimation for the whole euro area, finds that financial wealth effects could be a bit below 2 cents to the dollar, but has difficulties finding a significant result for housing wealth.[7] The estimation technique is comparable to the one in Caroll *et al.* (2006) who use US macro data in differences and find stock market wealth effects are lower than housing wealth effects after controlling for unemployment and interest rate-related effects (about 9 cents to the dollar for housing wealth and half that for stock market wealth). Case *et al.* (2006) also found larger wealth effects for real estate using a panel of US states, but their study lacks several controls, as argued by Muellbauer (2007). Conversely, Calomiris *et al.* (2009) suggest most of these results have largely overestimated housing wealth effects due to endogeneity biases. Their study re-uses the dataset of Case *et al.* (2006) and finds that housing wealth effects in the United States, if any, are likely to be smaller than stock market wealth effects. Comparable studies for the euro area considered as a whole have been relatively rare until recently. Apart from Slacalek (2006), both Sousa (2009) and Skudelny (2009) considered wealth effects in the euro area as a whole. They both use euro area-wide aggregated macro data. Skudelny (2009) runs regressions with un-logged data and finds the MPC out of financial wealth in the euro area could range from 2.4 to 3.6 cents to the dollar and could be a bit below 1 cent to the dollar for housing wealth. Sousa (2009) finds smaller results, and concludes housing wealth effects in the euro area are virtually nil. However, neither Sousa (2009) nor Skudelny (2009) tries to control for other potential macroeconomic determinants of consumption like inflation or interest rates.

15. The study of Muellbauer (2007) discusses most of these previous empirical findings, emphasises their lack of controls and tries to control for the massive change in financial conditions faced by households that occurred over the last 30 years. Using the same semi-log[8] specification as Boone *et al.* (2001) that is adopted in the present paper, he found the US wealth effects are likely to be about 10 cents to the dollar for liquid assets, a bit less than 2 cents for illiquid ones, and from 6 to 7 cents for housing wealth. This is consistent to what is found here, with differences of methodology probably explaining remaining differences. Finally, Davis (2010) gives a concise, but rather complete overview of this literature which serves to emphasise the wide range of results in this area. The literature review presented in this section is summarised in Table 1. Considerable variations can be found from one study to another, motivating the numerous robustness checks provided throughout the present paper and especially in the annex.

7. He finds significant housing wealth effects only in a sub-sample of his dataset.

8. Relating the log of the consumption-income ratio to the un-logged wealth-income ratio.

Table 1. Summary of some results from the literature

Paper	Main findings on wealth effects	Additional information
Boone *et al.* (2001)	Find positive housing and financial wealth effects, both estimated to be higher in Japan and Canada than in the United Kingdom or the United States. Estimated coefficients range from 0.04 (US) to 0.12 (Japan) for financial wealth and -0.06 (Italy) to as much as 0.34 (Japan) for housing wealth.	ECM, semi-log specification
Catte *et al.* (2004)	Report long-run MPCs of 1 and 5 cents for housing wealth in Japan and the United States respectively, and of 7 and 3 cents for financial wealth respectively. Smaller effects (one to two cents) are found for main continental European economies.	ECM, log specification
Belsky and Prakken (2004)	Provide a couple of impulse response functions of US consumption to US wealth that are somewhat comparable to Figures 12 and 14 of this paper in terms of magnitude (a bit above 5 cents to the dollar in the long-run)	ECM, specification with levels of the ratio of the variables to labour income, lack of control variables
Benjamin *et al.* (2004)	Estimate consumption effects of a one dollar increase in wealth in the United States are 8 cents for real estate assets and 2 cents for non-tangibles.	Find no cointegration linking consumption and wealth. Data used relative to personal disposable income
Slacalek (2006)	Reports generally lower housing wealth effects than financial wealth effects in most countries, though not in the United States and the United Kingdom. Estimated long-run MPC out of one dollar of wealth are 7 and 6.3 cents for housing wealth, respectively, for the United States and Japan, and of 5.3 and 9.5 cents, respectively, for financial wealth. Has difficulties finding positive housing wealth effects in the euro area.	Specification with data in difference. Underscores strong cross-country heterogeneity of wealth effects. Note weaker effect in the euro area.
Caroll *et al.* (2006)	Find stock market wealth effects are lower than housing wealth effects in the United States (about 9 cents to the dollar for housing wealth and half that for stock market wealth)	Specification with data in difference.
Case *et al.* (2006)	Find significant housing wealth effects in the United States and little evidence of stock market wealth effects	Use a panel of US States
Calomiris *et al.* (2009)	Find that housing wealth effects in the United States, if any, are likely to be smaller than stock market wealth effects. Argue that previous empirical results may have overestimated them.	Re-use the dataset of Case *et al.* (2006)
Skudelny (2009)	Finds MPCs of 2.4 to 3.6 cents per dollar of financial wealth, and of 0.7 to 0.9 cents per dollar of housing wealth in the euro area.	Uses euro area aggregate data. Specification without logs.
Sousa (2009)	Finds long-run MPCs about 2 cents per dollar of financial wealth but no significant housing wealth effects in the euro area.	Uses euro area aggregate data, log-specification.
Muellbauer (2007)	Emphasises the lack of controls in many of the empirical results in the literature, especially as regards the credit availability in the United States. Finds that US wealth effects are likely to be about 10 cents to the dollar for liquid assets, a bit less than 2 cents for illiquid ones, and from 6 to 7 cents for housing wealth.	Semi-log specification
Davis (2010)	Gives another concise overview of this literature and reports some other empirical results for MPCs out of wealth in his Table 2.	

Source: Compilation by the author.

2. Wealth effects in US consumer behaviour: the importance of housing and financial wealth

2.1 Some details about the dataset

16. The quarterly time series refer to the personal sector, which is made up of households and non-profit institutions serving households (NPISH). Real variables were deflated by the private consumption deflator. All the wealth aggregates (real estate and financial) were taken from the Federal Reserve Board (FRB) Flow of Funds accounts (FoF), Table B.100. Surveys of consumer expectations (and the Index of Consumer Sentiment) are from the survey of consumers conducted by the University of Michigan and Thompson Reuters. Stock market data are from Wren Investment Advisers.[9] The various consumer loan interest rates are from the FRB (when available monthly, the quarter-average is kept as the quarterly value). Demographic annual variables are from the US Bureau of the Census. Age share variables are linearly interpolated from annuals (yearly data stored in Q2) while the population series is interpolated from the quadratic algorithm of the Eviews software, yearly averages matching the annual values. The construction of the credit supply index is detailed further. The rest of the data come from the *OECD Economic Outlook* 86 database. The definition of US household savings is consistent with the *OECD Economic Outlook* and refers to the personal saving item in the US National Income and Product Accounts (NIPA) published by the Bureau of Economic Analysis (BEA). Note that FoF includes consumer durables in total tangible assets held by the personal sector, whereas they already are included in consumption. Hence, one should either take them out of consumption, add them to savings and eventually to wealth stocks, or else exclude them from wealth stocks. This latter solution was chosen. The non-property income was built for the United States by subtracting property income from total disposable income.[10]

17. The credit conditions (*CC*) index is inspired by the work of Muellbauer (2007). A reasonable index reflecting the supply side of the credit market can be built using the answers to the Senior Loan Officer Opinion Survey on Bank Lending Practices conducted by the Federal Reserve. It is based on the cumulative value of the net percentage of domestic banks reporting "more willingness to make consumer instalment loans", adjusted for its trend,[11] and normalised so that its highest value is one (in 2006Q3).

2.2 The long-run determinants of US consumption

2.2.1 Estimation

18. An error-correction model equation (ECM) is estimated to fit US consumption behaviour.[12] For the sake of simplicity, the two-step Engle-Granger (1987) method[13] was used. A particular emphasis is put

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9. http://www.wrenresearch.com.au/downloads/ consulted 1 December 2009.

10. These two flows exclude consumer debt service, "personal interest payments" item in the NIPA, which is non-mortgage interest paid by households. Their difference is then the same as if interest was first subtracted from both.

11. More precisely, noting η_t the series of Net percentage of Domestic Banks indicating more willingness to Make Consumer Instalment Loans, the *CC* index is built according to the recursive equation: $CC_0 = \eta_0$ and $\Delta CC_t = \eta_t - 6$ where the date 0 here refers to the quarter 1966Q3. Then, it is normalized so that its highest value is one.

12. Another difference with the work of Muellbauer (2007) is that the present paper did not introduce a "stochastic trend" in the regression, which interpretation is difficult and makes forecasting impossible.

on econometric checks for the relevance of the specification. First some statistical tests regarding the order of integration of some data[14] are shown: the results of some popular unit root tests, conducted on the longest available sample after 1960 are presented for most series. The tests are run with an intercept (and a time-trend if mentioned). Test statistics falling into the 1% critical value- 10% critical value interval are considered inconclusive and are reported by a blank. Both ADF tests use a Schwarz (1978) criterion (BIC) to select the lag length[15] in the test. Definitions of the variables are given in the annex, see Table A1.

Table 2. Order of integration in the US data, in level and in first difference (Δ)

Series	ADF[1] (level[7])	ADF-GLS[2] (level[7])	KPSS[3] (level[7])	ADF (Δ)	ADF-GLS (Δ)	KPSS (Δ)	Selected order of integration (level)
ln(c/pop)	I(1)	I(1)		I(0)	I(0)	I(0)	I(1)
ln(y/pop)	I(1)	I(1)		I(0)	I(0)	I(0)	I(1)
ln(c/y)	I(1)	I(1)		I(0)	I(0)	I(0)	I(1)
CC	I(1)	I(1)		I(0)	I(0)	I(0)	I(1)
hw/y	I(1)[8]	I(1)[8]	I(0)	I(0)	I(0)	I(0)	I(1)[6]
*CC*hw/y*		I(0)[4]	I(0)	I(0)	I(0)	I(0)	I(1)[6]
ofa/y	I(1)	I(1)	I(1)	I(0)	I(0)	I(0)	I(1)
fl/y	I(1)	I(1)	I(1)			I(0)	I(1)
liq/y	I(1)	I(1)	I(1)	I(0)	I(0)[4]	I(0)	I(1)
u	I(1)[5]		I(1)	I(0)	I(0)	I(0)	I(1)[6]
r_{new}	I(1)		I(1)	I(0)	I(0)	I(0)	I(1)[6]
r_{48}	I(1)	I(1)	I(1)	I(0)	I(0)	I(0)	I(1)[6]
*r*nfwh/y*	I(1)	I(1)[5]		I(0)	I(0)	I(0)	I(1)[6]
ln(nfwh/pop)	I(1)	I(1)	I(1)	I(0)	I(0)	I(0)	I(1)
π	I(1)	I(1)	I(1)	I(0)	I(0)[9]	I(0)	I(1)

1. Augmented Dickey-Fuller Test, see Dickey and Fuller (1979) and Said and Dickey (1984). Critical values from MacKinnon (1996).
2. Quasi-optimal ADF Test, see DF-GLS in Elliott et al. (1996), Critical values from MacKinnon (1996).
3. Test for stationarity proposed by Kwiatkowski et al. (1992). Asymptotic Critical values from Kwiatkowski *et al.* (1992).
4. The test for non-stationarity using the AIC criterion fell between the 5% and the 10% critical values.
5. Only AIC brought a conclusive statistic.
6. The correlogram of the differenced series did not show any sign of over-differentiation.
7. A time-trend was introduced to run the tests on the variables in level.
8. The test for non-stationarity using the AIC criterion fell between the 1% and the 5% critical values.
9. AIC indicates I(1).

Source: Author's calculations.

13. Another advantage over the Stock-Watson (1993) method is that not putting leads of the variables in the regression allows taking the last observations into account, for they are assumed rather informative.
14. Not all the regressors that were tried are listed and tested for unit roots.
15. When test results are not robust to the use of the Akaike (1973) criterion AIC instead, it is signaled. AIC will select more lags in the ADF regressions than BIC. Precisely, it is reported when the test with BIC was non-conclusive or if it brings a result on the other side of the 5% critical value.

19. First, the long-run drivers of consumption are examined by estimating a simple cointegration relation between consumption, wealth and income. The functional form used here ensures that any 1% increase both in wealth stocks and non-property income results in a 1% increase in consumption in the long run. It looks much like that adopted by Boone *et al.* (2001) but differs from the one in Muellbauer (2007) because the focus is here on the long term, while Muellbauer (2007) directly estimates dynamic equations in differences.

20. *A-priori*, one would expect the credit index to either appear alone (effects on first-time buyers) or else interact with the housing wealth (effects on home-owners). However, it cannot appear in both forms given the risk of colinearity (the correlation between *CC* and *CC*hw/y* is more than 0.96). As shown in regression (B) compared to the benchmark, (Table 3), the housing wealth effect appears to come mainly through the interacted *CC*hw/y* variable so that this latter was preferred over just *CC*. Note, however, that regression (E) could constitute a reasonable alternative to the benchmark. Introducing the long-term real interest rate is necessary to have cointegration [see the diagnostics of regression (A)]. This could explain why the cointegration tests failed in Benjamin *et al.* (2004). When controlling for the income effect of a rise in the long-term yield, a substitution effect becomes evident in regression (C) through a negative coefficient on the interest rate -- though it remains insignificant. Regression (D) shows that the benchmark incorporates almost all of the relevant control variables. It also shows that the coefficients on the main variables are quite robust across specifications, which is confirmed by further empirical analysis not reported here. Regression (F) can be compared to regression (E) and to the benchmark to assess the magnitude of the bias introduced by not controlling for the change in credit conditions.[16] Note that one also loses the cointegration (see for instance the ADF and DW statistics) relation in (F) which indicates the results in (F) may be "spurious" (see Phillips, 1986).

21. Among other specifications, fiscal variables were tried to capture Ricardian effects, and did not show up with significant or robust coefficients. Some household survey variables from the SCA (Reuters/University of Michigan -- Survey of consumers) were also tested, in order to better capture income growth expectations, but proved to be unsuccessful.[17] The theoretical MPCs should also be related to demographics through ageing. It is then tempting to try to capture demographic trends in consumption. However, demographic data are smooth trends and carry very little information. Putting them in a regression is likely to capture other forms of trend. Indeed, demographic age shares variables proved to be not robust in this setup.[18]

16. Note that the regression (F) indeed gives point estimates that are closer to what can be found, for instance, in the work of Caroll *et al.* (2006), among others.

17. More precisely, the balancing items of the answers to questions 6 and 9 of the SCA did not come out with significant coefficients. Question 6 is: "Now looking ahead -- do you think that a year from now you (and your family living there) will be better off financially, worse off, or just about the same as now?". Question 9 is: "How about the next year or two -- do you expect that your (family) income will go up more than prices will go up, about the same, or less than prices will go up?"

18. When significant, they led to extreme forecasts over the medium-run which could not be trusted. All in all, one may recall that some mild evidence of influence of demographic trends on house prices exist, see Lindh and Malmberg (2008) for instance, which could be used to make assumptions on future house prices. The use of demographics in non-stationary regressions is however always problematic, as these extremely-low frequency series are carrying very little information that may be correlated with some other long-run trends in the economy.

Table 3. Selected long-term US regressions[1]

Dependent variable : *ln(c/y)=ln(c/pop)-ln(y/pop)*

Regressors	Benchmark	(A)	(B)	(C)	(D)	(E)	(F)
Cste	-0.117***	-0.218***	-0.133***	-0.112***	-0.179***	-0.187***	-0.28***
	[-4.4]	[-6.8]	[-4.2]	[-2.9]	[-4.6]	[-7.1]	[-5.5]
*CC*hw/y*	0.044***	0.044***	0.040***	0.044***	0.046***		
	[16.0]	[12.3]	[8.8]	[16.5]	[8.3]		
(ofa-fl)/y	0.011***	0.02***	0.012***	0.009	0.015***	0.015***	0.033***
	[3.2]	[4.6]	[3.3]	[1.0]	[3.5]	[5.1]	[5.5]
liq/y	0.125***	0.218***	0.124***	0.123***	0.149***	0.124***	0.120***
	[5.6]	[8.4]	[5.7]	[4.9]	[5.3]	[6.1]	[2.9]
*r*nfwh/y*	0.174***		0.173***	0.223	0.184**	0.135***	0.180***
	[5.2]		[5.4]	[1.0]	[2.3]	[4.7]	[3.3]
hw/y			0.009			0.036***	0.081***
			[0.7]			[5.2]	[6.9]
r				-0.157			
				[-0.2]			
r_{new}					-0.324***[2]		
					[-2.9]		
r_{48}					0.168		
					[0.6]		
u					0.003		
					[1.3]		
π					0.146		
					[0.2]		
$u*\pi$					0.01		
					[0.1]		
r_s					0.08		
					[0.4]		
CC						0.077***	
						[11.2]	
Sample	66Q3-09Q2	66Q3-09Q2	66Q3-09Q2	66Q3-09Q2	72Q1-09Q2	66Q3-09Q2	66Q3-09Q2
R^2	0.93	0.89	0.93	0.93	0.94	0.94	0.83
BIC[3]	-5.650	-5.246	-5.633	-5.621	-5.722	-5.844	-4.824
SER	0.0135	0.0167	0.0135	0.0135	0.0120	0.121	0.0204
DW	0.545	0.267	0.548	0.548	0.641	0.636	0.272
ADF	-5.22	-2.62	-5.24	-5.24	-5.27	-5.79	-3.51

1. Newey-West (1987 and 1994) robust covariance matrix is used to compute t-statistics in brackets. The autocorrelation in the residual is significant, roughly, until the fourth order. This corresponds to the auto-selected Newey-West window. For convenience, "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution. However, note that such criterion does not constitute a formal test per se (see for instance Park and Phillips, 1986) since the regressors should be strictly exogenous to ensure asymptotic normality, which is not necessarily the case. Nevertheless, t-stats still should go to infinity almost-surely when the parameter is non-zero, while the contrary is not true. Anyway, DW and ADF statistics provide formal cointegration tests together with the Johansen (1996) tests.

2. Not statistically significant or with different order of magnitude in a one-step OLS procedure (as in Table A2) or multivariate Maximum Likelihood (Johansen, 1996) estimation. Also less robust across specifications of this same regression.

3. BIC is the Bayesian Information Criterion from Schwarz (1978). DW is the Durbin-Watson statistic and SER the standard error of the regression. ADF is the ADF statistic calculated with the residuals as done by Phillips and Ouliaris (1990) to test no-cointegration

Source: Author's calculations.

22. The use of a Stock-Watson (1993) procedure[19] did not really alter the order of magnitude or the sign of the estimated coefficients. When added in the regression, the coefficient on log income per capita is not significant, which supports the restricted functional form used here (see Table A2). Cointegration tests based on ADF and DW statistics validate the cointegration in the benchmark regression.[20] The Johansen (1996) cointegration tests also validate one and only one cointegration relation.[21] Along with this test, the estimated cointegration by this multivariate Maximum Likelihood method gives somehow different coefficients to the OLS and Stock Watson procedure. The main differences are that the coefficient on liquid assets is half the one in the benchmark, the one on "other financial assets" is not always statistically significant, and the coefficient on *r*nfwh* is higher than in the benchmark, about 0.30. When income is included, its coefficient is zero, validating again the structural form of the equation. The coefficient on housing wealth is always between, roughly 0.035 and about 0.055.

23. Overall and using various procedures (see also the results in Table A2 in the annex) one can see that wealth effects are likely to be significant and positive, but their size is difficult to assess with precision. Housing wealth effects could range from 2.5 to 7 cents increase in consumption for each dollar increase in housing wealth, while financial wealth effects would range from 3 to 9 cents to the dollar. The MPC of liquid assets seem to be higher than for other wealth stocks. The benchmark specification lies somewhere in the middle with about 5 cents to the dollar of housing wealth (provided the credit supply is unconstrained) and comparable financial wealth effects.

2.2.2 Interpretation of the results

24. The results tend to confirm that there is indeed a strong long-term relation between wealth and consumption, although (non-property) income remains its main driver. The need to control for the (property) income effect by introducing the interest rate appeared necessary to have strong cointegration results, especially when interacted with financial wealth. This term could capture a dependence of the MPCs on interest rates, or Keynesian effects not included in life-cycle models (that is, the higher the current income, the higher the consumption), or, more likely, both. Housing wealth and credit conditions also turned out to be necessary in the benchmark regression, while the inclusion of liquid assets and other net assets does not really change the *DW* statistic.

25. As illustrated by Figure 1, consumption as a share of income shows large historical variations. The first period, before 1980, is a period of low real interest rates. The wealth stocks are relatively flat as shares of income (Figure 2) as well as house and stock prices (Figure 3) and the saving rate is high. From 1980 onwards the sudden rise in real interest rates considerably increased the purchasing power of property income, rising consumption to non-property income ratio while the saving rate remained unchanged at over 8%. The easing in credit supply (see Figure 22) starting between 1980 and 1985 and the (possibly related) increase in house prices from 1985 to 1990 led to a rise in consumption, interrupted in 1990 by a fall in house prices. After 1985, disposable income was no longer supported by a rise in interest rates and savings fell from 9% to about 7% in five years. After stabilising for about three years, the saving rate appears to fall again after 1993, which could be attributed to a second sharp rise in credit conditions (house and stock prices were flat before 1995). After this year, the sharp increase in stock market wealth could account for high consumption until 2000, which drove the saving rate below 4% (see Figure 20). After the crash of stock markets and stagnation of credit supply, with falling real interest rates, consumption decreased after

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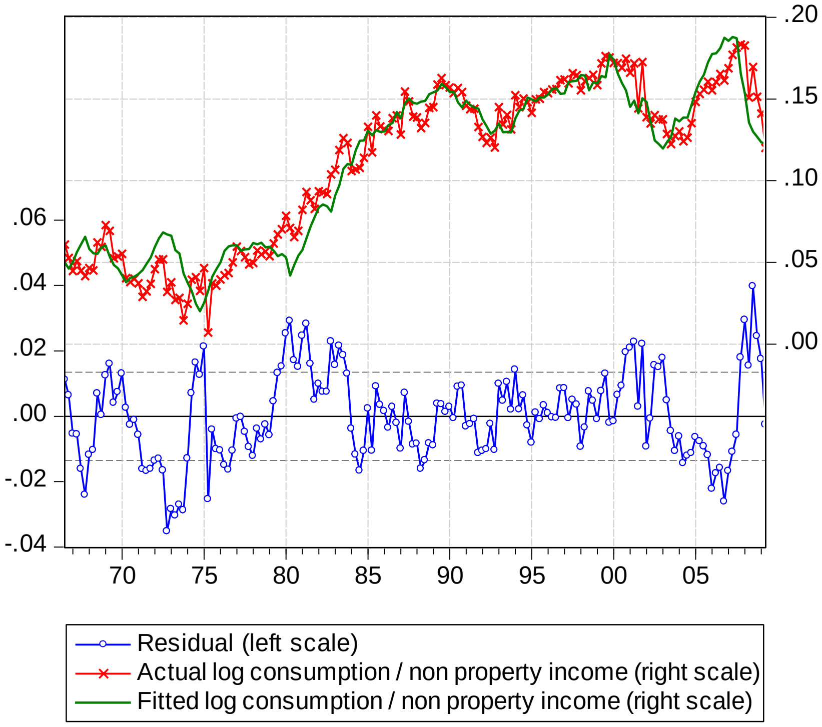
19. Lags and leads of differenced variables were put in the regression until the third order.

20. DW statistics are reported in all tables, and the 1% critical value for no-cointegration test in Engle-Granger (1987) is about 0.5. Asymptotic critical values for the no-cointegration test based on ADF statistic on residuals are reported in Phillips and Ouliaris (1990), table IIb, with n=4 regressors, equal to -4.5 and -5.1 at the 5% and 1% level while the actual ADF statistic is -5.2. No-cointegration is then strongly rejected in the benchmark regression.

21. Maximum Eigenvalue and Trace tests with any number of lags from 1 to 5.

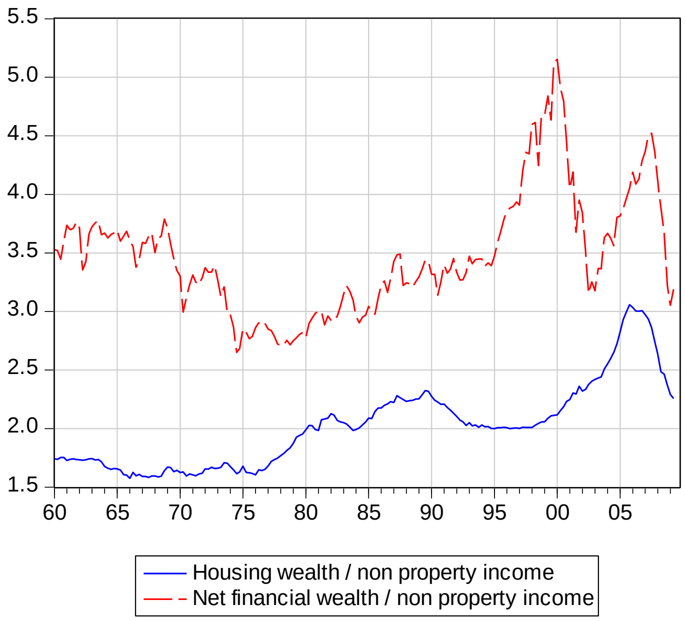
2000. The subsequent rise in housing wealth supported, from 2003 on, by a renewed easing of credit conditions, can account for further increases in consumption. Afterwards, all the drivers of consumption have positive influences: higher housing wealth with easy credit conditions, along with new peaks in stock market prices. Consumption followed with a lag, causing the saving rate to fall to under 2% after 2005. Conversely, consumption decreased during the recent crisis, following the drop in wealth stocks -- caused by the dramatic fall in asset prices -- leading the saving rate, after a lag, back to about 4%.

Figure 1. Residuals of the long-term benchmark regression for the United States



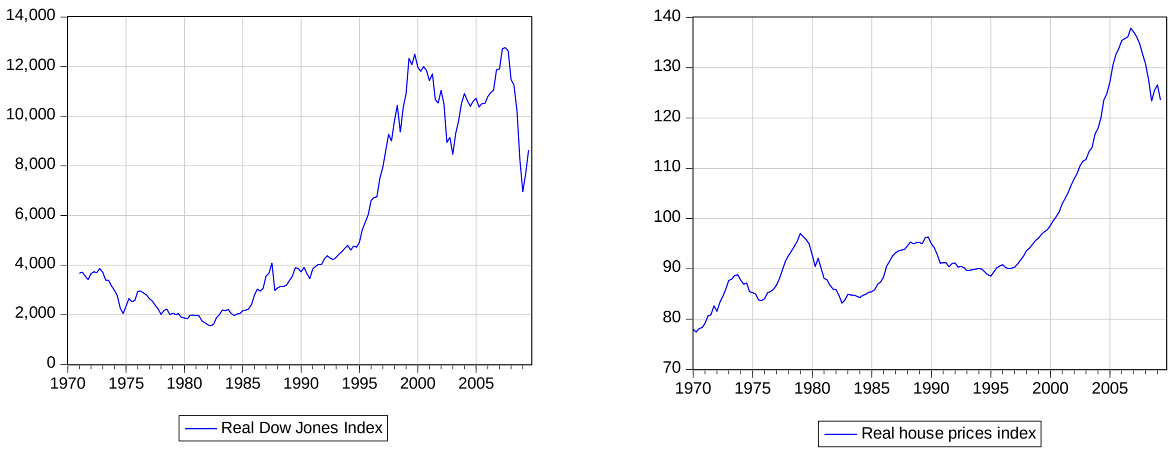
Source: OECD, author's calculations.

Figure 2. Stocks of US household wealth



Source: OECD, Federal Reserve Board.

Figure 3. Asset prices in the United States over historical periods



Source: OECD, Wren Investment Advisors.

2.3 Short-run behaviour of US consumption

26. The short-run equation driving the dynamics of consumption, that is its growth rate, is presented below. An extensive set of à-priori relevant regressors were tried, and a large number of tests such as those reported for the benchmark model were run. The robustness of a small number of parameters was preferred over a large set of regressors. Only few results are reported here with the set of regressors that proved most robust across specifications, samples and statistical tests.

27. Note that the benchmark specification proves robust to many statistical checks. A quantile regression estimating the conditional median (instead of the conditional mean in OLS regressions) brings coefficients similar to the benchmark equation.[22] Indeed, the obtained residuals are so much like the ones from the benchmark that their correlation is over 99.2%. The benchmark regression passes, for instance, various Chow Tests, the CUSUM test (Brown *et al.*, 1975)[23] over the whole available sample; recursive estimates show no sign of significant structural change in the coefficients, and confirm that the null hypothesis is not rejected for any subsample starting in 1972Q2. One can also compare regressions (I) and (J) estimated over different sub-samples with the benchmark model (see Table 4).

28. In the benchmark regression, the error correction term appears to be significant and about -0.10, meaning four years are needed to close 80% of any gap in the long-run relationship. The strong impact of change in income and unemployment may capture the effects of shifts in expectations. For instance, a rise in unemployment will raise precautionary savings as workers may then fear a loss of their own jobs. The interest rates on the car loan market appeared to be very significant, much more than financial market interest rates. However, they may well capture general short-term changes of credit opportunities more than the sole effect of car loans. The lagged change in financial wealth, mainly linked with developments in the stock market, should be seen as a significant driver of expectations, both in future employment and wage and in future change in wealth.

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22. Quantile regressions are, among other properties, less sensitive to extreme observations in the dependant variable than OLS. See for instance Koenker and Hallock (2000).

23. This recursive test uses sub-samples on which the model can be estimated, making the presence of dummies problematic. So the test was done either removing all the dummies, or all the dummies apart from the one of 1991.

Table 4. Selected short-term regressions for the United States[1]

Dependent variable: $\Delta ln(c/pop)$

Regressors	Benchmark	(G)	(H)[10]	(I)[7]	(J)[8]	Regressors	(K)[9]
Cste	0.004***	0.004***	0.005***	0.004***	0.004***	*Cste*	0.005***
	[11.3]	[10.2]	[7.6]	[6.8]	[8.4]		(0.001)
ect(-1)	-0.099***	-0.100***	-0.084**	-0.122***	-0.084*	$ect_{t-1}.1_{ect>0}$	-0.136***
	[-3.1]	[-3.0]	[-2.3]	[-3.2]	[-1.7]		(0.048)
$\Delta n(y/pop)$	0.163***	0.168***	0.177***	0.179***	0.161***	$ect_{t-1}.1_{ect<0}$	-0.069
	[4.4]	[4.5]	[4.5]	[4.1]	[3.6]		(0.043)
$\Delta u/100$	-0.68***	-0.732***	-0.789***	-0.696***	-0.739***	$\Delta ln(y/pop)$	0.162***
	[-3.8]	[-3.9]	[-3.3]	[-3.1]	[-5.4]		(0.038)
Δr_{new}	-0.108***	-0.149***	-0.129***	-0.123*	-0.1***	$\Delta u/100.1_{u>6}$	-0.615***
	[-2.7]	[-3.2]	[-3.2]	[-1.7]	[-3.2]		(0.202)
Δr_{48}	-0.401***	-0.366***	-0.362**	-0.324***	-0.32**	$\Delta u/100.1_{u<6}$	-0.797**
	[-3.8]	[-3.6]	[-2.2]	[-2.7]	[-2.2]		(0.342)
Wealth change[2]	0.020***	0.019***	0.019***	0.027*	0.015**	Δr_{new}	-0.105***
	[3.0]	[2.7]	[2.7]	[1.9]	[2.2]		(0.04)
$\Delta \pi$	-0.837***	-0.779***	-0.788***	-0.808***	-0.514***	Δr_{48}	-0.404***
	[-6.1]	[-5.5]	[-3.4]	[-4.3]	[-2.7]		(0.111)
$\Delta ln(c/pop)(-1)$			-0.133			*Wealth change*[2]	0.018**
			[-1.6]				(0.007)
$\Delta \pi(-1)$			0.014			$\Delta \pi.1_{\Delta \pi>0}$	-0.937***
			[0.2]				(0.173)
$\Delta ln(y/pop)(-1)$[6]			0.101**			$\Delta \pi.1_{\Delta \pi<0}$	-0.699***
			[2.2]				(0.176)
Public deficit[5]			-0.005				
			[-0.6]				
$\Delta u(-1)/100$			-0.026				
			[-0.2]				
$\Delta_4 r$			0.04				
			[0.8]				
$\Delta_4 r_s$			-0.039				
			[-1.3]				
$\Delta(hw/y)(-1)$			0.021				
			[1.4]				
$\Delta_4 ln(rp_{house})$			-0.006				
			[-0.3]				
Set of dummies[3]	yes	no	yes	no	yes	*Set of dummies*[3]	yes
Sample	72Q2-09Q3	72Q2-09Q3	72Q2-09Q2	72Q2-94Q3	87Q2-09Q3		72Q2-09Q2
R^2	0.65	0.58	0.68	0.62	0.65		0.66
BIC	-7.805	-7.717	-7.568	-7.425	-8.144		-7.719
SER	0.0042	0.0046	0.0042	0.0051	0.0033		0.0042
DW	2.084	1.978	1.973	2.111	1.697		2.104
Q-test[4]	0.199	0.047	0.455	0.412	0.001		0.181
Q²-test[4]	0.066	0.445	0.001	0.661	0.014		0.211
Jarque-Bera[4]	0.19	0.52	0.29	0.84	0.05		0.11
Chow F test 1990[4]	0.274	0.056	0.056	-	-		0.276
Chow F test 1985 1997	0.215	0.029	0.040	-	-		0.216

1. White (1980) robust covariance matrix is used to compute robust t-statistics in brackets. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.
2. "Wealth change" stands for the growth rate of per-capita real net financial wealth over a semester ending in the quarter *t-1*: *ln(nfwh(-1)/ nfwh(-3)) - ln(pop(-1)/pop(-3)).*
3. 2001Q4 and 2008Q4 observations were dummied out and a dummy equal to 1 between 1990Q4 and 1991Q4 was also included.
4. "Q-tests at lag k" for no-autocorrelation in the residual up to lag k are run for all *k*<30. The table reports the lowest p-value for all k. The "Q²-tests" reports the same for the squared residuals. "Jarque-Bera" reports the p-value of the Jarque-Bera test for normality of the residuals. The Chow tests detect structural breaks. They indicate the p-value of the test for no significant difference in the values of the parameters estimated over the subsamples separated by the indicated year(s).
5. *ln(debt(-1)/debt(-5))* where *debt* is the gross public debt in GDP points.
6. Not significant alone in addition to the benchmark.
7. Regression (I) covers the first observations (60%) of the whole sample while regression (J) covers the last observations (60%).
8. Newey-West (autocorrelation robust) t-stats instead of White t-stats for this regression.
9. In parentheses: robust standard errors.
10. P-value for the joint nullity F-test of the non-benchmark regressors: 0.42.

Source: Author's calculations.

29. Note the strongly significant negative effect of inflation on consumption. In theory, inflation can curb consumption by several channels. First, a hike in inflation may imply higher purchases made in anticipation of further increase in prices. It also implies a lower purchasing power and could hence cause lower consumption, which seems to be dominant here. However, this coefficient may rather capture the fact that households do not react to current Δr_{new} and Δr_{48} real interest rates changes, but rather, to the nominal changes.[24] A more accurate answer cannot be identified in this setup.

30. A set of dummies has also been included in the regression [compare regression (G) with the benchmark].[25] Other effects, like Ricardian savings when public spending raise, did not prove robust or significant. Regression (H) shows a mainly random example of candidate regressors that did not turn out to explain consumption growth well.

31. Regression (K) presents some answers about whether non-linear asymmetric effects should be included in the equation. Interestingly enough, the error correction seem to be stronger downward than upward. That is, decreases in wealth are likely to induce more rapid adjustment of consumption than do increases. The effect of rising unemployment may have been more pronounced when unemployment was not already high ("high" unemployment rate threshold was put at 6%). Increasing inflation may have a slightly higher effect on consumption than decreasing inflation, perhaps because of risk aversion or perhaps because it is covered differently by economic news. However, all in all, non-linearities did not prove large and robust enough to be kept in the benchmark.

32. The benchmark specification captures several negative blips in consumption as in 1974Q4 and 1980Q2 reasonably well. The recent fall in the consumption growth rate since 2007 is also captured (though with a lag) and so is the subsequent recovery. More generally, this specification always passes the N-step forecast recursive test.[26] Figure 4 shows the out-of-sample forecast of consumption per capita growth from 2004Q1 (conditional on the actual realisations of explanatory variables) through the crisis, using the benchmark specification.[27]

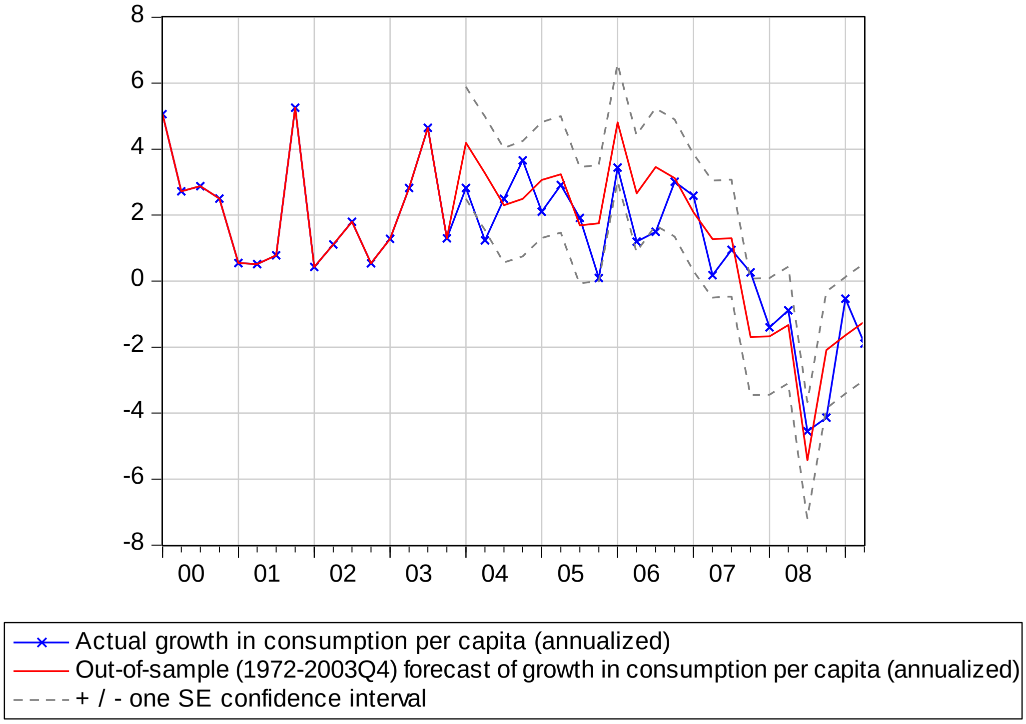
24. Recall private consumption deflator is used to calculate real interest rates.

25. The inclusion of a dummy for 2008Q4 was done to capture a large shock that could not be fully explained by this simple linear model, and hence could damage the estimation. However, all the conclusions are unchanged when this dummy is ignored.

26. N-Step forecast test: For each possible subsample starting from the first observation, the equation is estimated and used in out-of-sample forecast until the end of the whole dataset. If differences with the actual observations are statistically significant, the F-test should reject the null, indicating a likely structural change. The N-step forecast computes such F-tests recursively for all possible subsamples.

27. The 2008Q4 dummy included in the benchmark is completely ignored here, given that the estimation sample for this exercise stops in 2003.

Figure 4. United States: out-of-sample forecast over the crisis period



Source: Author's calculations.

3. Wealth effects in Japan: small effects from housing, but large financial wealth effects

3.1 Some properties of the data

33. Similar regressions and tests are run on Japanese data to those already described with US data. All the Japanese data used here come from the OECD Economics Department Analytical Database[28] apart from the lending survey data. The latter uses the long times series of the Tankan survey reported by the Bank of Japan, namely the business opinion survey concerning lending attitudes of financial institutions.[29] It was used to build a proxy for the tightness of credit supply. Non-property income is built by subtracting net property income from disposable income.

34. The order of integration in Japanese data is investigated in Table 5. The tests are run with an intercept (and a time-trend if mentioned). Overall, it seemed necessary to drop the pre-1976 part of the sample for some data in order to avoid large trend breaks. The long-run relationship between consumption, wealth and income is investigated with the same restriction as for the United States.[30]

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28. Version: 29 January 2010.

29. It is a diffusion index of "Accommodative" minus "Severe" in percentage points. The series for "All firms" was used.

30. Recall this restriction imposes that a 1% increase in both wealth and non-property income results in a 1% increase of consumption in the long-run. It was tested and accepted, as well as for the United States and the euro area. See Tables A2 and A4 in the annex.

Table 5. Order of integration in the Japanese data, in level and in differences (Δ and Δ²)

Series	ADF (level)[6]	ADF-GLS (level)[6]	KPSS (level)[6]	ADF (Δ)	ADF-GLS (Δ)	KPSS (Δ)	ADF (Δ²)	ADF-GLS (Δ²)	KPSS (Δ²)	Selected order of integration (level)
ln(c/pop)	I(1)	I(1)	I(1)	I(0)[1]	I(0)[1]	I(1)	I(0)	I(1)	I(0)	I(1)
ln(y/pop)	I(1)	I(1)	I(1)	I(0)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)[5]
ln(c/y)	I(1)	I(1)		I(0)	I(0)	I(0)	-	-	-	I(1)
hw/y	I(1)	I(1)	I(1)			I(0)	I(0)	I(0)	I(0)	I(1)
nfwh/y	I(1)	I(1)	I(0)	I(0)	I(0)	I(0)	-	-	-	I(1)
u	I(1)	I(1)		I(0)[1]	I(1)[2]	I(0)	-	-	-	I(1)[3]
r	I(1)	I(1)		I(0)	I(0)	I(0)	-	-	-	I(1)[4]
π	I(1)	I(1)	I(0)	I(0)	I(0)	I(0)	-	-	-	I(1)
r_s	I(1)	I(1)	I(1)	I(0)	I(1)	I(0)	-	-	-	I(1)

Note: "-" indicates that the test has not been run because it was unnecessary.

1. AIC indicates I(1).
2. Only AIC was conclusive.
3. Based also on sub-sample evidences.
4. No sign of over-differentiation in the correlogram of the series in first difference.
5. Based also on sub-sample from 1976Q1.
6. A time-trend was introduced to run the tests on the variables in level.

Source: Author's calculations.

3.2 Estimation of long-run consumption in Japan

35. Both of the Johansen (1996) tests for cointegration report one and only one cointegration relation between the benchmark regressors. Overall, and comparing with other estimation methods (see Table A3), there is strong evidence of cointegration. However, while financial wealth effects appear to be very high and significant, housing wealth effects may be either small or zero. This may be linked with the fact the credit supply index did not prove useful in this setup, either interacted with housing wealth or alone. It may not capture credit supply conditions properly.[31] Close comparison between methods would tend to confirm that there exists a long-term effect of housing wealth, though not much higher than 1 cent to the "dollar", while financial wealth effects would range from 4 to 7 cents to the "dollar". In total, wealth effects are smaller than in the United States in terms of MPCs, but still significant.

36. The positive coefficient on the interest rate should be interpreted as an income effect, and appears necessary to have strong cointegration (see regression C' and compare ADF statistics). It seemed to work better than the interest rate interacted with wealth, as regression (A') shows. The short-run interest rate could have played a comparable role, but it was not put into the regression due to high correlation with the long-term rate.

31. Neither in level, nor in cumulated sum, even after trend adjustments.

Table 6. Selected long-term regressions for Japan[1]

Dependent variable : *ln(c/y)=ln(c/pop)-ln(y/pop)*

Regressors	Benchmark	(A')	(B')	(C')
Cste	-0.279*** [-19.3]	-0.312*** [-21.3]	-0.31*** [-8]	-0.274*** [-11.7]
nfwh/y	0.061*** [20.5]	0.073*** [19.3]	0.068*** [11.2]	0.058*** [13.6]
hw/y	0.011*** [4.6]	0.015*** [5.5]	0.008*** [2.9]	0.017*** [5.8]
r	0.808*** [4.9]	1.869*** [5.9]	0.455* [1.7]	
*r*nfwh/y*		-0.594*** [-3.7]		
u			0.007 [1.3]	
π			-0.816** [-2.5]	
$u*\pi$			0.545*** [3.7]	
r_S			0.661*** [4.1]	
Sample	76Q1-08Q4	76Q1-08Q4	76Q1-08Q4	76Q1-08Q4
R^2	0.94	0.95	0.96	0.9
BIC	-5.436	-5.575	-5.668	-5.023
SER	0.0151	0.0138	0.0127	0.0188
DW	0.504	0.534	0.654	0.251
ADF[2]	-4.59	-4.63	-5.40	-3.16

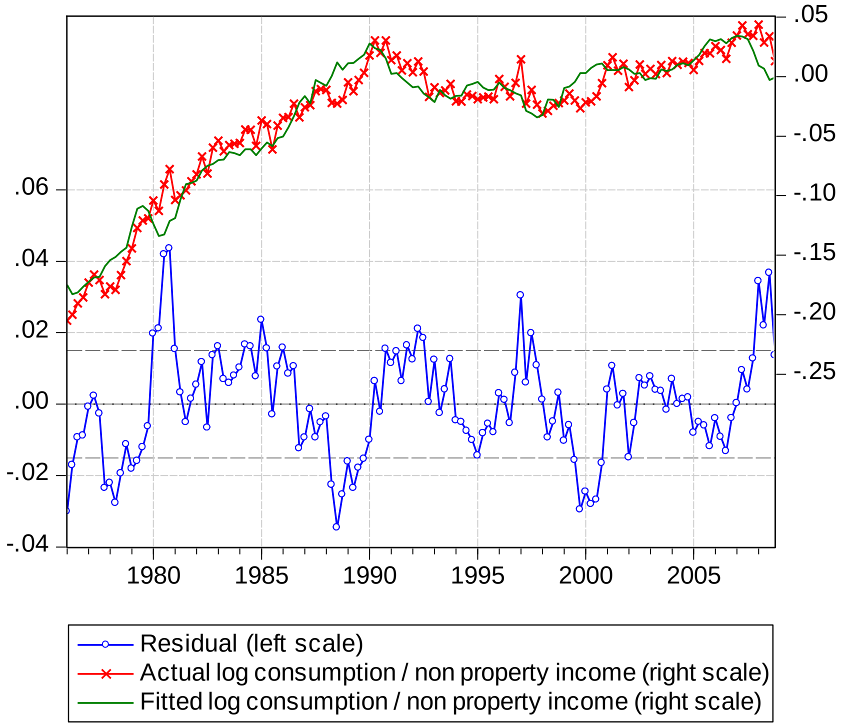
1. Newey-West (1987 and 1994) robust covariance matrix is used to compute indicative t-statistics in brackets. See footnote 1 of Table 3. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.
2. Asymptotic critical values reported in Phillips and Ouliaris (1990) are, for *n*=3 regressors in the long-run relation, -4.1 and -4.7 for nominal sizes of 5% and 1% respectively.

Source: Author's calculations.

3.3 Some interpretation of the long-run results

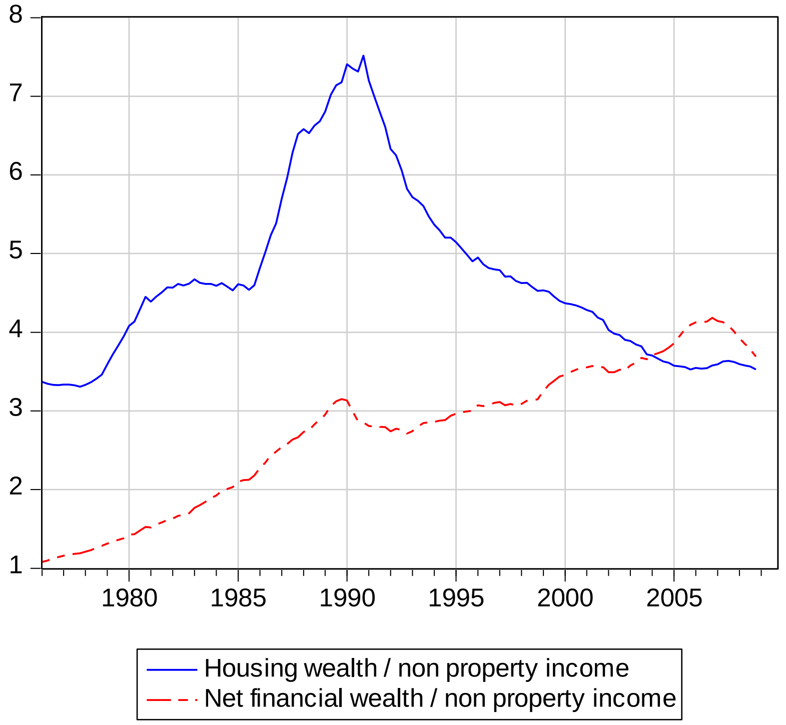
37. The striking, though not surprising, feature of Figure 5 is the change of trend after, roughly, 1990. Following the 1990 peak, decreases in financial asset prices and housing wealth suppressed consumption for a sustained period (see Figure 6). Lower real interest rates also contributed by implying lower property income. More recently, increases in the property-income ratio, especially around the year 2000 supported (sluggish) increases in consumption. This overall picture is consistent with the decrease in Japanese consumption during the crisis in 2008 while stock prices were dropping.

Figure 5. Residuals of the long-term benchmark regression for Japan



Source: OECD, author's calculations.

Figure 6. Stocks of Japanese households wealth



Source: OECD.

3.4 Short-term dynamics of Japanese consumption

38. In the benchmark equation (see Table 7), inflation changes $\Delta\pi$ were put to zero when inflation is negative, so that the actual regressors have the form $\Delta\pi.1_{\pi>0}$. The benchmark regression is quite robust to the addition of new regressors or to a change of sample. There is mild evidence of parameter instability [compare regressions (E') and (F')], but not large enough to result in the failure of Chow tests. A quantile regression for the median brings comparable estimates for all parameters, and the residual term tells a very similar story (correlation with the OLS residual from the benchmark is about 0.995). All N-step forecasts tests pass.[32]

39. The error correction coefficient is always large, showing the return to long-run equilibrium is rather fast, approximately twice as fast as in the United States. The effect of improving lending attitude to businesses, possibly capturing the prevailing economic climate, shows some positive effect on consumption growth.

40. As for the United States, regression (G') shows some signs of non-linearity in the adjustment to the long-run trend. It seems the adjustment of consumption is faster downward -- for instance, after a sharp fall in asset prices -- than upward (for a given gap in the long-run relationship). This may be linked to several factors. First, it is never a necessity to increase spending when ones wealth is increasing quickly, whereas the contrary could be. Second, consumption is above equilibrium when asset prices are low, which could be associated with a period of economic distress triggering sharper than usual revisions of coming economic prospects among households. Contrary to the US case, a Wald test mildly rejects the hypothesis that this asymmetry is zero. However the linear-in-wealth model was kept for the sake of simplicity and cross-country consistency. The regression (G') also illustrates the rationale for putting the inflation-change terms only when inflation is positive. It is not surprising to see the effects of inflation changes are different when inflation is actually *deflation*. The regression (G') shows that they may be about zero in a deflationary context.[33] As showed by Figure 7, the benchmark equation is relatively crisis-proof when used in an out-of-sample forecast throughout the crisis. Note that if, contrary to what is done in the benchmark equation, the variables $\Delta\pi$ and $\Delta\pi(-1)$ are included in the regression without putting them to zero when inflation is negative, then the Chow test fails and big forecasting errors are made through the ongoing crisis, while deflation is large.

32. See footnote 26.

33. This empirical experiment can only be done with Japan since only Japan has a large sample where inflation is negative.

Table 7. Selected short-term regressions for Japan[1]

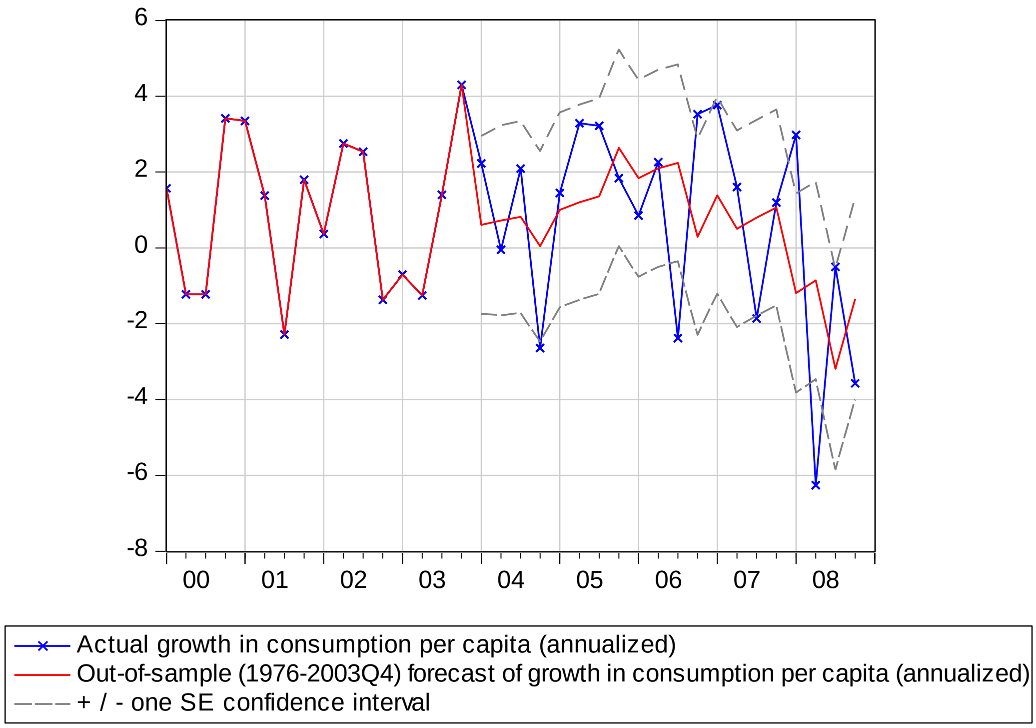
Dependent variable: $\Delta ln(c/pop)$

Regressors	Benchmark	(D')	(E')	(F')	Regressors	(G')
Cste	0.001	0.002*	0.001	0.001	*Cste*	0.002**
	[1.4]	[1.9]	[0.5]	[1.1]		[2.1]
ect(-1)	-0.236***	-0.213***	-0.239***	-0.276***	$ect_{t-1}.1_{ect>0}$	-0.338***
	[-5.9]	[-4.8]	[-4.8]	[-4.9]		[-5.6]
$\Delta ln(y/pop)$	0.271***	0.256***	0.282***	0.372***	$ect_{t-1}.1_{ect<0}$	-0.129**
	[4.1]	[3.8]	[3.7]	[3.7]		[-2.2]
$\Delta ln(c/y)(-1)$	-0.153***	-0.217***	-0.155**	-0.158**	$\Delta ln(y/pop)$	0.272***
	[-2.8]	[-3.2]	[-2.4]	[-2.0]		[3.9]
$\Delta i_s(-2)$	0.348***	0.325***	0.328***	0.274	$\Delta ln(c/y)(-1)$	-0.153***
	[4.8]	[4.1]	[4.4]	[1.2]		[-2.8]
$\Delta i(-4)$	-0.362***	-0.381***	-0.374***	-0.319***	$\Delta i_s(-2)$	0.366***
	[-3.4]	[-3.6]	[-3.2]	[-2.6]		[4.8]
$\Delta\pi.1_{\pi>0}$[4]	-0.604***	-0.56***	-0.626***	-0.642	$\Delta i(-4)$	-0.36***
	[-3.6]	[-3.3]	[-3.4]	[-1.6]		[-3.6]
$\Delta\pi(-1).1_{\pi>0}$	0.373***	0.258**	0.389***	0.444**	$\Delta\pi.1_{\pi>0}$	-0.611***
	[3.3]	[2]	[3.3]	[2.2]		[-3.6]
$\Delta ln(c/pop)(-3)$	0.172***	0.18***	0.206***	0.112*	$\Delta\pi(-1).1_{\pi>0}$	0.414***
	[3.7]	[3.8]	[3.0]	[1.8]		[3.5]
$\Delta(tankan/1000)(-1)$	0.196**	0.177*	0.202	0.069	$\Delta\pi.1_{\pi<0}$	0.053
	[2.3]	[1.9]	[1.6]	[0.5]		[0.4]
Δu		-0.001			$\Delta\pi(-1).1_{\pi<0}$	-0.134
		[-0.2]				[-0.9]
$\Delta u(-1)$		-0.007			$\Delta ln(c/pop)(-3)$	0.173***
		[-1.5]				[3.7]
$\Delta u(-2)$		0.004			$\Delta(tankan/1000)(-1)$	0.218***
		[0.9]				[2.6]
$\Delta ln(y/pop)(-1)$		-0.115				
		[-1.5]				
Wealth change[2]		0.004				
		[0.3]				
Set of dummies[3]	yes	yes	yes	yes	*Set of dummies*	yes
Sample	76Q2-08Q4	76Q2-08Q4	76Q2-97Q4	87Q1-08Q4		76Q2-08Q4
R^2	0.66	0.66	0.70	0.70		0.68
BIC	-7.101	-6.93	-6.86	-6.903		-7.041
SER	0.0057	0.0058	0.0061	0.006		0.0057
DW	2.198	2.049	2.293	2.085		2.34
Q-test	0.096	0.092	0.105	0.162		0.047
Q²-test	0.021	0.079	0.033	0.165		0.050
Jarque-Bera	0.71	0.55	0.88	0.55		0.73
Chow F test 1993	0.329	0.127	-	-		0.409

1. White (1980) robust covariance matrix is used to compute robust t-statistics in brackets. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.
2. Equals *ln(nfwh/pop)(-1) - ln(nfwh/pop)(-5).*
3. 1997Q1 and Q2 are dummied out. They are extreme observations with opposite signs. Also, a break in the constant term is allowed about the early nineties. More specifically, a dummy equal to one until 1992Q3 is introduced (the exact quarter was selected among other dates about 1991 by a BIC criterion).
4. When $\pi<0$, changes in inflation are not significant anymore. See regression (G').

Source: Author's calculations.

Figure 7. Japan: out-of-sample forecast over the crisis period



Source: Author's calculations.

4. Consumption in the euro area: smaller but significant wealth effects

41. This section expands the previous methodology used for the United States and Japan to the European Monetary Union (EMU) as a whole, referred to as the euro area.

4.1 The dataset

42. Empirical analysis for the euro area is frequently impeded by the lack of long and consistent historical data. Several sources were used for this study that relies on data starting about 1980. The euro area refers, for data-availability reasons, to the euro area with 12 members,[34] although some small countries among these may be missing in some time-series. In this case, proper adjustments were made. Overall, the economic and econometric behaviour of this euro area must be very similar[35] to the one of the current euro area. As for the data related to pre-EMU period -- before 1999 --, they relate to a fictitious "euro area" having the same composition as that after 1999. Housing wealth data was obtained by interpolation of annual data from the European Central Bank (ECB). Quarterly financial wealth data from the ECB was used, spliced with OECD estimates prior to 1999. Consumption, GDP, population, household

34. Namely: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain.

35. Which would not remain true if a new large European country joined the EMU.

account variables, unemployment rate and interest rates were aggregated from the OECD's Analytical Database.[36] House prices and fiscal data are also from the OECD.

43. The same investigations are made as for Japan and the United States. Testing the orders of integration was made using the longest available sample from 1980, with an intercept (and a time-trend if indicated).

Table 8. Order of integration in the euro area data, in level and in first difference (Δ)

Series	ADF (level)[4]	ADF-GLS (level)[4]	KPSS (level)[4]	ADF (Δ)	ADF-GLS (Δ)	KPSS (Δ)	Selected order of integration (level)
ln(c/pop)	I(1)	I(1)		I(1)	I(1)	I(0)	I(1)[2]
ln(y/pop)	I(0)[1]	I(1)	I(0)	I(0)	I(1)[1]	I(0)	I(1)[2]
ln(c/y)	I(1)	I(1)	I(1)	I(0)	I(0)[3]		I(1)[2]
nfwh/y	I(1)	I(1)	I(1)	I(1)	I(0)[1]	I(0)	I(1)[2]
hw/y	I(1)			I(1)		I(0)	I(1)[2]
ln(nfwh/pop)	I(1)	I(1)	I(1)	I(1)			I(1)[2]
ln(ydh/pop)	I(1)	I(1)	I(0)	I(0)	I(1)[1]	I(0)	I(1)[2]
π	I(1)	I(1)[1]		I(0)	I(1)[1]	I(0)	I(1)[2]

1. Only AIC brought a conclusive statistic.
2. The correlogram of the differenced series did not show any sign of integration or over-differentiation.
3. AIC indicates I(1).
4. A time-trend was introduced to run the tests on the variables in level.

Source: Author's calculations.

4.2 Estimation of the long-run consumption in the euro area

44. The benchmark regression in Table 9 shows evidence of cointegration. The Trace and Maximum eigenvalue sequential tests generally used with the Johansen (1996) method also indicate one and only one cointegration relation in the benchmark case. Other estimation methods bring similar, stable results, except for the housing wealth effect that is estimated to be near zero in a one-step-OLS equation (see Table A4). Overall, evidence of a cointegrating relation between wealth, income and consumption is strong, even though it is milder for housing wealth effects. Financial wealth effects appear to range from 5 to 7 cents to the dollar while housing wealth effects would be between zero and 2 cents. It is worth noting that the housing wealth variable is needed in the cointegration relation, and hence was included in the benchmark equation [see (D'') and (E'') and their ADF statistics]. Inflation also seems to have played a significant role in explaining consumption in the euro area [see the ADF statistic of (C'')], though its current level and volatility are lower in the recent period implying a much milder impact on consumption. Surprisingly, no effects of interest rates is discernable, contrary to the US and Japanese cases. This should not be attributed to the heterogeneity of the euro area in terms of real interest rates because they have been more or less following the same trend since 1980 in the four biggest countries of the zone. However, such a result could arise from heterogeneity in the *response* of each country to this interest rate, rendering the estimation of a single coefficient problematic.

36. Version: 15 February 2010

Table 9. Selected long-term regressions for the euro area[1]

Dependent variable : *ln(c/y)=ln(c/pop)-ln(y/pop)*

Regressors	Benchmark	(A'')	(B'')[3]	(C'')	(D'')	(E'')
Cste	-0.044***	-0.049**	-0.015	-0.117***	-0.032*	-0.023
	[-2.8]	[-2.3]	[-0.5]	[-8.5]	[-1.7]	[-1.3]
π	-0.774***	-0.757***	-1.955***		-0.732***	-1.101***
	[-5.0]	[-4.9]	[-8.7]		[-4.5]	[-9.9]
nfwh/y	0.053***	0.053***	0.063***	0.089***	0.071***	
	[6.7]	[6.5]	[9.8]	[9.6]	[10.9]	
hw/y	0.013***	0.014***	0.010***	0.006		
	[4.1]	[3.3]	[2.9]	[1.3]		
*r*nfwh/y*		0.022				
		[0.3]				
u			-0.006***			
			[-2.9]			
r			0.095			
			[0.4]			
$u*\pi$			0.183***			
			[6.7]			
r_s			-0.122			
			[-0.7]			
(nfwh+hw)/y						0.025***
						[10.8]
Sample	80Q4-08Q4	80Q4-08Q4	80Q4-08Q4	80Q4-08Q4	80Q4-08Q4	80Q4-08Q4
R^2	0.98	0.98	0.99	0.95	0.97	0.96
BIC	-6.394	-6.354	-6.770	-5.64	-5.954	-6.053
SER	0.0093	0.0093	0.0072	0.0137	0.0117	0.0112
DW	0.406	0.410	0.550	0.165	0.264	0.346
ADF[2]	-5.55	-5.60	-5.51	-3.39	-3.10	-3.34

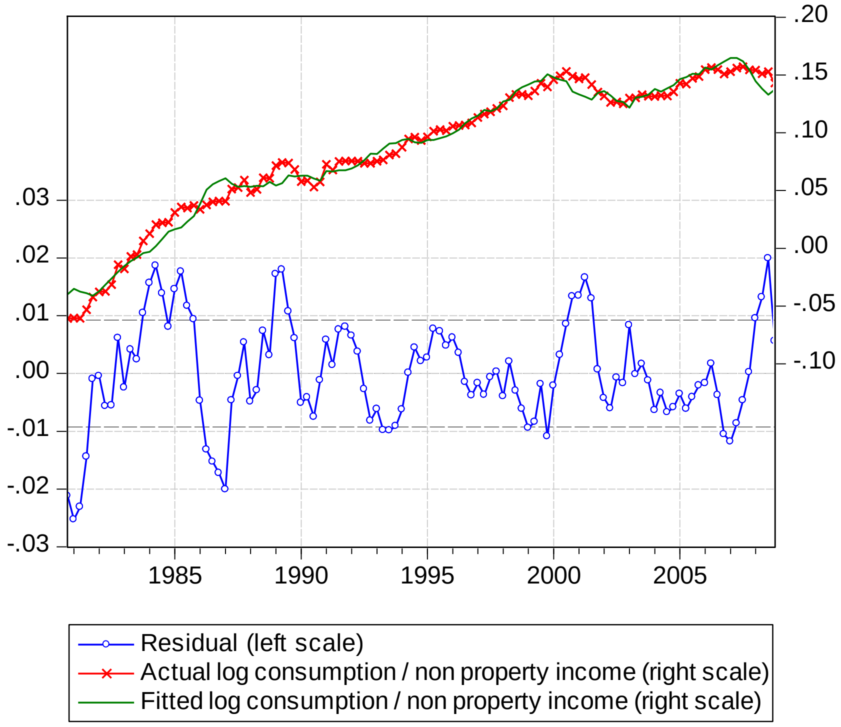
1. Newey-West (1987 and 1994) robust covariance matrix is used to compute indicative t-statistics in brackets. See footnote 1 of Table 3. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.
2. Asymptotic critical values reported in Phillips and Ouliaris (1990) are, for *n*=3 regressors in the long-run relation, -4.1 and -4.7 for nominal sizes of 5% and 1% respectively.
3. Unemployment put alone together with the benchmark does not bring a significant estimator.

Source: Author's calculations.

4.3 Some interpretation of the long-run results

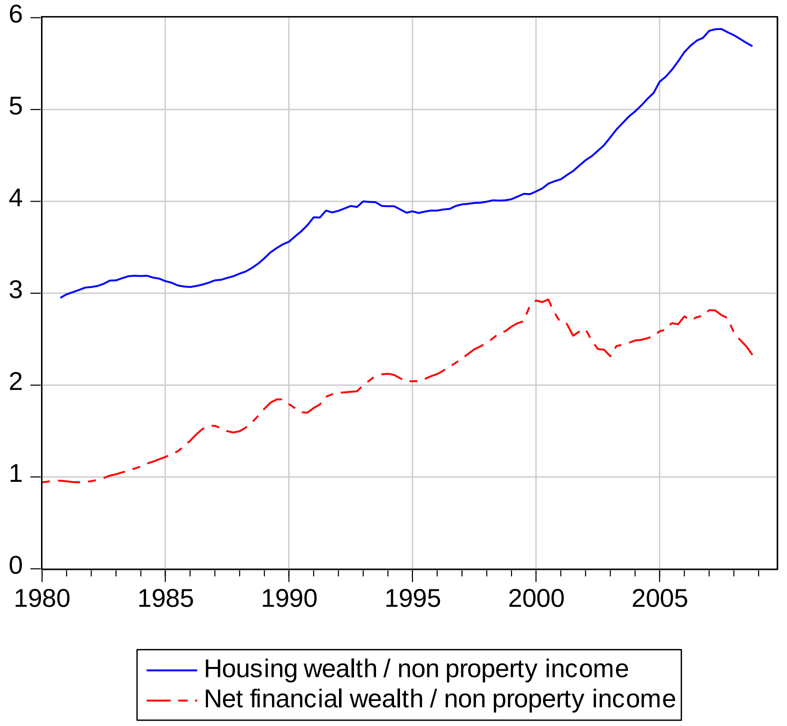
45. In the euro area too, the consumption increased a lot as a share of non-property income over the estimation period, and this upward trend did not abate about 1990 as it did in the United States and Japan, although it did about 2000. The high slope between 1980 and 1985 may be associated with a sharp fall in inflation. Otherwise this trend should be linked to the increase of wealth stocks as shares of non-property income (see Figure 9). This can account for the structural decrease in saving rate in the euro area, mitigated by a rise of the share of property income in total disposable income of more than one third. However, two turning points are noticeable, first after 2000 when stock market prices sharply decreased and second in 2007 with deflation of both housing and stock market prices, with consumption reacting with a lag to financial conditions. However, one must admit that the euro area data appears less informative than that for Japan and the United States, showing less peaks and troughs.

Figure 8. Residuals of the long-term benchmark regression for the euro area



Source: OECD, author's calculations.

Figure 9. Stocks of euro area households wealth



Source: OECD, European Central Bank.

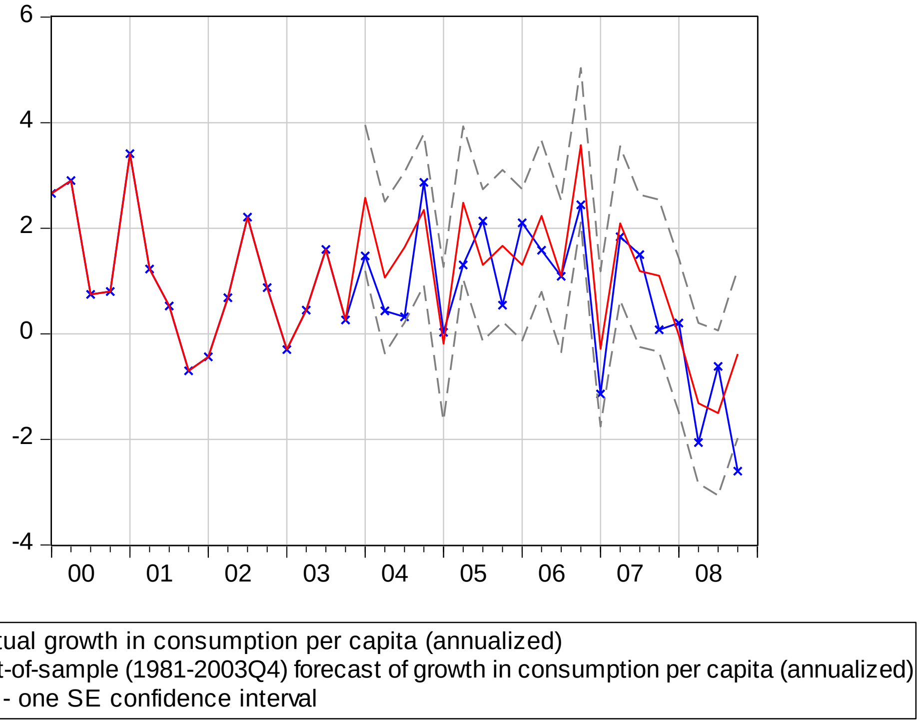
4.4 Short-term dynamics of euro area consumption

46. The benchmark regression (Table 10) is relatively robust to a shift of the estimation sample or to the addition of new regressors. A quantile regression for the median gives comparable estimates for all parameters, and brings residuals with 0.99 correlation with the OLS-residuals. N-step forecast tests[37] indicate no failure of the medium-term forecasting capabilities of the equation over time, supporting reasonable parameter stability.

47. The estimated error correction coefficient in not as large as in the Japanese case, but comparable to the US one. Although, it is less precisely estimated, all equations seem to show significant adjustment to the long-term equilibrium. Only a few regressors were found significantly partially correlated with consumption growth for the euro area. It is surprising to see that unemployment change does not impact consumption significantly. This may be due to an unemployment rate more sluggish than in the United States, making such effects difficult to estimate with precision, or due to the presence of more generous social safety nets that have been developed by European countries. Some -- but not statistically significant -- evidence of non-linearity is shown by equation (I''), as it seems the correction of the previous period error term is more rapid when households over-consume, as measured by the long-run relationship. This feature also shows for Japan and the United States.

48. The out-of-sample forecast exercise in Figure 10 shows that the equation has reasonable forecasting capabilities (recall all the regressors are taken exogenous for this forecast). The main problematic feature of the equation is the large negative residual in the last period. However, it is almost within the one standard error boundaries, and well within the two standard error boundaries.

Figure 10. Euro area: out-of-sample forecast over the crisis period



Source: Author's calculations.

37. See footnote 26.

Table 10. Selected short-term regressions for the euro area[1]

Dependent variable : $\Delta ln(c/pop)$

Regressors	Benchmark	(F'')[3]	(G'')	(H'')	Regressors	(I'')[4]
Cste	0.002***	0.002***	0.003***	0.001**	*Cste*	0.002**
	[4.3]	[3.9]	[3.5]	[2.5]		(0.001)
ect(-1)	-0.099**	-0.102**	-0.091*	-0.121**	$ect_{t-1}.1_{ect>0}$	-0.138***
	[-2.5]	[-2.4]	[-1.8]	[-2.3]		(0.053)
$\Delta ln(ydh/pop)$	0.703***	0.758***	0.725***	0.787***	$ect_{t-1}.1_{ect<0}$	-0.063
	[7.5]	[3.2]	[6.9]	[8.4]		(0.075)
$\Delta ln(c/pop)(-1)$	-0.179**	-0.336*	-0.234**	-0.085	$\Delta ln(ydh/pop)$	0.706***
	[-2.1]	[-1.9]	[-2.1]	[-1.1]		(0.093)
Wealth change[2]	0.026***	0.024***	0.022**	0.025***	$\Delta ln(c/pop)(-1)$	-0.181**
	[3.9]	[3.3]	[2.1]	[3.3]		(0.087)
$\Delta^2 ln(rp_{house})$	0.113***	0.109***	0.113***	0.048	*Wealth change*[2]	0.026***
	[3.2]	[3.0]	[2.9]	[1.0]		(0.007)
$\Delta\pi$		0.014			$\Delta^2 ln(rp_{house})$	0.11***
		[0.1]				(0.036)
$\Delta\pi(-1)$		0.013				
		[0.1]				
$\Delta ln(c/pop)(-2)$		-0.015				
		[-0.2]				
$\Delta i(-1)$		-0.001				
		[-0.5]				
$\Delta i_s(-1)$		0				
		[0.1]				
$\Delta u/100$		-0.285				
		[-1.1]				
$\Delta ln(y/pop)$		-0.092				
		[-0.5]				
$\Delta ln(ydh/pop)(-1)$		0.125				
		[0.8]				
$\Delta ln(c/y)(-1)$		0.12				
		[1]				
Set of dummies	no	no	no	no	*Set of dummies*	no
Sample	81Q1-08Q4	81Q1-08Q4	81Q1-99Q4	90Q1-08Q4		81Q1-08Q4
R^2	0.59	0.60	0.57	0.71		0.59
BIC	-8.447	-8.104	-8.167	-8.822		-8.412
SER	0.0032	0.0033	0.0036	0.0026		0.0032
DW	2.080	2.049	2.073	1.909		2.091
Q-test	0.580	0.489	0.557	0.337		0.492
Q²-test	0.003	0.003	0.043	0.355		0.001
Jarque-Bera	0.46	0.44	0.99	0.93		0.59
Chow F test 1995	0.713	0.563	-	-		0.826

1. White (1980) robust covariance matrix is used to compute robust t-statistics in brackets. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.
2. Equals *ln(nfwh/pop)(-1)- ln(nfwh/pop)(-4).*
3. F test for the joint nullity of non benchmark regressors has p-value equal to 0.90.
4. In parentheses: robust standard errors.

Source: Author's calculations.

5. Cross country comparison of overall wealth effects: short-term and long-term response functions

49. The benchmark equations are neither linear nor log-linear. This means it is not straightforward to estimate the (long-term) effects of a step increase in wealth, either in levels or in percentage terms. To do so, one can use an historical period and compare the benchmark forecast to the forecast with the step increase in wealth over the same period. However, such "impulse response" would depend on the path of many variables, mainly income, so it is not suitable for a cross country comparison. One can solve this problem by "freezing" the economy about its mean behaviour, that is, freezing all I(1) variables equal to their mean in a given period. Then, the benchmark equation can be used to calculate the difference between a forecast of consumption with and without a step increase of 1% in wealth, which no longer has a time dependence to income or any other series. Though this is a crude method to gauge the "response functions", such a method must be used as none of the estimated parameters delivers the MPCs or the elasticities relative to wealth directly. This allows giving figures related to a static context instead of dealing with time-dependent dynamics of the economies.

50. More specifically, the averages of population, consumption per capita, income per capita, wealth as a share of income, inflation, interest rates, credit supply index were calculated over 2000Q1-2008Q4. These variables were then set to these levels. Stationary regressors were set to zero or to a constant value. Dummies were removed. Figures 11, 13 and 15 show the percentage increase in real consumption resulting from a 1% increase in wealth, which is the estimated elasticity of consumption to wealth. Figures 12, 14 and 16 show this increase in the level of consumption divided by the step increase in wealth, and then multiplied by 100, so it represents the dollar increase in consumption in response to a 100 dollars increase in wealth, which is the MPC out of 100 dollars of wealth. The impulse responses are calculated on a ten-year window and are represented over the period 2000-10, as the means were taken on a comparable timeframe.

51. Recall that house and asset prices can be correlated through various channels, so that increases in one of the wealth stocks is likely to imply an increase in the other one. Figure 15 shows the percentage impact of the increase of 1% in total net worth of households, by just summing the results of Figures 11 and 13. Figure 16 gives the dollar increase in consumption due to a 100 dollars increase in total wealth by taking an average of the responses in Figures 14 and 12, using average financial and housing wealth stocks as weights.[38] Overall, the wealth elasticity of consumption is larger in Japan and in the United States than in the euro area, while the MPC out of wealth is significantly larger in the United States, mainly due to higher propensity to consume out of housing wealth. Note the difference between Figure 15 and 16 due to higher wealth to consumption ratio in Japan than in the euro area and in the United States (respectively, almost 7.5, 6.7 and 5.6, taking the means).

52. The broad conclusions arising from these figures are that:

- The MPCs out of $100 of financial wealth are relatively similar in the United States, in Japan and in the euro area (about $5 to $6) – Figure 12.
- Only the United States displays housing wealth effects of a comparable order of magnitude to those from financial wealth – Figure 14.
- The elasticity of consumption to total wealth seems to be higher in Japan and in the United States than in the euro area – Figure 15.
- The MPC out of $100 of total wealth is higher in the United States (about $5) than in Japan (just over $3.5) and the euro area (about $3) – Figure 16.

38. Accordingly, the weight attributed to the financial wealth curves relative to housing wealth curves are 1.529 for the United States, 0.991 for Japan and 0.515 for the euro area.

Figure 11. Percentage increase in consumption following a 1% step increase in net financial wealth (elasticity)

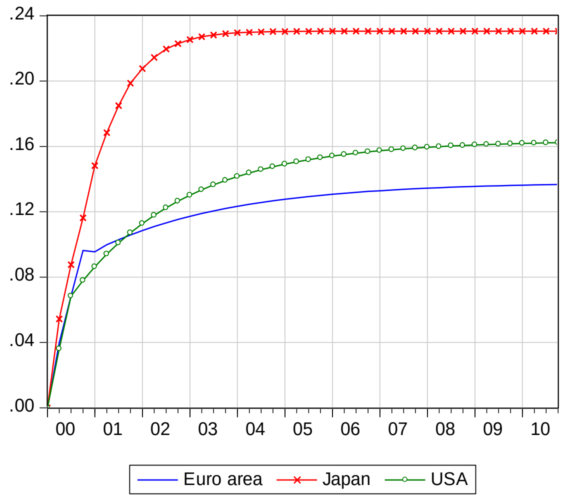
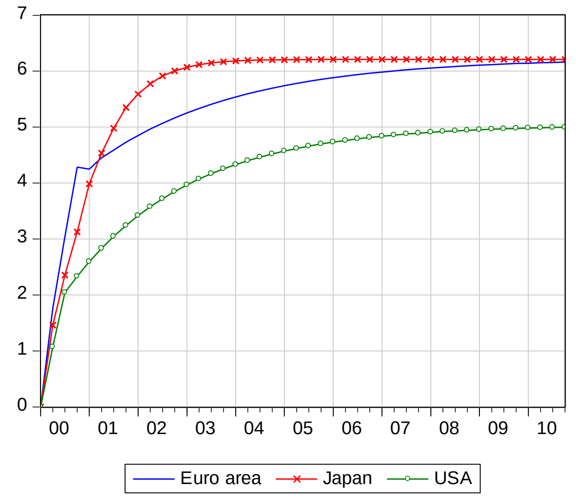


Figure 12. Dollar increase in consumption following a $100 step increase in net financial wealth (MPC)



Source: Author's calculations.

Figure 13. Percentage increase in consumption following a 1% step increase in housing wealth (elasticity)

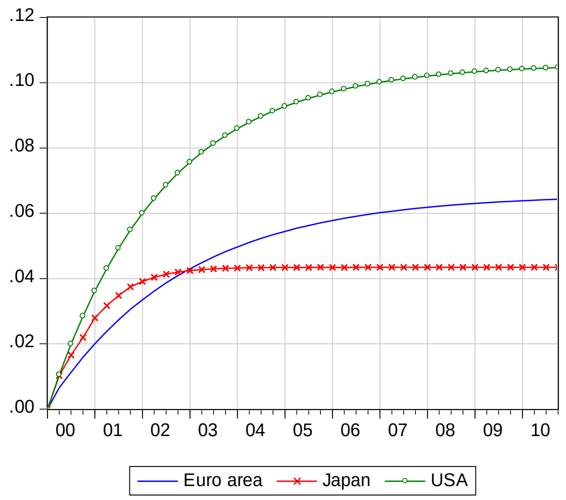
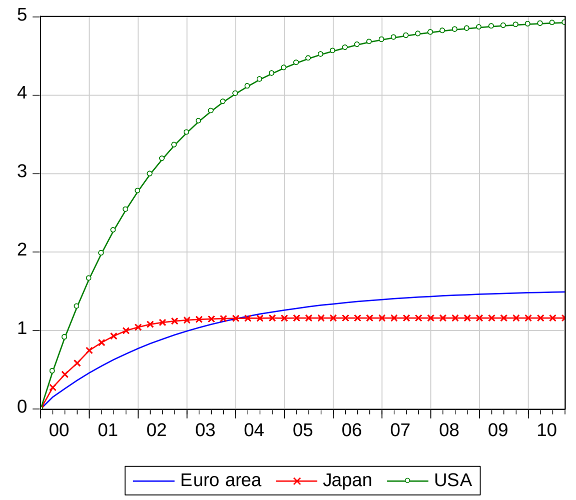


Figure 14. Dollar increase in consumption following a $100 step increase in housing wealth (MPC)



Source: Author's calculations.

Figure 15. Percentage increase in consumption following a 1% step increase in total wealth (elasticity)

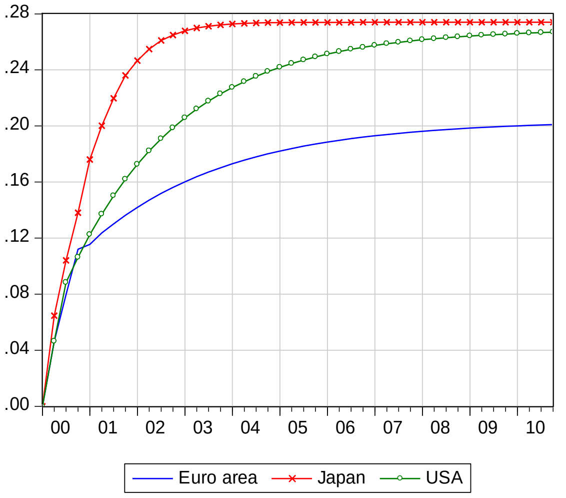
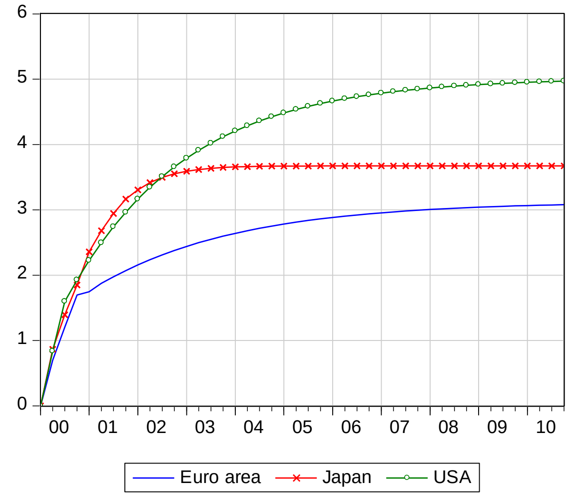


Figure 16. Dollar increase in consumption following a $100 step increase in total wealth (MPC)



Source: Author's calculations.

6. How important are changes in wealth in forecasting consumption over the recent crisis?

53. The above figures give a good idea of the magnitude of wealth effects that can be expected in the short and long run. It is of interest to use the recent events to gauge how big would have been the forecasting error if wealth-income (W/Y) had been kept constant from 2006Q1 onwards, all else equal.[39] This exercise compares the dynamic forecasting errors using the benchmark regression with the actual data and with the same data where wealth-income ratios were kept constant.

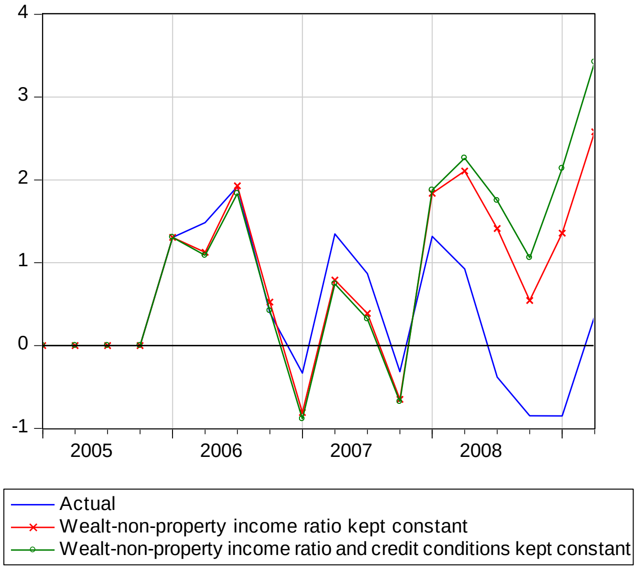
54. Overall, one can see that consumption behaviour is less well tracked by the equation when wealth variations are not taken into account both in terms of errors in the forecast level of consumption and in the variability of errors (see Tables 11 to 13).[40] This is well illustrated over the year 2008 for the United States, Japan and the euro area, though the errors are much larger for the United States, where wealth has decreased much more compared to Japan and the euro area.[41] This underpins the importance of taking wealth into account, especially when MPCs are large and also when it displays considerable and sharp variations. The larger volatility of households' wealth in the United States makes it of greatest importance to capture households' wealth effects in the recent downturn and going forward.

39. This means the increase in unemployment or the drop in interest rates are kept equal to their actual values in this exercise.

40. Note that from a mathematical point of view, it is not bound to be the case, since only a small part of the whole estimation sample is being considered here.

41. Overall, wealth to non-property-income ratios were respectively for the United States, Japan and the euro area, in 2006Q1, 4.2, 4.1 and 2.7 for net financial assets, 3.0, 3.6 and 5.6 for real estate, while at the end of the sample they were 3.2, 3.7 and 2.3 for net financial assets, and 2.3, 3.5 and 5.7 for real estate (housing wealth increased in 2007 in the euro area), implying variations of -24%, -10% and -15% for net financial assets and -26% in US real estate and almost no change in Japan or euro area housing wealth.

Figure 17. Compared forecasting errors on the (annualised) growth rate of US consumption (window: 2005Q1-2009Q2)



Source: Author's calculations.

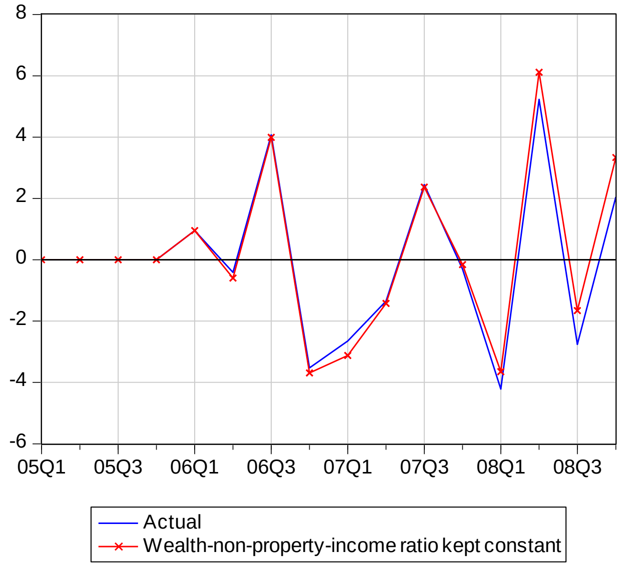
Table 11. Compared forecasting errors of US consumption

	% error on level consumption by the end of the sample[1]	Root mean square error of forecast relative to actual[2]
Actual	+1.80	1
W/Y constant	+3.61	1.36
W/Y and credit conditions constant	+4.16	1.59

1. Approximated by cumulated sum of errors on Figure 17 (divided by 4).
2. Averages taken from 2006Q1 to the end.

Source: Author's calculations.

Figure 18. Compared forecasting errors on the (annualised) growth rate of consumption in Japan (window: 2005Q1-2008Q4)



Source: Author's calculations.

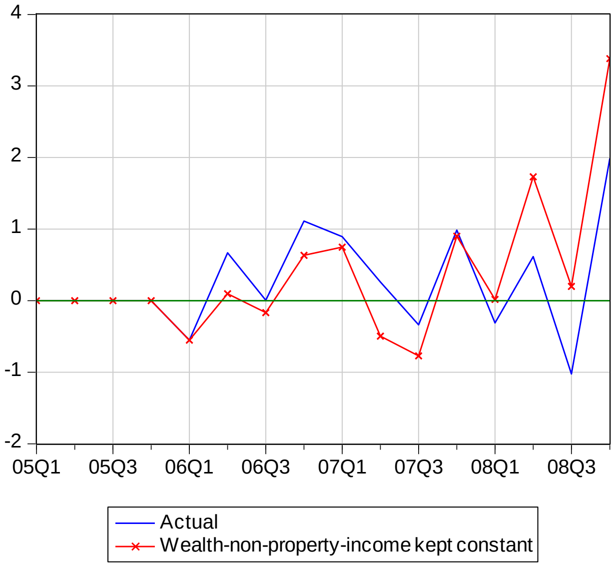
Table 12. Compared forecasting errors of consumption in Japan

	% error on level consumption by the end of the sample[1]	Root mean square error of forecast relative to actual[2]
Actual	-0.11	1
W/Y constant	+0.62	1.05

1. Approximated by cumulated sum of errors on Figure 18 (divided by 4).
2. Averages taken from 2006Q1 to the end.

Source: Author's calculations.

Figure 19. Compared forecasting errors on the (annualised) growth rate of euro area consumption (window: 2005Q1-2008Q4)



Source: Author's calculations.

Table 13. Compared forecasting errors of euro area consumption

	% error on level consumption by the end of the sample[1]	Root mean square error of forecast relative to actual[2]
Actual	+1.08	1
W/Y constant	+1.43	1.36

1. Approximated by cumulated sum of errors on Figure 19 (divided by 4).
2. Averages taken from 2006Q1 to the end.

Source: Author's calculations.

7. How sensitive are long-run household savings to asset prices?

55. The work here sheds additional light on the sensitivity of aggregate household savings to asset price changes in the United States. These changes can dramatically shift households' wealth or their ability to use it to purchase consumption goods, and therefore have an impact on the saving rate. Since the beginning of the crisis in 2008, the US saving rate has increased a lot towards pre-2000 levels. This section aims at gauging how the US savings could evolve over the coming years, given various assumptions on asset prices and credit conditions.

56. To gauge the sensitivity of US consumption to asset prices and financial conditions according to the US equation, a medium-term projection is generated in which explanatory variables are taken from the OECD Medium Term Baseline (MTB) scenario [see OECD (2009) Chap.4].[42] The variables left endogenous in this partial equilibrium model are savings (derived from consumption[43]) and wealth stocks, accumulated from savings[44] and inflated by changes in asset prices. The main goal here is not the (stylised) projection itself, which is very conditional on the recovery scenario described in the MTB, but its sensitivity to the path of asset prices and credit supply. In the baseline projection scenario, non-liquid asset prices increase by 2% per year in real terms over the projection, while house prices remain flat in real

42. Also, the share of property income in disposable income is supposed to converge towards reasonable recent-historical average. Population forecasts are from the US Census Bureau.

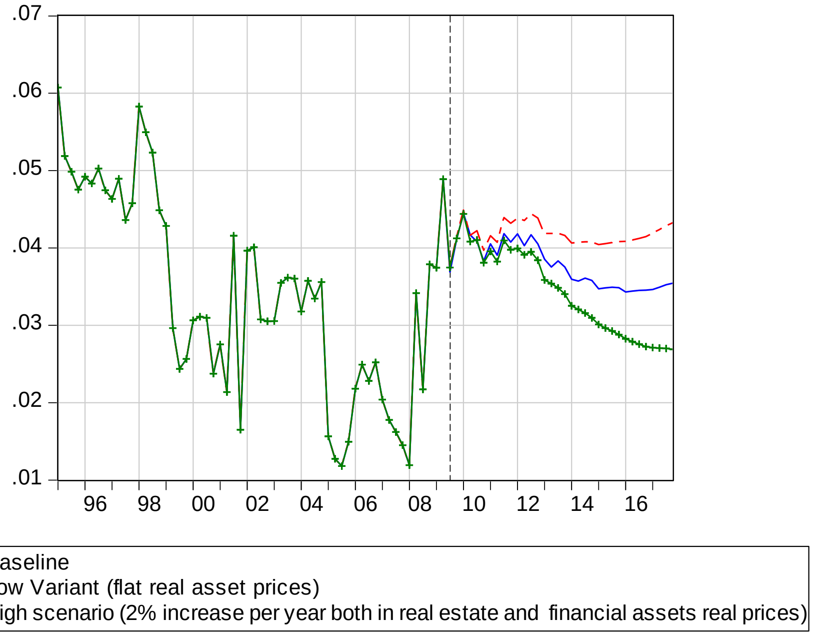
43. Savings in US national accounts is not equal to disposable income minus consumption. Small transfers are also taken into account and projected constant as shares of disposable income when not available in the OECD MTB.

44. Capital transfers like remittances are neglected and introduce a small bias in savings. But the precision of these forecasts could not reach this level of precision anyway.

terms. In the meantime, credit conditions smoothly improve towards pre-crisis levels. The prices of liquid assets are assumed constant in nominal terms. At each period savings are shared between all the wealth stocks.[45]

57. In this baseline scenario, the saving rate would stabilise about 3.5%, well above its 2000-08 average (2.7%) but below its historical average level of the 1990s (5.3%). Figure 20 shows the influence of asset prices on US savings. The two other curves are based on the same scenario provided that house prices also increase by 2% in real annualised terms (high scenario -- low savings) or that both house and stock market prices remain flat in real terms (low scenario -- high savings). The projected profile of the saving rate is subject to particular uncertainty as regards the recovery of asset prices. If asset prices move in line with the consumption deflator, and so remain constant in real terms, the saving ratio is projected to remain close to recent levels, just over 4%. Conversely if there is a recovery in real house prices, equity prices and other financial prices, then the saving ratio would gradually decline towards 2000-08 historical averages below 3%, while remaining above the low levels reached in the years preceding the present crisis. Overall, the household saving rate would be about 1.5 percentage points higher in the low scenario than in the high scenario.

Figure 20. US household saving rate and its sensitivity to asset prices



Source: OECD, author's calculations.

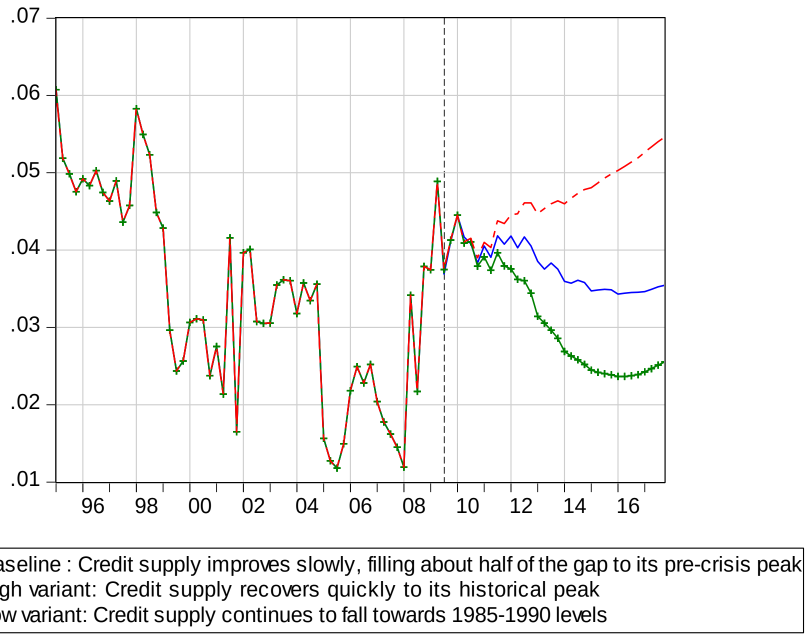
58. Figure 21 considers the influence of credit supply conditions on consumption. Here, the low and high scenario are based on the same assumptions on asset prices as in the baseline, but the credit conditions index path varies according to Figure 22.[46] If credit conditions continued to tighten, the saving rate might be expected to increase towards 1990-2000 historical levels over 5%. Conversely, if the availability of credit were to recover quickly and entirely to its pre-crisis peak, personal savings would fall about 2.5%. The evolution of asset prices could therefore be a major driver of the saving rate in the United-States in the coming years. Would asset prices follow the low scenario, US households would keep saving, trying to restore their pre-crisis wealth-income ratios. Similarly, the tightness of consumer credit could constrain the

45. The share of each stock is related to its share in total wealth. More specifically, 36% of the savings go to housing wealth, 14% in liquid assets and 50% in other assets net of liabilities.

46. Note that both limit scenarios are rather extreme.

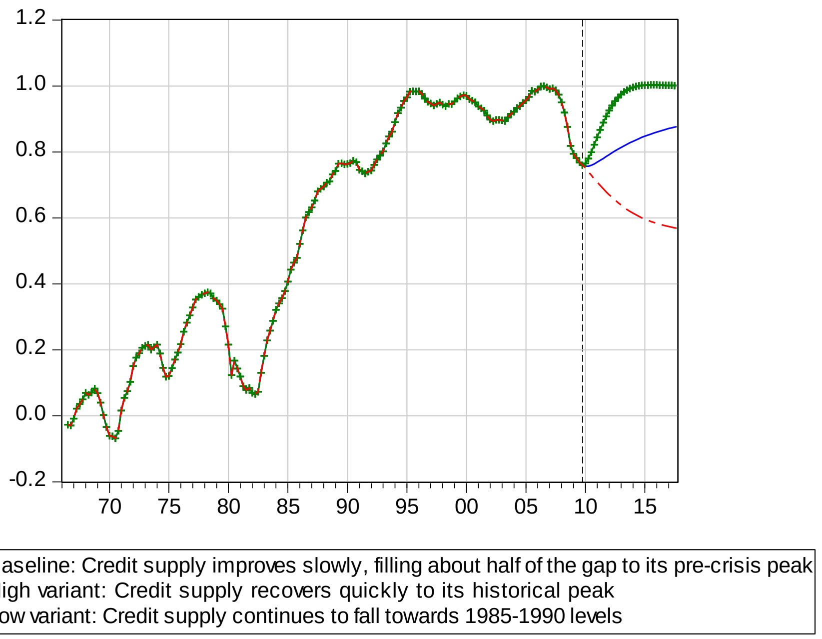
consumption plans of some households and push savings up. Overall, more than 2 percentage points of the household saving rate could be conditional on whether financial conditions are closer to the high or the low scenario.

Figure 21. US household saving rate and its sensitivity to credit conditions



Source: OECD, author's calculations.

Figure 22. The US credit supply index and alternative projections used in the scenarios



Source: Federal Reserve Board, author's calculations.

8. Conclusion and further work

59. This paper put considerable emphasis on robustness checks in order to provide clear empirical results. They support the view that financial and housing wealth matter for household consumption. Cointegration between wealth, consumption, non-property income and some necessary controls is found, using various statistical methods and functional forms.[47] Complete consumption equations are estimated in order to allow fine inference on the short-term dynamics of consumption, and prove robust to many checks such as out-of-sample forecast tests or Chow stability tests. The cases of the United States, Japan, and the aggregated euro area are treated in this way.

60. First, the estimated long-run MPCs for financial wealth are approximately of the same size in the United States, Japan, and the euro area, about 5 to 6 cents to the dollar increase in wealth. Second, the United States displays housing wealth effects of about 5 cents to the dollar, while they seem to be as low as 1 to 1.5 cents to the dollar in the euro area and Japan. Overall wealth effects, as measured by MPCs, are higher in the United States (about 5 cents to the dollar) than in euro area and Japan (3 to 4 cents to the dollar). The higher volatility of wealth observed in the United States puts additional weight on the need to take wealth into account for the US household consumption. These effects must be accounted for in short-term and long-term forecast exercises.

61. A comparison between the equations presented in this paper and the consumption equations currently used in the OECD Global Model (Hervé *et al.*, 2010) yields interesting results. In this latter, a full-log specification is used, so that estimated parameters are elasticities (see Box 1). Also, wealth enters consumption equations only through total net worth. This implicitly assumes that the different stocks of wealth affect consumption in the same way, which is proved inaccurate by the present paper. For these reasons, wealth effects of both sets of equations can be easily compared using the estimated consumption elasticities to total wealth (Figure 15). For the United States, a long-term elasticity approximately equal to 0.3 is found in this paper as well as in the Global Model. For Japan, larger effects are found here (almost 0.3) than in the Global Model equation (almost 0.2). For the euro area, the results are comparable, as long-run elasticities about 0.2 are found in both equations. Short term dynamics can also be compared. The United States show similar adjustment-to-equilibrium speed in the two equations (the error correction parameter is estimated about -0.1). Conversely, for Japan, the speed of adjustment is found significantly higher here (parameter about -0.2) than in the Global Model (parameter about -0.1). The dynamics for the euro area are very different here compared to the Global Model, where consumption is found to over-adjust to wealth changes (it increases above equilibrium and then decreases). Such features are not consistent with the results of the present paper (see Figure 15), where a moderate error correction parameter is found (about -0.1) while it is extremely large in the Global Model equation (about -0.7).

62. As considerable heterogeneity is found between the estimated wealth effects across countries, there is a good rational to use time series instead of pooled data. Future research could try to extend this approach to other countries. Better understanding the structural factors behind these differences would be of great interest. Comparing the respective roles of the banking and financial markets sectors, or controlling for risk aversion, demographics or for national pension schemes could shed additional light to these results. A more solid understanding of consumer behaviour related to wealth would help better understanding the relations between the financial sphere and the real economy, and would provide considerable weight to economic simulations, forecasts and to policy advises.

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47. See Tables A2 to A4 in the annex. They present additional evidence consistent with the above results by using full-log functional forms, and by estimating the equations either with one-step OLS method or with the Stock and Watson (1993) DOLS method. The validity of the restriction on the functional form used in the main text is also tested (and accepted) in these tables, for all "countries" at study (the United States, Japan and the euro area).

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ANNEX

A.1 Notations

63. If the contrary is not mentioned, all the variables are in real terms, deflated by private consumption deflator.

Table A1. Definition of variables

Symbol	Definition	Symbol	Definition
ln	Natural logarithm	r	Long-term interest rate[8]
Δ	Difference operator	r_s	Short-term interest rate
		r_{new}	New car loan rate[6]
c	Consumption	r_{48}	48-month consumer loan rate[5]
pop	Population	i	Nominal long-term interest rate
y	Non-property income	i_s	Nominal short-term interest rate
CC	Credit conditions index	π	Year-on-year inflation rate[4]
hw	Housing wealth	u	Unemployment rate
liq	Liquid Financial assets[1]	rp_{house}	House prices[2]
ofa	Other financial assets	*tankan*	Tankan survey[7]
fl	Financial liabilities	*ect*	Error term[3]
nfwh	*liq+ofa-fl*	*ydh*	Disposable income

1. Liquid assets are defined by the sum of households and NPISHs Deposits and currency, Treasury securities and Agency securities. The respective codes in FoF table B.100 are FL154000005.Q, FL153061505.Q and FL153061705.Q.
2. CPI deflated.
3. Residuals from the long-term regression. The error term is the gap to the long-term equilibrium.
4. Based on private consumption deflator: $\pi = ln(price/price(-4))$.
5. Finance Rate on Consumer Instalment Loans at Commercial Banks - New Autos, 48 Month Loan, Not Seasonally Adjusted, from the FRB (G19). Code : RIFLPBCIANM48_N.M (quarterlised).
6. New Car Average Finance Rate at Auto Finance Companies, Not Seasonally Adjusted, from the FRB (G19). Code: RIFLPCFAN_N.M (quarterlised).
7. Bank of Japan's Main Time-series Statistics (Quarterly), TANKAN/Judgement Survey/Lending Attitude (all industries), DI/Lending Attitude/All/All industries/Actual result, CO'COAEF0000612GCQ00000@, %points.
8. Based on government bonds yield.

Source: Compilation by the author.

A.2 Some additional evidence on cointegration between consumption, wealth and income

64. This section presents additional results linking wealth and consumption in the long-run. Overall, they prove that the broad conclusions derived from the estimation results shown in the main text are fairly robust. The first set of results refers to the United States, the second, to Japan, and the last, to the euro area. For each of them, the following three sets of results are reported in Tables A2 to A4:

1) The same short-run equation as in the main text is estimated in one step, instead of two steps in the main text. This means that instead of estimating first the long-run equation and second throwing its residuals into the short-run equation, all regressors are directly put into the short-run regression, including the integrated regressors. See the estimation results in column *"One-step OLS regression : estimated coefficients"*.[48] The implied long-run relation between integrated regressors is then shown together with the implied ECM parameter. See column *"Implied error correction term and long-run coefficients on the various long-term variables"*.[49]

2) The long-run equation is re-estimated using a full-log specification instead of the semi-log approach used until now. This means that log consumption-income ratio is still the dependant variable, but logs of wealth-income ratios are used in the regression, instead of the level of these ratios as done previously. Note that inflation, credit constraints and interest rates are not put in logs when they appear in the long-term equation. The output of this regression is reported in the column *"Results with the long-run equation estimated with logs of wealth-income ratios instead"*. As explained in Box 1, the reported coefficients for wealth are now elasticities. The corresponding MPCs were computed and reported in {} below the estimated elasticities. Below this is shown the error correction parameter obtained when the error of this full-log relation is thrown into the short-term regression as the error correction term.

3) The long-run equation, this time with the usual semi-log specification, is estimated using the Stock-Watson (1993) DOLS technique. See the column "*Estimated long-run coefficients when the long-run equation is estimated with a Stock-Watson procedure*".[50] The t-ratios of this equation are asymptotically valid for Gaussian inference in the standard Stock-Watson (1993) framework. The log of income per-capita *ln(y/pop)* is put in the long-run equation to test the restriction used throughout this paper that a 1% increase in both wealth and non-property income imply a 1% increase in consumption in the long-run. This test is accepted when the absolute t-stat on *ln(y/pop)* is below 1.96. These results validate this restriction for the United States, Japan and the euro area. Also, the long-run coefficients estimated by DOLS can be compared to the output of "Selected long-term regressions" tables in the main text: Table 3 for the United States, Table 6 for Japan and Table 9 for the euro area.

All equations are estimated including a constant term, although it is not always shown. Overall, the effects estimated here are similar to those reported in the main text, and confirm that the broad conclusions of the paper are relatively robust to the choice of the model and the estimation technique.

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48. The estimated coefficients reported in this column can be compared to the output of "Selected short-term regressions" tables in the main text: Table 4 for the United States, Table 7 for Japan and Table 10 for the euro area.

49. The estimated coefficients reported in this column can be compared to the output of "Selected long-term regressions" tables in the main text: Table 3 for the United States, Table 6 for Japan and Table 9 for the euro area.

50. The estimated coefficients reported in this column can be compared to the output of "Selected long-term regressions" tables in the main text: Table 3 for the United States, Table 6 for Japan and Table 9 for the euro area.

Table A2. Other methods to estimate the US long-run consumption

One-step OLS regression: regressors	**One-step OLS regression: estimated coefficients**[2] *Δln(c/pop) =*		**Implied error correction term and long-run coefficients on the various long-term variables:** *ln(c/y) =*	**Results with the long-run equation estimated with logs of wealth-income ratios instead. Estimated parameters for wealth are elasticities and corresponding MPCs are given in {}. Below is the corresponding estimated error correction coefficient**[1] *ln(c/y) =*	**Long-run variables**	**Estimated long-run coefficients when the long-run equation is estimated with a Stock-Watson procedure**[3] *ln(c/y) =*
Cste	-0.006	*hw/y*	0.027*CC*	0.120*CC*	*ln(y/pop)*	0.031
	[-0.8]			{0.055*CC*}		[1.6]
ln(c/y)(-1)	-0.112***	*(ofa-fl)/y*	0.024	0.029	*CC*hw/y*	0.034***
	[-3.3]			{0.012}		[6.8]
*(CC*hw/y)(-1)*	0.0030**	*liq/y*	0.087	0.127	*(ofa-fl)/y*	0.009***
	[2.1]			{0.159}		[3.4]
((ofa-fl)/y)(-1)	0.0027**	*nfwh/y*	0.222*r*	0.500*r*	*liq/y*	0.080***
	[2.5]			{0.151*r*}		[4.2]
(liq/y)(-1)	0.0099				*r*nfwh/y*	0.305***
	[1.4]	*ect(-1)*	-0.112	-0.082		[16.4]
*(r*nfwh/y)(-1)*	0.0249***					
	[2.8]					
Δln(y/pop)	0.153***					
	[4.3]					
Δu/100	-0.676***					
	[-4.0]					
Δr_{new}	-0.104***					
	[-2.9]					
Δr_{48}	-0.419***					
	[-4.1]					
Wealth change[4]	0.016**					
	[2.4]					
$\Delta\pi$	-0.853***					
	[-6.1]					
Set of dummies[4]	yes					
Sample	72Q2-09Q3			66Q3-09Q2		67Q3-08Q3
R^2	0.68					
BIC	-7.741					
SER	0.0041					
DW	2.197					
Q-test	0.092					
Q²-test	0.001					
Jarque-Bera	0.90					
Chow F test 1990	0.20					
Chow F test 1985 1997	0.25					

Note: "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.

1. See Box 1. *Estimated coefficients, elasticities and marginal propensities to consume*: the MPCs for wealth are derived from the estimated parameters using the consumption – wealth ratio from 2000 to the end of the sample. The error term from a long-term relationship with logs, when put with other benchmark regressors in the short-term equation, is still significant. But it does not change much the regression output, since all statistical properties are left roughly unchanged and the error term is very similar to the benchmark one (correlation about 0.988). **These MPCs are on wealth stocks, not their ratio to income.**
2. The dependent variable is the growth rate of consumption (per-capita). White (1980) robust covariance matrix estimator.
3. With leads and lags until order three. See Stock and Watson (1993). Newey-West (1987 and 1994) robust covariance matrix is used to compute t-statistics in brackets.
4. See Table 4.

Source: Author's calculations.

Table A3. Other methods to estimate the Japanese long-run relation

One-step OLS regression: regressors	**One-step OLS regression: estimated coefficients[3]** ***Δln(c/pop) =***		**Implied error correction term and long-run coefficients on the various long-term variables:** ***ln(c/y) =***	**Results with the long-run equation estimated with logs of wealth-income ratios instead. Estimated parameters for wealth are elasticities and corresponding MPCs are given in {}. Below is the corresponding estimated error correction coefficient[2]** ***ln(c/y) =***
Cste	-0.069***	*nfwh/y*	0.064	0.143
	[-6.1]			{0.0385}
ln(c/y)(-1)	-0.246***	*hw/y*	0.0088	0.0339
	[-6.5]			{0.0091}
(nfwh/y)(-1)	0.016***	*r*	1.058	0.551
	[6.0]			
(hw/y)(-1)	0.002***			
	[3.4]	*ect(-1)*	-0.246	-0.243
r(-1)	0.26***			
	[6.6]			
Δln(y/pop)	0.285***			
	[3.6]			
Δln(c/y)(-1)	-0.162***			
	[-3.2]			
$\Delta i_s(-2)$	0.359***			
	[4.6]			
Δi(-4)	-0.347***			
	[-5.1]			
$\Delta\pi.1_{\pi>0}$	-0.659***			
	[-4.6]			
$\Delta\pi(-1).1_{\pi>0}$	0.389***			
	[5.2]			
Δln(c/pop)(-3)	0.149***			
	[2.9]			
Δ(tankan/1000)(-1)	0.192**			
	[2.0]			
Set of dummies[1]	yes			
Sample	76Q1-08Q4			76Q1-08Q4
R^2	0.67			
BIC	-7.026			
SER	0.0057			
DW	2.254			
Q-test	0.024			
Q^2*-test*	0.047			
Jarque-Bera	0.70			
Chow F test 1993	0.640			

1. See Table 7.
2. See Box 1 *Estimated coefficients, elasticities and marginal propensities to consume*: the MPCs for wealth are derived from the estimated parameters using the consumption – wealth ratio from 2000 to the end of the sample. **These MPCs are on wealth stocks, not their ratio to income.**
3. The dependent variable is the growth rate of consumption (per-capita). Newey-West (1987 and 1994) robust covariance matrix estimator. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.

Source: Author's calculations.

Table A3. Other methods to estimate the Japanese long-run relation (cont'd)

Long-run variables	Estimated long-run coefficients when the long-run equation is estimated with a Stock-Watson procedure[1] *ln(c/y) =*	Estimated long-run coefficients when the long-run equation is estimated with a Stock-Watson procedure[1] *ln(c/y) =*
ln(y/pop)	-0.056	
	[-1.4]	
nfwh/y	0.067***	0.059***
	[10.7]	[45.7]
hw/y	0.01***	0.0062***
	[6.4]	[5.2]
r	1.222***	1.325***
	[16.8]	[17.9]
Sample	76Q1-08Q1	76Q1-08Q1

1. With leads and lags until order three. See Stock and Watson (1993). Newey-West (1987 and 1994) robust covariance matrix is used to compute t-statistics in brackets. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.

Table A4. Other methods to estimate the euro area long-run relation

One-step OLS regression: regressors	One-step OLS regression: estimated coefficients[2] $\Delta ln(c/pop)$ =		Implied error correction term and long-run coefficients on the various long-term variables: $ln(c/y)$ =	Results with the long-run equation estimated with logs of wealth-income ratios instead. Estimated parameters for wealth are elasticities and corresponding MPCs are given in {}. Below is the corresponding estimated error correction coefficient[1] $ln(c/y)$ =
Cste	0.004	π	-0.987	-0.527
	[1.0]			
$ln(c/y)(-1)$	-0.098**	*nfwh/y*	0.050	0.104
	[-2.5]			{0.0463}
$\pi(-1)$	-0.097**	*hw/y*	0.001	0.0635
	[-2.2]			{0.0147}
$(nfwh/y)(-1)$	0.005**			
	[2.0]	*ect(-1)*	-0.098	-0.0637
$(hw/y)(-1)$	0			
	[0.1]			
$\Delta ln(ydh/pop)$	0.679***			
	[6.9]			
$\Delta ln(c/pop)(-1)$	-0.219**			
	[-2.4]			
Wealth change[3]	0.021***			
	[2.9]			
$\Delta^2 ln(rp_{house})$	0.103***			
	[2.8]			
Sample	81Q1-08Q4			80Q4-08Q4
R^2	0.61			
BIC	-8.385			
SER	0.0032			
DW	2.117			
Q-test	0.467			
Q^2*-test*	0.0004			
Jarque-Bera	0.34			
Chow test 1995	0.606			

1. See Box.1 *Estimated coefficients, elasticities and marginal propensities to consume*: the MPCs for wealth are derived from the estimated parameters using the consumption – wealth ratio from 2000 to the end of the sample. **These MPCs are on wealth stocks, not their ratio to income.**
2. The dependent variable is the growth rate of consumption (per-capita). White (1980) robust covariance matrix estimator. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.
3. See Table 10.

Source: Author's calculations.

Table A4. Other methods to estimate the euro area long-run relation (cont'd)

Long-run variables	Estimated long-run coefficients when the long-run equation is estimated with a Stock-Watson procedure[1] *ln(c/y) =*	Estimated long-run coefficients when the long-run equation is estimated with a Stock-Watson procedure[1] *ln(c/y) =*
ln(y/pop)	-0.116	
	[-1.0]	
π	-0.654***	-0.821***
	[-3.8]	[-5.7]
nfwh/y	0.066***	0.051***
	[6.6]	[10.7]
hw/y	0.016***	0.011***
	[3.7]	[4.6]
Sample	82Q1-07Q4	82Q1-07Q4

1. With leads and lags until order four. See Stock and Watson (1993). Newey-West (1987 and 1994) robust covariance matrix is used to compute t-statistics in brackets. "*","**", and "***" denote significance at 10, 5, and 1% levels assuming Gaussian distribution.

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