Why net worth is the wrong concept for explaining consumption: evidence from Italy*

Riccardo De Bonis†, Danilo Liberati‡, John Muellbauer§ and Concetta Rondinelli¶

8 August 2025

Abstract

Most econometric policy models at central banks use a 'representative agent' aggregate consumption function in which net worth (summing all household assets less debt) proxies for balance sheet effects and borrowing and liquidity constraints are absent. This conventional formulation is inconsistent with heterogeneous agent behaviour under uncertainty in incomplete markets. Using aggregate quarterly Italian data from 1980-2019, our more general model rejects the above formulation. We instead find different marginal propensities to consume out of liquid and less liquid wealth components. Moreover, rising housing prices have a dual effect on consumption, increasing housing wealth for owners, and worsening housing affordability.

JEL Classification: E21, E32, E44, E51

Keywords: central bank models, consumption function, financial wealth, liquid and illiquid assets, permanent income, housing wealth.

* We are very grateful for comments from the referee and editor and wish to thank G. Albareto, A. Brandolini, G. Bulligan, F. Busetti, A. Cristini, F. Lilla, G. Marinelli, S. Neri, A. Silvestrini, R. Zizza and F. Zollino for discussions and/or access to data, and above all to Janine Aron for her extensive comments and insights. We are grateful to the participants who attended the Workshop "Household Consumption, The Role of Heterogeneity and Policies" held at Bergamo University, 3 June 2019, and those who attended the Italian Economic Association 60*th* Annual Scientific Conference held at Palermo University, 24-26 October 2019, for their useful comments. The views expressed herein are solely those of the authors and do not necessarily reflect the views of the Bank of Italy.
† Bank of Italy, Consumer Protection and Financial Education Directorate.
‡ Bank of Italy, Economics, Statistics and Research Directorate.
§ Nuffield College and Institute for New Economic Thinking, Oxford Martin School, University of Oxford.
¶ Bank of Italy, Economics, Statistics and Research Directorate.

1. Introduction

Most central banks now include 'semi-structural' econometric policy models, with scope to learn from data, in their policy modelling platforms. Consumption typically accounts for 60 percent or more of GDP in advanced economies, and hence the aggregate consumption function should have a crucial role in these models. However, even for central banks with 'semi-structural' models, most retain a specification of the consumption function based on the simple textbook form for a representative agent. This consumption function assumes that the relevant concept of wealth for households is net worth, which treats liquid and illiquid financial assets and housing as if they each have the same expenditure implications.[1] These consumption models also assume that households face no borrowing or liquidity constraints (especially relevant when there is income uncertainty).

A more general approach is needed which can relax these incredible assumptions. The probability is small that any household's asset holdings approximates the 'average' portfolio of all households, given transactions costs and differences in risks and returns. For example, heavily indebted households typically have very different asset holdings from those with low debt to income ratios. Many households have no direct stock market participation, and substantial proportions of households are renters rather than owner-occupiers. Hence when aggregate data on household portfolios is split into different asset types this reveals considerable distributional information. These different types of households are likely to have different consumption behaviours, amalgamated in aggregate consumption.

A more general formulation for the consumption function was proposed in Aron et al. (2012) and Duca and Muellbauer (2014), also see Muellbauer (2020). The textbook permanent income model was generalized by setting out a 'credit-augmented consumption function', with four important features. First, the credit channel, typically absent, is explicitly incorporated by the inclusion of 'credit conditions indices' for both unsecured credit and mortgage credit. This is because, for given levels and distributions of income and portfolios, easier lending conditions (more relaxed loan standards) will tend to raise aggregate consumer spending.[2] Second, rather than using the net worth concept, household balance sheets are split into liquid assets and debt, illiquid financial assets and housing wealth, with an additional control included for housing affordability (relevant for renting households). This allows a more realistic measurement of the different propensities to consume from the components of wealth. This formulation is supported by micro-evidence from Fagereng et al. (2021) and Crawley and Kuchler (2023), that the marginal propensity to consume out of income varies according to the structure of household portfolios.

A third feature is that rather than ignoring income expectations as in some central bank models, these have a role through permanent income. Income expectations are captured in a permanent income construct using an innovative econometric model that also adjusts for the structural breaks in the income

[1] The Federal Reserve's FRBUS, ECB BASE, the Bank of Canada's LENS and the Reserve Bank of Australia's MARTIN models are among the 'semi-structural' econometric models (SEMs) relying on net worth; the policy models at the Bank of Italy and the Bundesbank instead use 'net financial wealth', which excludes housing wealth.

[2] Easier credit also affects household portfolios, especially raising debt levels, and may raise asset prices and thus feed into future spending decisions.

process, such as were caused by the Global Financial Crisis and the subsequent European debt crisis. Since households could not have anticipated these shifts, the model takes account of over-optimism before the crises and posits a subsequent learning process in which expectations adjust to new realities. A far higher discount rate than in the standard textbook model is applied to the expected future income streams (high discount rates are also used by those central bank models which include income expectations - though they do not address structural breaks - FRBUS, ECB BASE and Canada's LENS model). Adopting a higher discount rate is consistent with the view of Friedman (1963), that in aggregate empirical applications, permanent income should be based on far shorter horizons. This approach does not claim that all households discount cash-flows identically, but it argues that it is better to control for an average effect rather than to ignore such effects altogether.

A final feature addresses a short-term role for income insecurity, proxied by the change in the unemployment rate and different measures of income volatility, and potential cash-flow effects on indebted households facing floating interest rates are captured by changes in nominal interest rates.

This paper applies the above generalized consumption function approach (that encompasses the conventional formulation), using quarterly data for Italy from 1980 to 2019. The model controls for permanent income, includes the effective interest rate on household loans, and tests for the relevance of an index of credit conditions. The model splits net worth into liquid, semi-liquid and illiquid financial assets and debt, controlling separately for housing affordability (measured by the house price to income ratio) and for housing wealth. The conventional form of the consumption function implies a set of linear restrictions on this credit-augmented consumption function. We relax these restrictions in a series of regressions, beginning with the conventional net worth formulation. Striking improvements in statistical fit and coherence are demonstrated with the sequential disaggregation of net worth, clearly rejecting the conventional formulation.

This generalized consumption function approach is also supported by qualitative insights for modelling aggregate consumption from the buffer-stock saving and heterogeneous agent literatures. Buffer-stock saving theory explains how rational behaviour under income uncertainty and liquidity constraints radically undermines the simple textbook permanent income model of consumption (Deaton, 1991; Carroll, 1992). Indeed, after the Global Financial Crisis, research has burgeoned on the behaviour of heterogeneous households operating in an incomplete market setting with asymmetric information. For example, new theory and evidence on heterogeneity comes from models of wealthy 'hand-to-mouth' consumption, presented by Kaplan and Violante (2014) and Kaplan et al. (2014), and generalized into a general equilibrium model in Kaplan et al. (2018) - with an otherwise conventional New Keynesian production and pricing side of the economy. These models describe the short-horizon consumption behaviour of income-constrained, though asset-rich, consumers, facing credit constraints and trading costs in an illiquid asset. Monetary policy conclusions are radically transformed in the 'heterogeneous agent New Keynesian' (HANK) model version of Kaplan et al. (2018)[3] as compared to the standard representative agent rational expectations life-cycle/permanent income version of the New Keynesian

[3] See Kaplan and Violante (2018) for a non-technical overview.

model.[4] A comprehensive overview in Pascal (2024) shows how theoretical developments in the heterogeneous agent literature have altered thinking about monetary transmission, since the earlier view summarized by Gertler et al. (1999). For transmission via households, the new view puts far more weight on the transmission of interest rate changes through general equilibrium effects on current income than on intertemporal substitution by households.

Heterogeneous agent theory suggests that, in principle, long time-series of micro-data on households are needed to analyse fluctuations in aggregate consumption. However, once the above generalized approach is adopted, many policy-relevant insights from the theory can be obtained from a consumption function using aggregate data with the household balance sheets separated by portfolio composition.

There is widespread disagreement about the size and nature of the influence of housing and financial wealth on household consumption. One of the contributions of this paper is to provide new evidence on these questions and the associated channels of monetary transmission, though for an economy with a very different credit market architecture than the US and many other major OECD economies. Boone and Girouard (2002) find large differences from both housing and financial wealth on consumption across OECD countries. Case et al. (2005) argue that in the US, the wealth effect for housing exceeds that of the stock market. Cooper and Dynan (2016) point to the distinction between the collateral and wealth effects for housing for several countries and acknowledge the mixed and heterogeneous nature of the evidence. There is also no consensus about housing wealth effects on consumption in Italy. These are positive and rather small according to Catte et al. (2004) and Guiso et al. (2006), and sometimes even negative (Boone and Girouard, 2002; Slacalek, 2009). Financial wealth effects appear to be stronger and more statistically significant than housing effects (Bassanetti and Zollino, 2010). While the ratio of Italian household financial wealth to GDP has increased in the last 40 years, housing wealth remains the main asset for Italian households, as in other advanced economies (see Caprara, De Bonis, and Infante, 2020).

The paper is structured as follows. Section 2 explains the theoretical and empirical frameworks and Section 3 describes the data (together with detailed online appendices). The role of the different assets in explaining households' consumption dynamics is reported in Section 4, with strong robustness checks. The results reported begin with the fully restricted standard net worth formulation. In a series of regressions, with progressive relaxation of the restrictions, striking improvements are demonstrated. Section 5 concludes. Online appendices provide information on ancillary modelling, as well as describing data transformations.

[4] The version of HANK Kaplan et al. (2018) does not incorporate endogenous asset prices, e.g. of equities and real estate, through which monetary policy also operates. See Garrida and Hedlund (2020) for a micro-founded extension of this model, in which balance sheets also matter.

2. Consumption models

The cornerstone of the standard aggregate consumption model is the classical Ando-Modigliani-Brumberg-Friedman consumption function based on the optimizing behaviour of a representative household. This has the following form:

$$c_t = \left(\gamma^* A_{t-1} + y_t^p\right) \tag{2.1}$$

Aggregate consumption (c) depends on household real net worth (A) at the end of the previous period and permanent non-property household disposable income (y^p). The marginal propensities to consume out of these are, respectively, $\gamma^*$and one. Non-property disposable income (y) refers to labour income plus net transfers, both after tax. In the basic version of the theory, $\gamma^*$is the return on net worth (embodying property income such as dividends, rent and interest). Then it is appropriate to exclude property income from the measure of household income used in the consumption function (Blinder and Deaton, 1985). All variables are real per capita levels.

With some manipulation, see Aron et al. (2012), a log approximation of equation (3.1) becomes:

$$\ln c_t/y_t \approx \alpha_0 + \gamma A_{t-1}/y_t + \varphi E_t \ln\,(y_t^p/y_t) \tag{2.2}$$

where the log consumption to income ratio, $\ln c_t/y_t$ is driven by three elements. First, there is (time-invariant) autonomous consumption, α_0. Second, there is the ratio of household net worth (at end of period t-1) to income, A_{t-1}/y_t, scaled by the marginal propensity to consume out of assets ($\gamma \approx \gamma^*$). Third, there is the log ratio of permanent to current non-property income, $E_t \ln\,(y_t^p/y_t)$, scaled by (φ) to allow for a generalisation of the strict permanent income hypothesis under which $\varphi = 1$. $\ln\,(y_t^p/y_t)$ can be approximated by equation (2.3) where k is the time horizon and δ is the rate at which households discount future income[5]:

$$\ln\,(y_t^p/y_t) = \left(\sum_{s=0}^{k} \delta^s \ln y_{t+s}\right)/\sum_{s=0}^{k} \delta^s) - \ln y_t \tag{2.3}$$

With partial adjustment,[6] the dynamic version of equation (2.2) is:

$$\Delta \ln c_t \approx \lambda(\alpha_0 + \gamma A_{t-1}/y_t + \varphi\, E_t \ln\,(y_t^p/y_t) + \ln y_t/c_{t-1}) \tag{2.4}$$

where λ is the speed of adjustment to a target defined by the right-hand side of equation (2.2). Note that equations (2.2) and (2.4) are homogenous of degree one, so that doubling income and assets doubles consumption. Equation (2.4) assumes constant real interest rates, no distributional effects, homogeneous net assets so that γ is the same for all asset types including liquid and illiquid financial

[3] Campbell (1987) shows that this is equivalent to a weighted moving average of forward-looking income growth rates.
[6] Partial adjustment can be rationalised through costs of adjustment, habit formation or sticky information, Reis (2006).

wealth and housing wealth; and that there are no credit constraints.[7]

To introduce controls for time-varying credit access and the real interest rate and to generalise the role of liquid, semi-liquid and illiquid assets, debt and housing wealth and affordability, we augment equation (2.4) as follows:

$$\Delta \ln c_t = \lambda(\alpha_0 + \alpha_c CCI_t + \alpha_1 r_t + \gamma_1 LA_{t-1}/y_t + \gamma_2 LOANS_{t-1}/y_t + \gamma_3 SLFA_{t-1}/y_t + \gamma_4 ILFA_{t-1}/y_t + \gamma_5 HA_{t-1}/y_t + \gamma_6 \ln hp_{t-1}/y_t + \varphi\, E_t \ln (y_t^p/y_t) + \ln y_t/c_{t-1}) + \ other\ factors \quad (2.5)$$

Here, CCI is an index of non-price credit conditions index, or loan standards[8]; r is the effective real household borrowing rate; LA_{t-1}/y_t is the ratio to income of the most liquid assets, always including bank and saving deposits; $LOANS_{t-1}/y_t$ is the ratio to income of debt; $SLFA_{t-1}/y_t$ is the ratio to income of semi-liquid financial assets; $ILFA_{t-1}/y_t$ is the ratio to income of illiquid financial assets; HA_{t-1}/y_t is the ratio to income of gross housing assets; and $\ln hp_{t-1}/y_t$ is the log ratio of real house prices to income. In most specifications, after testing, we constrain liquid assets and debt to appear in a net liquid assets form, with equal and opposite signed coefficients, i.e. assuming $\gamma_2 = -\gamma_1$. This reflects evidence also from France (Chauvin and Muellbauer, 2018), the UK and the US (Aron et al., 2012) and South Africa (Aron and Muellbauer, 2025a), which support this constraint. As it turned out, for Italy, unlike France, the UK, the US, South Africa and Germany, no long-term effect could be found for our indicator of credit conditions, which was found to be relevant only for short-term dynamics.[9]

The other long-term factors, discussed further below, include a shift in seasonality in 1990 and a 2012 shift dummy to control for the Fornero pension reforms.

We now explain equation (2.5) in more detail. Whereas equation (2.1), (2.2) and (2.4) assumed homogenous assets, equation (3.5) adopts a five-part disaggregation of household net worth[10] whereby the marginal propensity to consume (MPC) out of assets depends on the liquidity of financial assets (see Otsuka 2004 for an optimizing model with trading costs) and is different for housing (see Berger et al. 2018 for an optimizing model of the housing collateral effect in the presence of credit constraints).

An important feature of this formulation of the long-run drivers of consumption is consistency with

[7] In an ad hoc extension of the model, one can add a term $\Delta \ln\ y_t$ to represent a fraction of households who just spend income perhaps because they are credit constrained. However, as Deaton (1991) makes clear, there is no micro-foundation for such an interpretation of credit-constrained behaviour.

[8] Where loan standards have evolved differently for mortgage and non-mortgage credit, an index for each is needed.

[9] The CCI term controls for fluctuations in long run $\ln c/y$ due to shifts in household credit access, particularly mortgage credit access. Due to asymmetric information, lenders tend to rely on non-price constraints such as down-payments ('skin in the game'), debt servicing capabilities and other screening technologies, that in many countries have evolved through time, to mitigate default risk. These non-price constraints can vary with the capital position of the lenders, with regulatory constraints they face, with competitive pressure from other lenders and with their risk appetite.

[10] This is a more refined disaggregation than adopted in Aron et al. (2012) and in Chauvin and Muellbauer (2018), where semi-liquid assets were combined with illiquid financial assets.

micro-evidence and buffer-stock theory that the marginal propensity to consume out of current income varies according to the asset position of households. As Crawley and Kuchler (2023) show, the MPC out of current income is highest for asset-poor households, intermediate for those holding illiquid but not liquid assets, and lowest for the doubly asset-rich. That is consistent with equation (2.5) which implies the following expression for the MPC out of current income of an individual household:

$$\frac{\partial c_t}{\partial y_t} = (c_t/y_t)[1 - \varphi - \gamma_1 LA_{t-1}/y_t - \gamma_2 LOANS_{t-1}/y_t - \gamma_3 SLFA_{t-1}/y_t \\ -\gamma_4 ILFA_{t-1}/y_t - \gamma_5 HA_{t-1}/y_t - \gamma_6 \ln hp_{t-1}/y_t] \tag{2.6}$$

For a given value of φ which captures how forward-looking is a household, one without financial assets or housing wealth but with debt, will have the highest MPC as γ_2 is negative. This is reinforced by the fact that such resource-poor households will tend to have a high average propensity to consume, c/y. A household with no debt but with both financial assets and housing wealth will have low MPCs, reinforced by the fact that such households will tend to have lower values of c/y. Moreover, as γ_1, the MPC for liquid assets, exceeds the MPCs for less liquid assets and housing, where the fraction of the portfolio in liquid assets is high, the MPC will be especially low.

Credit availability potentially also has an influence on the MPC. Since easy access to credit should make households more forward-looking as they can more easily borrow in advance of higher expected income, the MPC out of current relative to permanent income would be expected to be lower when credit access is easy or where households have the means to meet down-payment constraints on access to a mortgage. However, this could be somewhat offset by a tendency of such households to have higher values of the average propensity to consume, c/y.

Differences in MPCs with respect to wealth between different asset types, i.e. the γ_i, can also reflect the different characteristics of households dominating ownership of the different assets. Indeed, as noted earlier, there is a good deal of distributional information buried in the aggregate data on the different types of assets. It is likely that many households with mortgage debt have relatively low holdings of deposits and other financial assets. This would make their reaction to higher debt levels more negative than implied by the net worth formulation. As it is likely that the MPC for illiquid financial assets such as stock market wealth is lower for very wealthy households, the fact that such wealth is very unequally held is likely to be a factor affecting the estimated magnitude of γ_4 on aggregate data, in addition to its illiquidity characteristics such as trading costs or riskiness.

In principle, the model should include the perceived present value to households of the generous Italian state pension in the illiquid financial asset category. Since we do not have such data, all we can do is control for likely shifts in that present value. As the 2011 Fornero pension reforms, tightening the rules, would have reduced the perceived present value of pension benefits for many households, the Fornero dummy can be interpreted as a kind of wealth effect.

Even without liquidity issues and credit constraints, intertemporal optimisation shows that, because housing is a consumption good as well as an asset, the response of consumption to an increase in house prices is different from the response of consumption to, for example, the share price index (see Aron et al. (2012), Buiter (2010) and Calomiris et al. (2009)). This is one motivation for including in the consumption function the log ratio of house price to income as well as the housing wealth to income ratio. The other reason arises because of mortgage credit constraints, particularly salient in Italy, and because owner-occupiers and renters are likely to have different reactions to housing market developments. For a given loan standard, higher house prices relative to incomes increase the need to save for a deposit to enter the mortgage market. Moreover, renters are likely to see an increase in house prices relative to incomes as sign that future rents will rise and therefore be more cautious in their spending.

It is possible that some of the coefficients in the model could interact with non-price credit conditions. For example, Aron et al. (2012) and Duca and Muellbauer (2014) find that in the US and UK, where home equity withdrawal is common, the MPC for housing wealth increases with easier non-price credit conditions. This finding is corroborated for South Africa where home equity withdrawal is also widespread, see Aron and Muellbauer (2025a). But this is not the case in France and Germany where equity withdrawal has scarcely existed, see Chauvin and Muellbauer (2018) and Geiger et al. (2016). In this respect, Italy should be like France and Germany, as mortgage equity withdrawal is also absent. It is possible that if lenders ease the down-payment requirement, the negative effect of higher house prices relative to incomes could moderate. If access to credit in general increases, households should be able to engage in more intertemporal substitution, which would increase the weight φ on permanent income. These are testable propositions.

An important advantage of the general formulation of the consumption function in (2.5) is that it nests conventional models. With $\alpha_c = 0$, and $\gamma_i = \gamma$, for i=1, 5, and $\gamma_6 = 0$, equation (2.5) reduces to equation (2.4).

3. Data background

3.1 Consumption and income

The Italian ratio of total consumer spending to household disposable income has experienced a significant evolution. Before about 1993 it was rarely above 80 percent, implying a household saving ratio of over 20 percent, well above the saving ratios for most OECD countries (Ando et al., 1994 and De Bonis and Marinucci, 2023). In the later 1990s, the ratio rose to levels nearer 90 percent, and saving ratios moved closer to those in OECD countries after Italy joined the Eurozone in 1999. The ratio remained fairly constant until the Covid-19 pandemic, when it declined dramatically, due to a huge drop in Italian private consumption – the sharpest in the euro area together with Spain. The contraction in disposable income was somewhat smaller as it was sustained by a generous stimulus package implemented by the government (Guglielminetti and Rondinelli, 2023).

Textbook consumer theory suggests that non-property income rather than overall disposable income is the appropriate income concept for explaining consumption, with net worth capturing permanent property income. We define a measure of scaled income, which averages labour income plus transfer income, and overall disposable income (including property income), as our preferred income measure to approximate non-property income (see the Appendix A.1 for details on data sources and definitions). Rather than give labour and transfer income a weight of 1, with property income replaced by a fixed rate of return multiplied by net worth, we overweight labour and transfer income, giving property income as measured in the national accounts a reduced weight in explaining consumption. Several reasons underly this compromise. First, the text-book assumption that property income is a fixed proportion of net worth is quite unrealistic, as rates of return vary by property type and are far from constant. Second, in a context where many households are credit or liquidity constrained, the cash-flow from property income matters to them, rather than a real rate of return applied to the underlying asset. In the national accounts, many types of property income are recorded as cash flows, which are then deflated by the consumption deflator to convert to real terms. Many economists regard that as mismeasurement (especially in the context of price inflation) in contrast to a measure of real property income defined as the real rate of return (the nominal return on the asset minus the expected rate of inflation) multiplied by a constant price measure of the asset value, e.g. Jump (1980). However, our view is that for many households, the standard national accounts cash-flow treatment is more relevant. The third argument for our compromise is that some types of income, e.g. those of small business owners, contain a difficult-to-separate mix of labour and property income, and moreover, national accounts data do not permit separate tax adjustments for different types of income.

Figure 1 shows the log ratios of consumption to disposable and scaled income, taken as deviations from their values in 1996Q1.[11] The figure shows that they moved closely together from the late 1970s to the mid-2000s but then diverged as interest rates fell and property income as measured in the national accounts declined relative to labour and transfer income.

[11] This reference date of 1996Q1 reflects when it became clear that Italy qualified for and would adopt the euro.

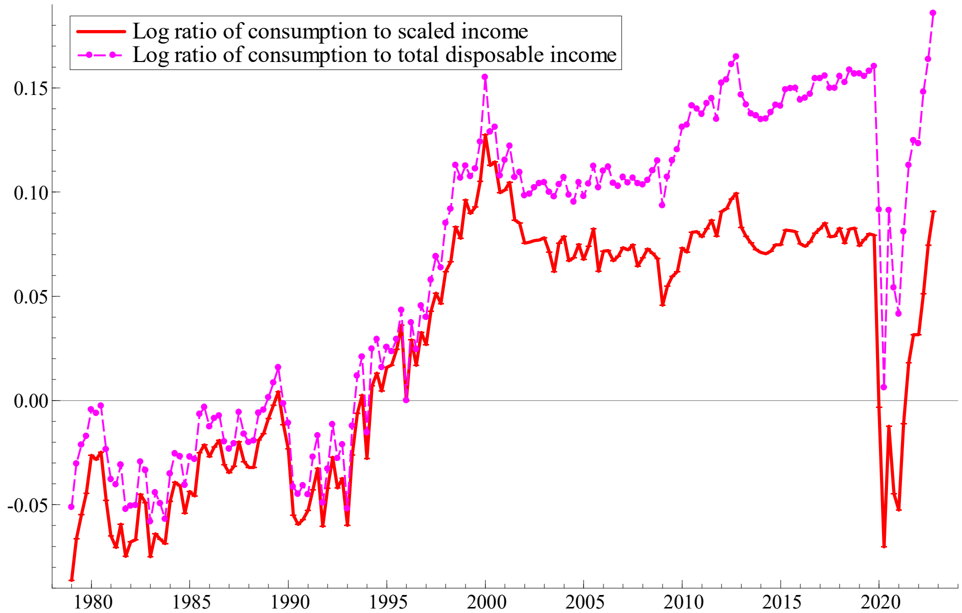

Figure 1: The log ratios of consumption to disposable and to scaled income, relative to 1996Q1.
Source: National Accounts, ISTAT, and authors' weighting of raw data.

3.2 Household balance sheet data

From 1995, quarterly financial balance sheets from the Bank of Italy give a detailed breakdown of household portfolios. Before 1995, annual balance sheets from Bonci and Coletta (2008) give a cruder breakdown in which short ('bills') and medium and long-term debt securities ('bonds') are aggregated. The split of total debt securities into bills and bonds before 1995 is based on information collected by the Bank of Italy. After quarterly interpolation of the data (see details in Appendix A.2) and adjusting for data breaks[12], we obtain continuous quarterly household balance sheets for eight categories of financial assets and debt. These are deposits plus cash, bills, bonds, quoted equities, mutual funds, unquoted shares, life insurance reserves and pensions, and household debt in the form of loans.

For empirical modelling on time series data, it is desirable to group the nine elements of the portfolio into broader groups to avoid overfitting. In our preferred empirical formulation, as explained below, we group these eight categories into three groups: net liquid assets (defined as deposits including cash, plus bills plus mutual funds, minus debt); semi-liquid assets (defined as bonds); and illiquid assets (defined as quoted shares plus pension and life-insurance assets). Figure 2 displays ratios of the market values of these three asset groups, measured at the end of the previous quarter, and divided by annualised scaled income.

[12] Several structural breaks were adjusted for, see the Appendix A.2 for details. First, the 1995 European System of Accounts (ESA95) introduced methodological innovations which affected the continuity of the time series of national accounts. Second, corrections had to be made for a break in the pensions data in 1984 and in the loans data in 1989.

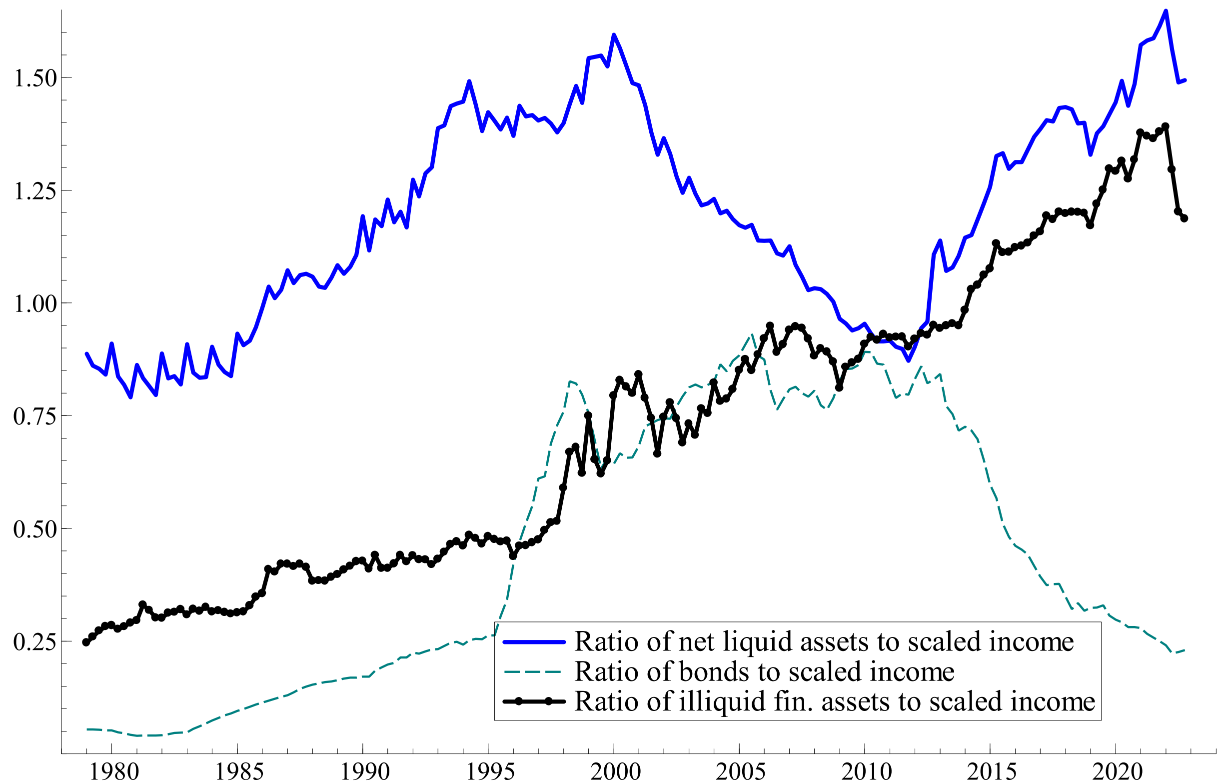

Figure 2: Ratios to income of net liquid assets, semi-liquid assets and illiquid assets.
Source: National Accounts and quarterly household balance sheets, ISTAT. Before 1995, balance sheet data linked with break-adjusted annual data are from Bonci and Coletta (2008) and internal Bank of Italy data. Net liquid assets are defined as deposits plus mutual funds plus bills minus loans; semi-liquid assets are defined as bonds; and illiquid financial assets are defined as quoted shares plus pension and life insurance funds. Unquoted shares are excluded.

The rationale behind grouping the eight portfolio elements into the three groups shown in Figure 2, is influenced by the liquidity characteristics of Italy's different assets. *A priori*, it is plausible that unquoted shares and pension wealth should be the least liquid wealth components, followed by quoted shares. Bills, bonds and mutual funds could be an intermediate semi-liquid category, with deposits the most liquid of all. However, in Italy mutual funds are mainly bank-managed funds, containing a mix of money market instruments as well as less liquid securities (Albareto et al.2020). These come in a variety of forms, including some partially invested in quoted equities. With low withdrawal restrictions, they have a high degree of effective liquidity making them a candidate for inclusion in the liquid category. Short-term treasury bills carry little capital risk and have very low transactions costs, which also makes them suitable for inclusion in the liquid category. Bonds make up the semi-liquid category, and quoted shares plus pensions and life insurance reserves comprise the illiquid financial asset category. We also consider an alternative grouping in which deposits define liquid assets, the semi-liquid category consists of bills plus mutual funds, and the illiquid group contains bonds, quoted shares and pension wealth, while unquoted shares drop out altogether.

In all plausible groupings, unquoted shares have a negative coefficient, suggesting serious problems in valuation, and are therefore omitted. Liquid assets minus loans can be combined into 'net liquid assets' in both of the plausible groupings.

Figures 3 to 5 show ratios to annualised scaled income for the full available disaggregation of nine categories of balance sheet data. We summarise the main factors driving variations in these ratios. Notable features of the data include the surge relative to income in holdings of bonds and mutual funds after 1995, while holdings of bills –short-term securities- as well as of bank deposits, declined as certainty increased regarding Italy's membership of the currency union and price inflation fell (see Caprara, De Bonis and Infante 2020). Holdings of mutual funds decreased after the collapse of the dotcom stock market boom in 2001, which also sharply lowered the value of quoted shares, and bottomed out in the financial crisis of 2008-9, but rose again with the low interest rates and stock market recovery that followed. For net liquid assets, as defined above, the rise in the value of mutual funds in the 1995 to 2000 period was offset by the relative decline in deposits and bills and the rise in household debt. The decline in the ratio of net liquid assets to income from 2001 to about 2012 is substantially due to the relative decline in mutual funds after the dotcom bust, to a decline in holdings of short-term bills, and to a degree, to the rise in debt relative to deposits. Subsequently, the rise in the ratio of net liquid assets to income is due to the recovery of mutual funds and the rise in deposits relative to debt.

Turning to semi-liquid assets, i.e. bonds, there was a striking rise after 1995 with lower inflation, lower interest rates and increasing confidence that Italy would join the euro, broadly sustained in the relatively low interest rate period until the sovereign debt crisis, after which the value of holdings of Italian long bonds fell sharply. For the composite of illiquid financial assets, the gradual upward trend of holdings of pensions and life-insurance funds looks relatively unrelated to economic events and dominates the much more volatile value of holdings of directly-held quoted shares. That showed a strong rise in the late 1990s boom, with a temporary hiccup in the financial crisis affecting emerging markets and Russia in 1998-9, followed by a sharp decline in the dotcom bust, broadly sustained in the subsequent years.

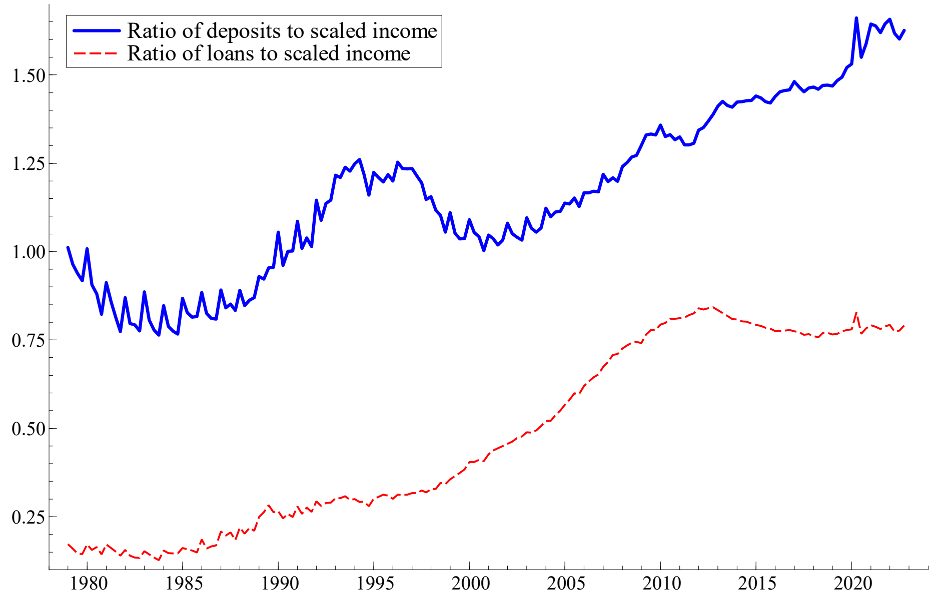

Figure 3: Ratios to scaled income of household deposits and loans.
Source: National Accounts and quarterly household balance sheets, ISTAT. Before 1995, balance sheet data linked with break-adjusted annual data are from Bonci and Coletta (2008) and internal Bank of Italy data.

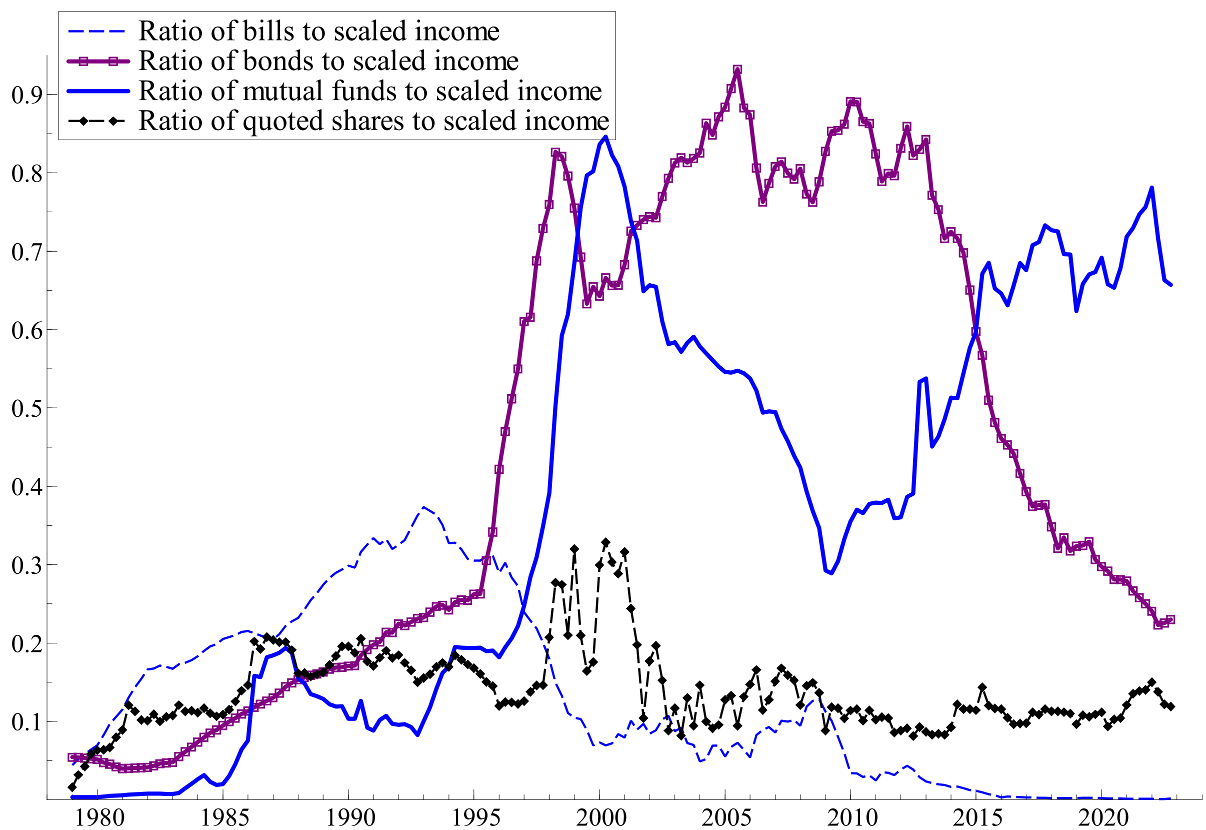

Figure 4: Ratios to scaled income of bills, bonds, mutual funds and quoted shares.
Source: National Accounts and quarterly household balance sheets, ISTAT. Before 1995, balance sheet data linked with

break-adjusted annual data are from Bonci and Coletta (2008) and internal Bank of Italy data.

The estimates of housing wealth including land are obtained from Istat from 2001 and are linked with estimates by Cannari et al. (2017). Figure 5 plots the ratios to scaled income of housing wealth, unquoted shares and pension wealth, showing the dramatic scale of housing wealth relative to other wealth components.

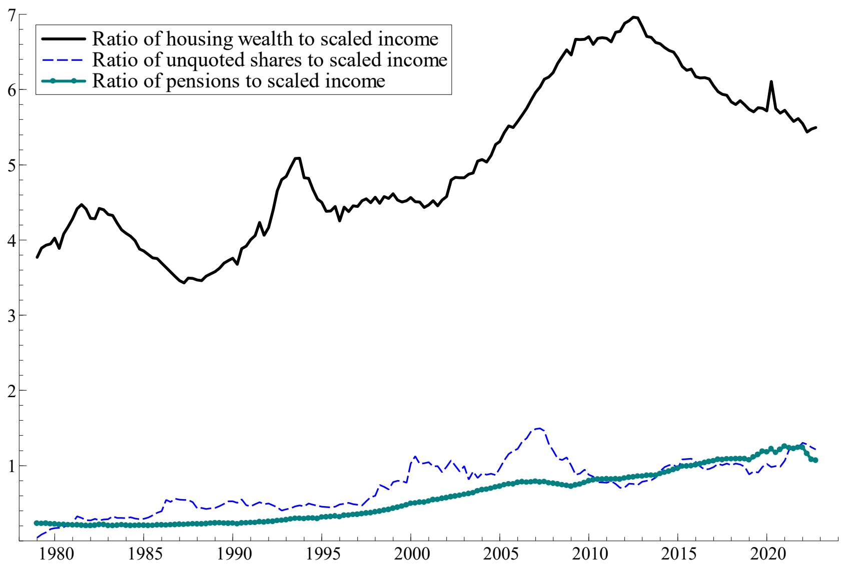

Figure 5: Ratios to scaled income of housing wealth, unquoted shares and life insurance and pension wealth.

Source: National Accounts and quarterly household balance sheets, ISTAT. Before 1995, balance sheet data linked with break-adjusted annual data are from Bonci and Coletta (2008) and internal Bank of Italy data.

House prices affect not only housing wealth but also housing affordability which we measure by the ratio of house price to per capita income, see Figure 6. Unsurprisingly, the ratio has a cyclical pattern like that of the housing wealth to income ratio but does not show the upward trend of the latter (which reflects the cumulative effect of real residential investment). A comparison of Figure 5 with the index of credit conditions discussed below, see Appendix 4, Figure A3, suggests that shifts in credit conditions help to explain the cyclical pattern of the house price to income ratio. As in a number of European countries, house prices rose relative to income in the late 1980s and declined in the early 1990s, with the ERM crisis. Interestingly, house prices did not join the boom in prices of financial assets in the late 1990s, but rose strongly after Italy joined the euro. With the GFC and the sovereign debt crisis, the Italian house price to income ratio fell sharply, and failed to join the post-crises recovery in house prices experienced in most other European countries.

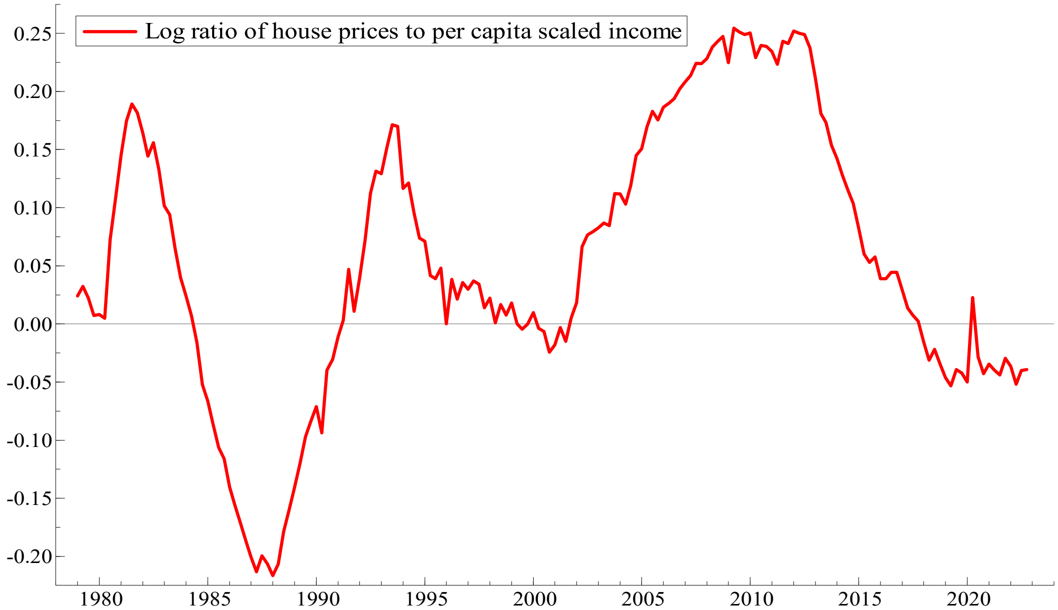

Figure 6: The log ratio of house prices to income, relative to 1996Q1.
Source: National Accounts, ISTAT and the house price index from Bank of Italy.

While major evolutions in aggregate balance sheets have occurred, it is important to recognize the heterogeneity of ownership of these assets. In Italy, financial wealth is very concentrated, with almost 80 per cent of financial wealth belonging to the upper 40 per cent of the income distribution. As shown in the Appendix A.3, higher income households are far more likely to be holders of the different categories of wealth. Indeed, the proportion of households who own no wealth in a particular asset class varies strikingly by income quantile. Unsurprisingly, deposits are the most universally held type of asset, though even for this category, poorer households have far lower holdings than higher income households. As noted earlier, the MPC measures depend not only on liquidity characteristics of the different assets but also on the behaviour patterns of the dominant owners.

3.3 Estimates of permanent income

Permanent income is prominent in the classic texts and important in our generalized consumption function. We define the log of permanent income as the discounted stream of log income over the next 40 quarters, using a quarterly discount factor of 0.95, implying an annual discount rate of a little over 20 percent. Following Chauvin and Muellbauer (2018) and Geiger et al. (2016), we use a reduced form forecasting equation for log permanent income.[13] One independent variable is the log ratio of the working age population measured relative to the total population, since our consumption and income data are per capita: a higher proportion in the working age group implies higher per capita

[13] The forecasting method used is a type of 'local projection', Jordà (2005). Instead of forecasting quarterly log income separately at each horizon t+h, where h runs from 1 to 40 quarters, we forecast the weighted average of quarterly log income over a horizon of 40 quarters (see further details in Appendix A.6).

incomes. Other drivers in levels are a measure of international competitiveness, the unemployment rate, the log of real oil prices, logs of current and recent income. and the log of the real stock market index. For the dynamics, changes are included in the unemployment rate, in log real oil prices and in a credit conditions indicator (see Appendix A.6). In the long run, real per capita incomes trend upwards with improvements in technology and working practices, and hence a time trend is also included.

Our forecasting model for the permanent income construct requires adjustment for unexpected structural breaks through learning-adjustment. The Global Financial Crisis (GFC) of 2008-2009 put a serious and unforecastable dent into the long-run performance of income. This is manifested as a mix of a one-off step down in income and a lower level of trend growth. Households could not have foreseen the GFC, and to address this, the fitted values of the forecasted log permanent income are adjusted beginning 40 quarters earlier to remove the unanticipated break in the income process. We aim to replicate what a forecaster would likely have predicted prior to the GFC. According to our estimates, without any adjustments, the permanent income assessments just before the crisis would have been around 3.5 percent too optimistic. We thus build in a mix of quick and gradual learning from 2008Q3. We assume that, given the severe shock of the financial crisis, half of the learning adjustment was instantaneous. As the European sovereign debt crisis followed the GFC with some delay, we then assume there are two phases of a gradual learning process, with equal weights: the first running from 2008Q3 to 2010Q2 and the second from 2010Q3 to 2012Q2. From 2012Q2, the learning-adjusted forecasts are fully based on the model incorporating the structural break, see Figure 7. The learning-adjusted version demonstrates the over-optimism before the GFC and the relatively rapid downward adjustment thereafter. By contrast, the rational expectations approach, which ignores the large forecast error everyone made just before the crisis, would not have been a good proxy for realistic income expectations.

The forecast method uses actual data on future income for 40 quarters ahead, where available. As the estimation period ends in 2019Q4, this means that from 2020Q1 we need forecasts of income from some external source to fill in the missing values. We take these from income growth forecasts for the following 40 quarters from vintage forecasts made at the end of 2019 by *Oxfordeconomics.com*. These forecasts correctly do not take account of the imminent unforecastable pandemic. Robustness is checked by comparing the results where these forecast growth rates are reduced by 50 percent.

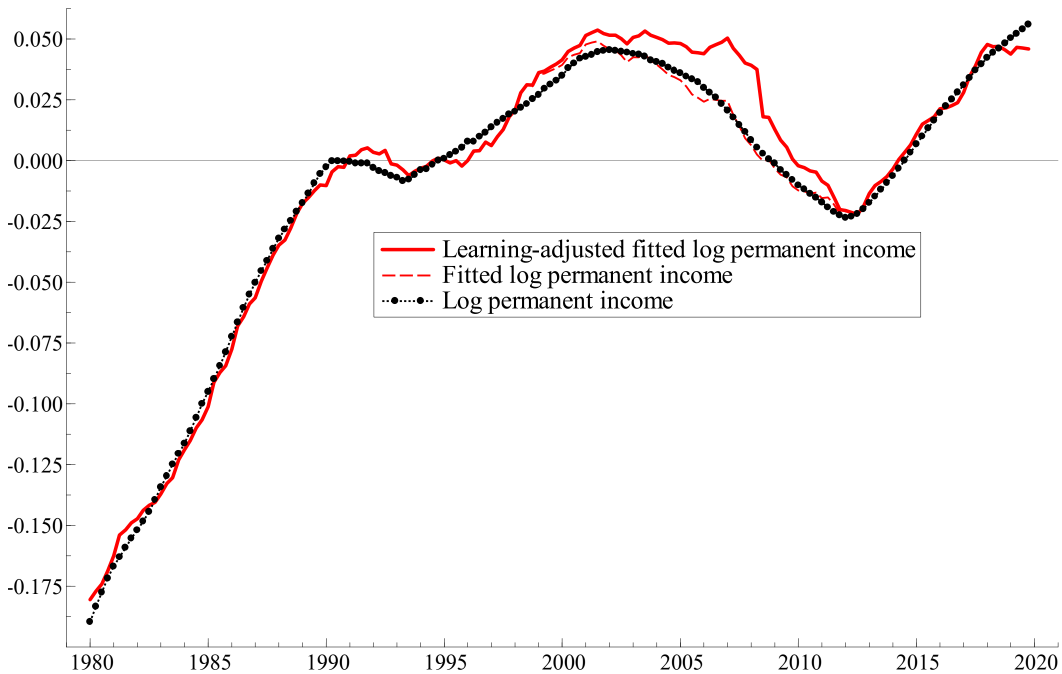

Figure 7: Estimates of log permanent income: actual values, fitted values from an econometric model, and learning-adjusted fitted values, relative to fitted value in 1996Q1.
Source: National Accounts, ISTAT and learning-adjusted permanent income estimates described in the text.

3.4 Interest rates

Further data relevant for the consumption function's long-run solution are interest rates. Textbook theory suggests a negative coefficient for a real interest rate in the consumption function, interpreted in terms of intertemporal substitution. However, there are conditions under which a positive coefficient can occur, see Aron et al. (2012). To define the nominal effective interest rate, we adapt the Bank for International Settlements (BIS) approach to the measurement of the household debt service ratio, where Drehmann (2017), defines the relevant interest rate for an amortising mortgage with a remaining life of 18 years as follows:

$$adjR = R/(1-(1+R)^{-18}) \tag{3.1}$$

where R is the annual percentage mortgage rate/100. We instead take 12 years as a better approximation to the average maturity of Italian household debt. In the equation, $adjR$ is the interest rate for which the stream of payments on debt service and repayment would be constant over time. We define the real interest rate by subtracting the 4-quarter change in the log of the consumer expenditure deflator. Figure 8 shows that the real rate remained positive even during the high inflation period in the 1980s, turning briefly negative only in the inflation surge following Russia's attack on Ukraine.

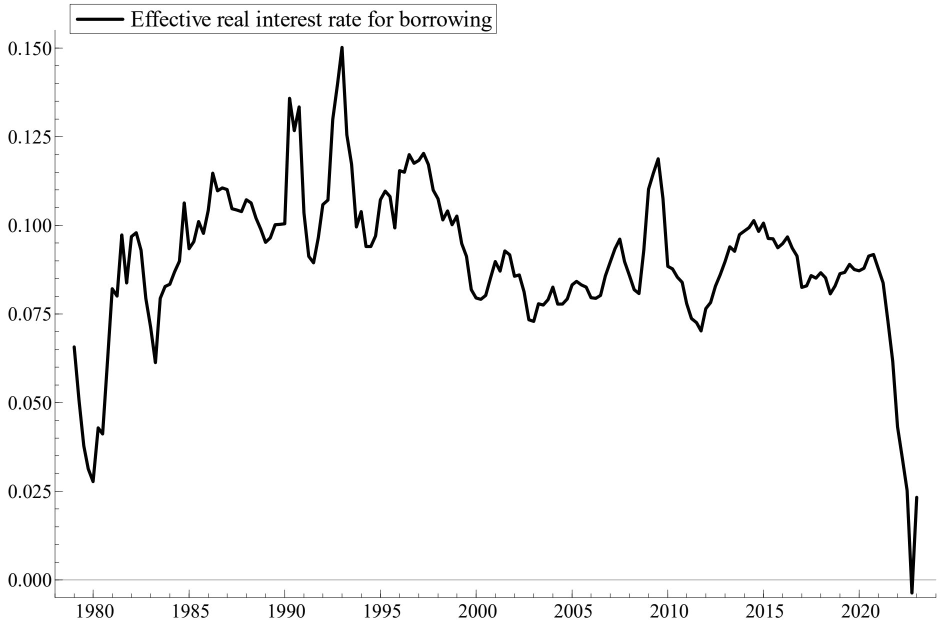

Figure 8: The real interest rate on housing loans.
Source: The interest rate is for housing loans, from Bank of Italy; the BIS definition is used for the effective rate on a loan but with 12 years remaining duration, see text; the real rate is the nominal effective rate minus the 4-quarter change in the log consumer expenditure deflator.

3.5 Credit conditions

As shown in Figure 3, the ratio of household loans to income has drifted up since the 1980s, and especially from the late 1990s to the onset of the GFC. As the introduction and the model set-up in Section 2 have emphasised, controlling for variations in non-price credit conditions (credit standards) is important for modelling consumption, especially in countries where liberalization of credit conditions occurred through financial deregulation, and through shifts in financial technology and industry structure. The ratio of household debt to income in Italy has risen over time, but it is not clear to what extent credit liberalization played a role as opposed to rising income and house prices, and lower interest rates. Unfortunately, data on the composition of household debt (mortgage versus non-mortgage debt) only begin in 1999. We therefore resort to a measure of overall credit conditions for the economy rather than one specific to households. Our measure is based on data on credit lines available to the private sector relative to a moving average of GDP, see Appendix A.4 for details.

3.6 Seasonality, structural breaks and short-term dynamics

While the consumption and income data are meant to be seasonally adjusted, there is a pronounced shift in seasonality in the recorded consumption data before 1990, especially in the third quarters.

We control for this with a third quarter seasonal dummy which is zero from 1990Q1.

Finally, there are structural breaks controlled for by dummies. The long run solution for the consumption to income ratio is likely to be affected by pension reform. While the 1992 Amato and 1995 Dini reforms, see Hamann (1997), appear to have had few detectable long-run effects on aggregate data, this is not so for the Fornero reforms passed in 2011, which reduced the generosity of pension benefits (Fornero, 2020). These reforms succeeded the sovereign debt crisis, which had deeply affected Italy. The reforms came into effect in 2012 and appear to have lowered the consumption to income ratio, see Lepinteur et al. (2022).[14] We therefore introduce a shift dummy in 2012, but allow gradual phasing in of its full effects over two years. A softening of the Fornero reforms was anticipated from 2017 with the election of a new government seen to be likely and eventually realised in May 2018. Checking for consumption effects of a softening of the pension reforms uses a 2017 shift dummy, also phased in over two years (though this proved to be insignificant).

Four dummies capture outliers in the consumption function. These are negative impulse dummies for 1983Q1, 1993Q1 and 2013Q1, and a dummy, which is 1 for the two years: 1992-93. In 1983Q1 the spread between the Italian and German T-bill rates reached a record. Consumer confidence was probably affected by political uncertainty with the forthcoming election. The years 1992-93 saw an exchange rate crisis resulting in Italy's temporary departure in 1992Q3 from the Exchange Rate Mechanism (ERM), accompanied by a record jump in the T-bill spread. Measures of consumer confidence available since 1985 show the most pessimistic view of economic conditions and of employment prospects in 1993Q1, except for the sovereign debt crisis in 2012.[15] The 2013 impulse dummy reflects the large tax rises of that year including the temporary re-introduction of a property tax.

Controls for short-term dynamics are discussed in the Appendix A.4 and A.5. These include the growth rate of an indicator of credit conditions, the acceleration of income growth, a measure of income volatility, changes in the unemployment rate and in the log of per capita constant price public consumption – a partial substitute for private consumption.

4. Empirical findings.

4.1 Main regressions

[14] Lepinteur et al. (2022) use a difference-in-differences framework based on a firm-size discontinuity, with individual data coming from the Italian Survey on Household Income and Wealth, to examine the consequences of the 2012 reforms. Their evidence suggests that the resulting greater income insecurity reduced consumption and increased savings.

[15] It is also possible to incorporate such expectations effects into the model for permanent income. However, this is quite complex, given a downward shift in the underlying trend in 1992, and an upward shift in 1994, and corresponding learning mechanisms. As the 1992 crisis was temporary, introducing a 1992-3 dummy into the consumption function was the simpler alternative.

Tables 1 and 2 give the main findings for our estimation of a more generalized consumption function for Italy, that encompasses the conventional textbook version. In Table 1, we follow a step-by-step approach, progressively splitting the conventional measure of net worth into its liquid and illiquid asset components and introducing debt.

The first regression (in column 1) captures the simple textbook version of the consumption function, where every asset that households own is summarised by one variable: the log of net worth to income ratio. There is no distinction between financial and physical assets in the long run and in the short-term dynamics. Without most of these controls, there is almost no relationship to be found, suggesting quite erroneously that consumption might well follow a martingale process. Column 1 shows a poor fit of the model with high levels of residual autocorrelation, as shown in the Durbin-Watson statistic and tests of autocorrelation up to 4 quarters, and poor parameter stability revealed in the Chow test.[16] The speed of adjustment is very low at 0.05 (a typical sign of misspecification).[17] The log of net worth is only just significant, and one could accept the hypothesis of a unit coefficient on permanent income (as implied by the textbook model). Column 2 adds the lagged change in the log of credit conditions to the short-term controls. Though outside the framework of the standard model, it is significant, raises the speed of adjustment a little and improves the fit and the significance of log net worth. However, parameter stability remains very poor and residual autocorrelation very high. Column 3 demonstrates that replacing the log of net worth simply by net worth causes a deterioration in the already poor results. All three specifications so far suggest a *positive* sign on the real effective interest rate on borrowing, contrary to conventional wisdom.

It is crucial to disaggregate net worth. Column 4 shows a radical improvement when housing wealth is separated from net financial assets. The speed of adjustment rises to 0.10 and is strongly significant. The relative weight of permanent income is now estimated to be 0.7, well below the unity coefficient implied by the (textbook) permanent income model. The effect of the real effective interest rate on borrowing is now negative, though insignificant. The long-run MPC out of net financial assets is estimated at 0.06 with a t-ratio of 7.7. The housing wealth to income ratio and the log of the house price to income ratio are insignificant, both with negative effects. One could conclude from these findings that housing wealth has, if anything, a small *negative* effect on consumption in Italy, as found by Boone and Girouard (2002) and Slacalek (2009). However, this model still fails the parameter stability test and the residual autocorrelation remains very high.

A dramatic improvement comes with a further disaggregation of the household balance sheet, shown in column 5. We now split net financial assets into three components: net liquid assets (defined as liquid assets - deposits plus mutual funds plus bills minus debt); semi-liquid assets (defined as

[16] In addition to the Durbin-Watson statistic, the tests include Lagrange-Multiplier (LM) tests for residual autocorrelation for both the moving average and for the autoregressive form for lags up to 4 quarters, see Godfrey (1988). The RESET2 test, or Ramsey Regression Equation Specification Error Test, is a test of potential misspecification in a linear regression model. It checks for issues related to the functional form of the model, such as omitted non-linear terms.

[17] As suggested by the review of central bank econometric policy models in Muellbauer (2022), low speeds of adjustment are a symptom of specification problems, especially of omitted variables (see also Hendry et al. (1984) and Hendry (1995, ch.7)).

bonds); and illiquid financial assets (defined as quoted shares plus pension wealth).[18] The equation standard error falls from 0.0043 to 0.0032, and the speed of adjustment jumps from 0.10 to 0.29. Moreover, parameter stability according to the Chow test is excellent and tests for residual autocorrelation are at last satisfactory. The relative weight of permanent to current income is estimated at 0.57; this is strongly against the extreme textbook view which bases all the weight on permanent income. The marginal propensities to consume are estimated with a relatively high degree of accuracy. The MPC of liquid assets minus debt is estimated at 0.14 (with a t-ratio of 14.8), the MPC of semi-liquid assets (bonds) is 0.054 (with a t-ratio of 6.7), and on the MPC of illiquid financial assets is 0.0.037 (with a t-ratio of 2.2). These results are broadly in line with other countries (discussed below) except for the relatively high MPC on illiquid financial assets. There is a strongly significant positive housing wealth effect but it is substantially offset by the negative affordability of higher house prices relative to income. Hence, high house prices have different consumption consequences for home-owners versus renters, and the net outcome for Italy is that the weight of the owners somewhat exceeds that of renters. Overall, this evidence strongly contradicts the conventional net worth restriction.[19] The real effective interest rate effect is now negative and strongly significant. However, the negative estimated effect of the Fornero reforms on the long-run consumption to income ratio is implausibly large at over 6 percent.

The GFC and the subsequent European sovereign debt crisis had severe domestic consequences for fiscal policy, likely to have raised income insecurity among Italian households. Probably households became less forward looking and more focused on current income A simple way of testing this within the model is to permit a shift in the relative weights of permanent and current income from the GFC onwards. We interact the 2008Q3 shift dummy with the log-ratio of permanent to current income (with this ratio defined relative to its value in 2008Q2). The interaction effect is significant at the 10 percent level and reduces the relative weight on permanent income from a point estimate of 0.57 pre-GFC, to 0.21 post-GFC, see column 6. The speed of adjustment is 0.30 and the estimated long-run effect of the Fornero reforms becomes a more plausible 0.035 reduction in the consumption to income ratio (compared with a 0.061 reduction, when omitting the post-GFC shift in the weight of permanent income in column 5).

[18] Note that unquoted shares are excluded from our preferred specifications, as the estimated implied MPC would have been negative in all specifications. This is related to mismeasurement of unquoted shares, see discussion in Section 3.2

[19] For comparison, the Bank of Italy's consumption equation in BIQM uses net financial wealth (excluding housing wealth) and includes a *negative* real interest rate effect. The speed of adjustment is only around 0.07 in contrast with ours of 0.30, and the MPC out of net financial wealth is round 0.08, not far from ours of 0.06, though the results are not strictly comparable, since the wealth coverage is for the whole private sector (including businesses) rather than just households, and the definition of consumption includes the imputed service flow of durable goods instead of expenditure on durables.

Table 1. Italian Consumption Function Estimates, 1980-2019

$\Delta \ln c_t$	Symbol	(1)	(2)	(3)	(4)	(5)	(6)
				Long-run effects			
Speed of adjustment	λ	0.0531 (0.0219) [2.43]	0.0632 (0.0216) [2.92]	0.0398 (0.0184) [2.16]	0.103 (0.0263) [3.92]	0.289 (0.0276) [10.46]	0.298 (0.0273) [10.91]
$\ln y_t - \ln c_{t-1}$	-	1.00	1.00	1.00	1.00	1.00	1.00
Constant	α_0	-0.293 (0.180) [-1.63]	-0.299 (0.146) [-2.04]	-0.00999 (0.0155) [-0.0643]	0.151 (0.416) [0.364]	0.369 (0.1254) [2.94]	0.298 (0.1225) [2.43]
1990 seasonal pattern dummy	α_{0S}	0.105 (0.0568) [1.85]	0.0854 (0.0414) [2.06]	0.139 (0.0816) [1.70]	0.060 (0.0223) [2.69]	0.024 (0.00477) [5.11]	0.023 (0.00451) [5.19]
Adjusted r_t	α_1	0.769 (0.569) [1.35]	0.667 (0.448) [1.49]	0.982 (0.765) [1.28]	-0.0308 (0.272) [-0.113]	-0.361 (0.0769) [-4.70]	-0.350 (0.0733) [-4.78]
$\ln (y_t^p / y_t)$	α_2	1.48 (0.473) [3.13]	1.28 (0.352) [3.63]	1.43 (0.596) [2.41]	0.704 (0.197) [3.56]	0.574 (0.0544) [10.55]	0.615 (0.0548) [11.22]
2008q3 GFC dummy x $\ln (y_t^p / y_t)$	$\alpha_{2,int}$						-0.406 (0.1583) [-2.57]
1992 Exchange rate crisis dummy	τ_{1992}	-0.0425 (0.0400) [-1.06]	-0.0322 (0.0325) [-0.994]	-0.0385 (0.0528) [-0.729]	-0.02399 (0.0200) [-1.19]	-0.0363 (0.00644) [-5.63]	-0.0345 (0.00612) [-5.64]
2012 Fornero reform dummy	τ_{2012}	-0.0525 (0.0290) [-1.81]	-0.0308 (0.0261) [-1.18]	-0.00545 (0.0462) [-0.118]	-0.0151 (0.0216) [-0.702]	-0.0614 (0.00794) [-7.73]	-0.0346 (0.0129) [-2.68]
$\ln A_{t-1}/y_t$	γ	0.172 (0.0819) [2.10]	0.173 (0.0668) [2.60]				
A_{t-1}/y_t	γ_{NETW}			0.00766 (0.0176) [0.436]			
NFA_{t-1}/y_t	$\gamma_{1,NFA}$				0.0594 (0.00768) [7.74]		
NLA_{t-1}/y_t	$\gamma_{1,NLA}$					0.137 (0.00926) [14.84]	0.137 (0.00878) [15.57]
$BONDS_{t-1}/y_t$	$\gamma_{1,SMLB}$					0.0540 (0.00808) [6.68]	0.0623 (0.00817) [7.63]
$ILFA_{t-1}/y_t$	$\gamma_{1,ILAPEN}$					0.0365 (0.0169) [2.15]	0.0301 (0.0165) [1.83]

HA_{t-1}/y_t	γ_2				-0.00716 (0.0142) [-0.505]	0.0411 (0.00559) [7.35]	0.0392 (0.00533) [7.36]
$\ln HP_{t-1}/y_{t-1}$	γ_3				-0.0305 (0.0760) [-0.402]	-0.109 (0.0230) [-4.74]	-0.0962 (0.0224) [-4.29]
			Short-run effects				
$\Delta_2 \ln CCI_{t-2}$	β_0		0.0379 (0.0129) [2.94]	0.0365 (0.0130) [2.82]	0.0114 (0.0153) [0.750]	0.0295 (0.0112) [2.62]	0.0315 (0.0111) [2.85]
$\Delta_2 \ln c_t^P$	β_1	-0.0230 (0.00875) [-2.63]	-0.0237 (0.00853) [-2.78]	-0.0229 (0.00859) [-2.67]	-0.0294 (0.00847) [-3.48]	-0.0173 (0.00656) [-2.63]	-0.0166 (0.00644) [-2.57]
$\Delta\Delta_4 \ln y_t$	β_2	0.0858 (0.0232) [3.69]	0.0879 (0.0227) [3.88]	0.0891 (0.0228) [3.91]	0.0848 (0.0221) [3.84]	0.0803 (0.0168) [4.77]	0.0769 (0.0166) [4.64]
$\Delta_2 \ln ur_t$	β_3	-0.477 (0.0993) [-4.80]	-0.412 (0.0992) [-4.15]	-0.367 (0.101) [-3.64]	-0.155 (0.113) [-1.37]	-0.297 (0.0855) [-3.48]	-0.304 (0.0839) [-3.62]
$\lvert\hat{\varepsilon}_{t-1}\rvert$	β_4	-0.0579 (0.0530) [-1.09]	-0.0437 (0.0518) [-0.843]	-0.0463 (0.0524) [-0.883]	-0.0551 (0.0505) [-1.09]	-0.169 (0.0406) [-4.15]	-0.159 (0.0400) [-3.97]
$d_{1983:q1}$	β_5	-0.0193 (0.00471) [-4.11]	-0.0190 (0.00459) [-4.13]	-0.0185 (0.00461) [-4.01]	-0.0182 (0.00447) [-4.08]	-0.0214 (0.00340) [-6.28]	-0.0216 (0.00334) [-6.48]
$d_{1993:q1}$	β_6	-0.01692 (0.00501) [-3.37]	-0.0166 (0.488) [-3.40]	-0.0166 (0.00492) [-3.37]	-0.0154 (0.00475) [-3.23]	-0.0148 (0.00360) [-4.13]	-0.0144 (0.00353) [-4.09]
$d_{2013:q1}$	β_7	-0.0137 (0.00479) [-2.87]	-0.0131 (0.00467) [-2.80]	-0.0133 (0.00470) [-2.83]	-.0134 (0.00454) [-2.96]	-0.0125 (0.00346) [-3.62]	-0.0116 (0.00341) [-3.41]
			Diagnostics				
Standard error x 100		0.454	0.442	0.445	0.429	0.325	0.319
Adjusted R^2		0.510	0.535	0.528	0.562	0.749	0.758
LM Het. test (p-value)		0.474	0.822	0.891	0.701	0.182	0.162
Durbin-Watson		1.53	1.53	1.56	1.60	2.11	2.19
AR1/MA1 (p-value)		0.002	0.002	0.004	0.012	0.509	0.236
AR4/MA4(p-value)		0.001	0.000	0.000	0.000	0.450	0.401
Chow test (p-value)		0.005	0.001	0.000	0.000	0.962	0.960
RESET2 test (p-value)		0.04	0.139	0.088	0.816	0.841	0.480
Schwarz Criterion		-606.1	-608.2	-607.1	-609.1	-649.6	-650.7
Log likelihood		644.2	648.8	647.7	654.7	700.3	704.0

Notes: standard errors in parenthesis; t-stats in square brackets. An ogive dummy based on 8 quarters starting from 2017:q1 is not significant. A 2-quarter moving average transformation is applied to the 1992 exchange rate crisis dummy. Acronyms are as follows: *A* is net worth including housing wealth; *NFA* is net financial assets, defined as total financial assets minus loans; *NLA* is net liquid assets defined as deposits plus mutual funds plus bills minus loans; *BONDS* is bonds; *ILFA* is illiquid assets defined as pensions plus quoted shares, *HA* is housing wealth, *HP* is the real house price index, *CCI* is the credit conditions index, c^p is public consumption , *ur* is the unemployment rate, $\hat{\varepsilon}_{t-1}$ is lagged income volatility.

Table 2. Italian Consumption Function Estimates on different sample periods.

$\Delta \ln c_t$	Symbol	1980-1999	1980-2007	1980-2010	1985-2019	1980-2019	$\hat{y}$
		(1)	(2)	(3)	(4)	(5)	(6)
				Long-run effects			
Speed of adjustment	λ	0.324 (0.0505) [6.41]	0.297 (0.0362) [8.22]	0.326 (0.0353) [9.23]	0.305 (0.0302) [10.10]	0.298 (0.0273) [10.91]	0.296 (0.0279) [10.59]
$\ln y_t - \ln c_{t-1}$	-	1.00	1.00	1.00	1.00	1.00	1.00
Constant	α_0	0.501 (0.3630) [1.38]	0.231 (0.1873) [1.23]	0.184 (0.1512) [1.22]	0.578 (0.2077) [2.78]	0.298 (0.1225) [2.43]	0.293 (0.1274) [2.30]
1990 seasonal pattern dummy	α_{0S}	0.020 (0.0056) [3.63]	0.023 (0.0051) [4.56]	0.021 (0.0046) [4.60]	0.025 (0.0060) [4.14]	0.023 (0.00451) [5.19]	0.024 (0.0046) [5.16]
Adjusted r_t	α_1	-0.219 (0.1074) [-2.04]	-0.294 (0.0881) [-3.33]	-0.292 (0.0793) [-3.68]	-0.371 (0.1011) [-3.67]	-0.350 (0.0733) [-4.78]	-0.344 (0.0790) [-4.35]
$\ln (y_t^p/y_t)$	α_2	0.601 (0.0820) [7.32]	0.634 (0.0612) [10.37]	0.633 (0.0563) [11.23]	0.564 (0.0616) [9.15]	0.615 (0.0548) [11.22]	0.623 (0.0747) [8.34]
2008: q3 GFC dummy x $\ln (y_t^p/y_t)$	$\alpha_{2,int}$			-0.422 (0.3954) [-1.07]	-0.294 (0.1674) [-1.76]	-0.406 (0.1583) [-2.57]	-0.384 (0.1927) [-1.99]
1992 Exchange rate crisis dummy	τ_{1992}	-0.042 (0.0110) [-3.77]	-0.036 (0.0076) [-4.76]	-0.032 (0.0064) [-4.95]	-0.031 (0.0061) [-5.11]	-0.035 (0.00612) [-5.64]	-0.035 (0.0063) [-5.51]
2012 Fornero reforms dummy	τ_{2012}				-0.0452 (0.0140) [-3.22]	-0.035 (0.0129) [-2.68]	-0.037 (0.0153) [-2.41]
NLA_{t-1}/y_t	$\gamma_{1,NLA}$	0.126 (0.0230) [5.47]	0.137 (0.0100) [13.65]	0.134 (0.0090) [14.88]	0.140 (0.0093) [15.05]	0.137 (0.00878) [15.57]	0.137 (0.0096) [14.32]
$BONDS_{t-1}/y_t$	$\gamma_{1,SMLB}$	0.065 (0.0192) [3.38]	0.066 (0.0119) [5.60]	0.069 (0.0107) [6.43]	0.066 (0.0082) [8.11]	0.062 (0.00817) [7.63]	0.061 (0.0089) [6.87]
$ILFA_{t-1}/y_t$	$\gamma_{1,ILAPEN}$	-0.012 (0.0466) [-0.26]	0.020 (0.0254) [0.81]	0.023 (0.0224) [1.04]	0.027 (0.0182) [1.46]	0.030 (0.0165) [1.83]	0.031 (0.0167) [1.85]
HA_{t-1}/y_t	γ_2	0.060 (0.0249) [2.43]	0.040 (0.0995) [4.02]	0.038 (0.0073) [5.11]	0.045 (0.0067) [6.74]	0.039 (0.00533) [7.36]	0.040 (0.0058) [6.77]
$\ln HP_{t-1}/y_{t-1}$	γ_3	-0.143 (0.0750) [-1.91]	-0.086 (0.0363) [-2.37]	-0.076 (0.0285) [-2.68]	-0.149 (0.0388) [-3.83]	-0.096 (0.0224) [-4.29]	-0.096 (0.0230) [-4.17]
				Short-run effects			

$\Delta_2 \ln CCI_{t-2}$	β_0	0.049 (0.0183) [2.68]	0.041 (0.0145) [2.85]	0.044 (0.0141) [3.08]	0.027 (0.0125) [2.11]	0.032 (0.0111) [2.85]	0.031 (0.0111) [2.83]
$\Delta_2 \ln c_t^P$	β_1	-0.020 (0.0099) [-2.04]	-0.016 (0.0073) [-2.20]	-0.014 (0.0073) [-1.89]	-0.021 (0.0089) [-2.34]	-0.017 (0.00644) [-2.57]	-0.017 (0.0065) [-2.55]
$\Delta\Delta_4 \ln y_t$	β_2	0.099 (0.0247) [3.98]	0.081 (0.0192) [4.22]	0.075 (0.0192) [3.92]	0.084 (0.0181) [4.67]	0.077 (0.0166) [4.64]	0.074 (0.0213) [3.49]
$\Delta_2 \ln ur_t$	β_3	-0.338 (0.1803) [-1.87]	-0.311 (0.1314) [-2.36]	-0.415 (0.1244) [-3.34]	-0.306 (0.0929) [-3.29]	-0.304 (0.0839) [-3.62]	-0.308 (0.0902) [-3.42]
$\lvert\hat{\varepsilon}_{t-1}\rvert$	β_4	-0.168 (0.0594) [-2.83]	-0.143 (0.0468) [-3.06]	-0.171 (0.0465) [-3.66]	-0.163 (0.0412) [-3.95]	-0.159 (0.0400) [-3.97]	-0.159 (0.0402) [-3.97]
$d_{1983:q1}$	β_5	-0.022 (0.0040) [-5.50]	-0.022 (0.0036) [-6.07]	-0.022 (0.0036) [-6.08]		-0.022 (0.00334) [-6.48]	-0.022 (0.0033) [-6.45]
$d_{1993:q1}$	β_6	-0.017 (0.0043) [-3.82]	-0.015 (0.0038) [-3.89]	-0.015 (0.0039) [-3.94]	-0.015 (0.0036) [-4.18]	-0.014 (0.00353) [-4.09]	-0.014 (0.0036) [-4.04]
$d_{2013:q1}$	β_7				-0.012 (0.0034) [-3.44]	-0.012 (0.00341) [-3.41]	-0.012 (0.0034) [-3.37]
Diagnostics							
Standard error x 100		0.376	0.340	0.346	0.315	0.319	0.319
Adjusted R^2		0.675	0.695	0.705	0.754	0.758	0.758
LM Het. test (p-value)		0.303	0.088	0.148	0.334	0.162	NA
Durbin-Watson		2.40	2.46	2.16	2.13	2.19	2.19
AR1/MA1 (p-value)		0.072	0.019	0.355	0.418	0.236	0.222
AR4/MA4(p-value)		0.241	0.035	0.316	0.262	0.401	0.353
Chow test (p-value)		0.932	0.988	0.992	0.980	0.960	0.925
RESET2 test (p-value)		0.948	0.496	0.871	0.372	0.480	NA
Schwarz Criterion		-305.8	-446.9	-492.4	-569.8	-650.7	-701.4
Log likelihood		345.2	489.4	538.2	619.3	704.0	704.0

Notes: standard errors in parenthesis; t-stats in square brackets. An ogive dummy based on 8 quarters starting from 2017: q1 is not significant. A 2-quarter moving average transformation is applied to the 1992 exchange rate crisis dummy. See Table 1 for acronyms.

The preferred model is estimated over various samples to test for parameter stability. In Table 2, columns 1 to 4 contrast the results for the periods 1980-1999, 1980-2007, 1980-2010, and 1985-2019, respectively, with that of the final model of Table 1, repeated in column 5 of Table 2, for 1980-2019. The Chow test confirms that the model is stable, and all the parameter estimates lie within one standard error of the full sample estimates in column 5 of Table 2. As current household income is potentially endogenous, as strongly argued by Hall (1978), it is helpful to compare results of instrumental variable estimates where current income is instrumented by its fitted value from an equation using lagged data. Column 6 of Table 2 shows these estimates. The results are very close to the OLS estimates, suggesting that income endogeneity is not an issue.

To help interpret the quantitative relevance of the long-run solution implied by the preferred specification shown in Table 1, column 6, we examine the fitted contributions to the log consumption to income ratio in Figures 9 and 10. Figure 9 shows the fitted contribution of the effective real interest rate on borrowing, the log ratio of permanent to current income, the 1992-3 step dummy and the Fornero reform dummy. Figure 10 shows the fitted contributions of the key portfolio ingredients. These are the ratios to income of net liquid assets, semi-liquid assets, illiquid financial assets and the combined positive effect of the ratio to income of housing wealth and the negative offset of housing affordability, measured by the log house price to income ratio.

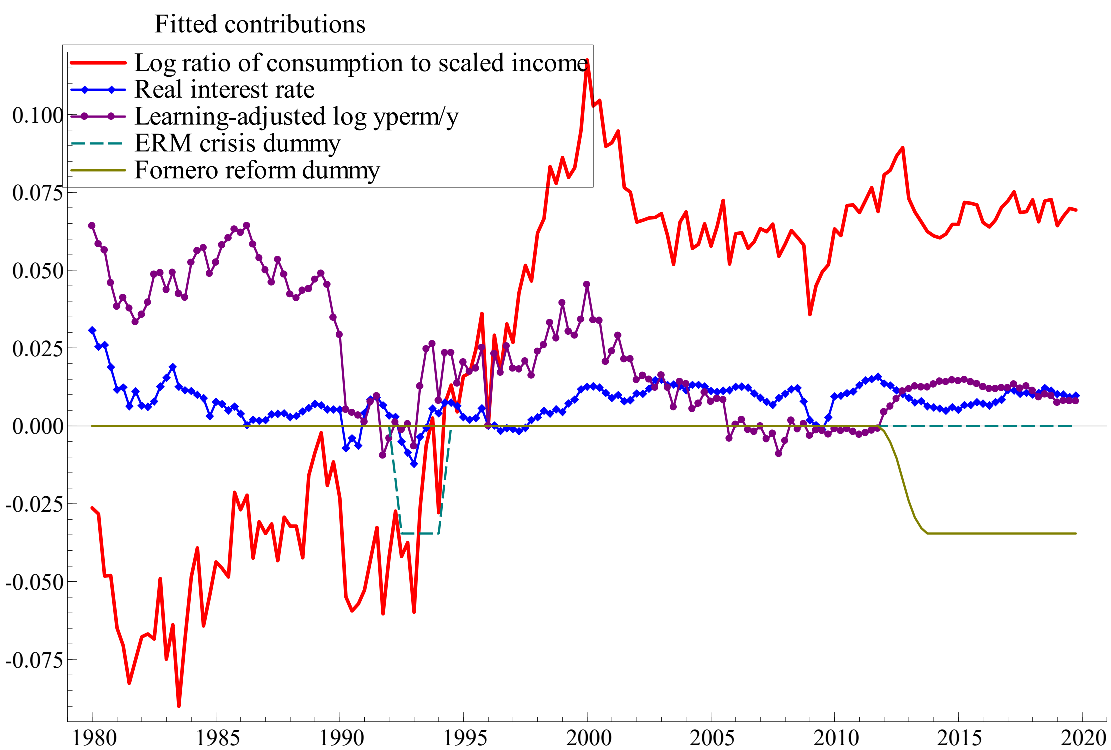

Figure 9: Fitted long-run contributions to log ratio of consumption to scaled income, part 1.
Notes: Fitted contributions are for the effective real interest rate, the deviation of log permanent income from current income, the 1992-3 step dummy and the Fornero reform dummy.

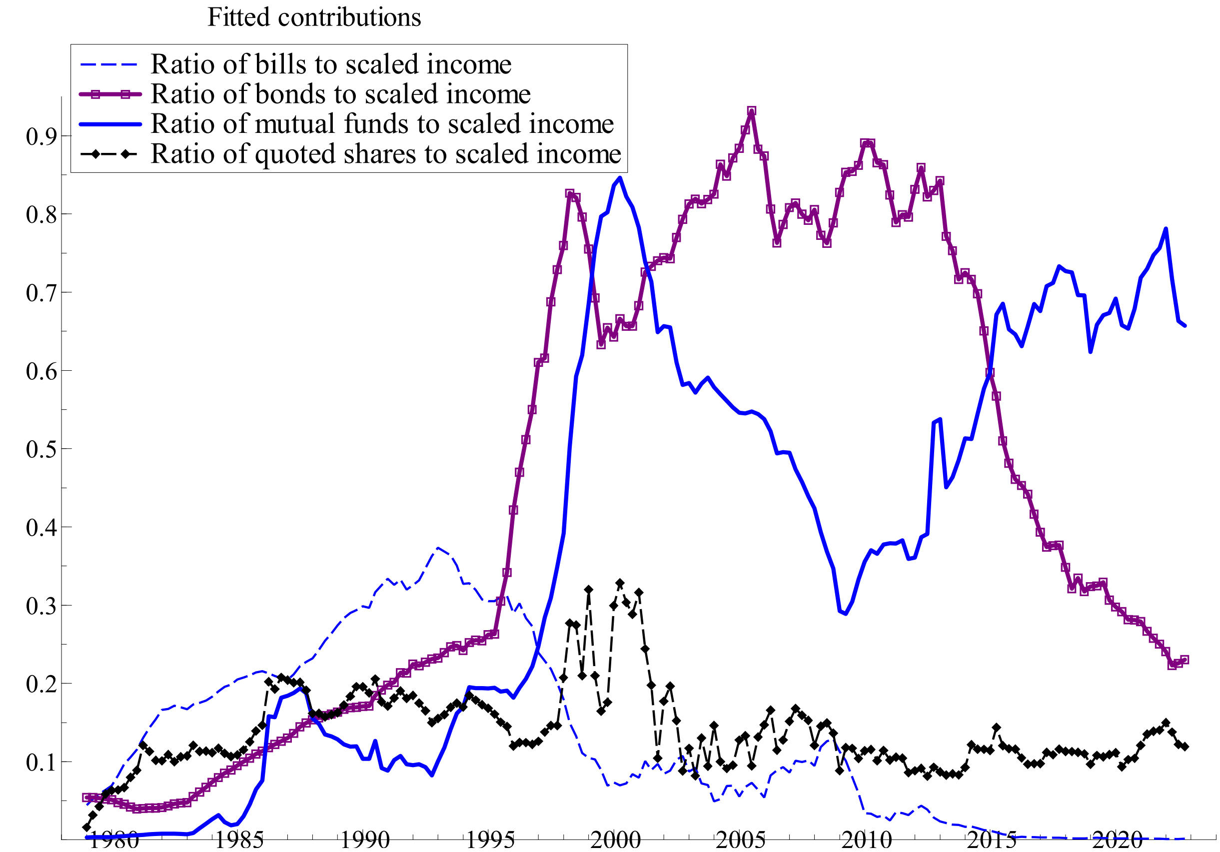

Figure 10: Fitted long-run contributions to log ratio of consumption to scaled income, part 2.
Notes: The fitted contributions are for the ratios to income of net liquid assets, semi-liquid assets (bonds), illiquid financial assets and the combined effect of the ratio to income of housing wealth and the log house price to income ratio.

The two graphics together suggest that the fall in the early 1990s in the consumption to income ratio was mainly driven by the fall in income relative to permanent income, together with the 1992-3 step dummy (see the discussion of the ERM crisis in Section 3.6). The rise in the ratio that followed was partly due to the ending of the ERM crisis, and partly due to the rising contributions of net liquid assets (especially by the mutual funds component), bonds, and illiquid financial assets in the boom period prior to joining the Euro. However, after the early 2000s, a faster rise of household debt than household deposits, and the decline in mutual funds, caused the ratio of net liquid assets to income to decline, offsetting some of the upwards drivers of other parts of the portfolio. Thus, following a peak in the consumption to income ratio in early 2001, the ratio stabilised at a lower level.

After 2011, one would have expected consumption to rise relative to income, given some recovery in the ratios of permanent income to income, and especially in net liquid assets to income. The fact that it failed to do so and instead fell moderately, our model attributes to the Fornero reforms of 2011, see Section 3.6.

Some of the fluctuations in the log of the ratio of consumption to income are sharper than can be accounted for by the long-run drivers pictured here. This is partly because of dynamic adjustment to shifts in these drivers, and partly because short-run factors, such as changes in the unemployment rate, the acceleration of income and income volatility, also play a role, most often at turning points in the consumption to income ratio.

4.2 A range of robustness checks

An initial round of robustness tests on the model are performed in Table 3. The first column examines the effect of omitting the change in the credit conditions index. With one exception, the results are very similar to those shown in Tables 1 and 2: however, the estimated coefficient on illiquid financial assets rises from 0.030 to 0.048. This is probably because the easing of credit constraints was associated with higher asset prices. Regarding other short-term controls, an interesting feature of the Table 1 results is that, even with the poorly specified formulations in columns 1 to 4, the short-term controls, including dummies for three outliers, have the same signs as in the favoured specification, and are mostly significant. This finding is evidence for the robustness of the short-term effects and the dummies selected. Reinforcing this, when all eight short term controls (including the dummies for the outliers) are omitted in column 2 of Table 3, the fit deteriorates, but the estimated speed of adjustment, and most of the long-run coefficients, are little changed. The one exception is the coefficient on illiquid financial assets, which rises to 0.066. One interpretation is that the stock market reacts to short-run shocks, changes in credit conditions and indicators of income insecurity, so that when their direct effects on consumption are omitted, there is an upward bias in the estimated wealth effect for illiquid assets.

In column 3, the net liquid assets restriction is relaxed, with separate coefficients on liquid assets and household debt. The results suggest that the restriction, of equal and opposite signed coefficients, is easily accepted. The final column shows results for the alternative grouping of semi-liquid and illiquid financial assets discussed in Section 2. Here liquid assets are restricted to deposits, semi-liquid assets are defined as bills plus mutual funds, and illiquid financial assets are bonds plus quoted shares plus pensions and insurance company reserves. The fit and parameter stability are slightly worse, though the speed of adjustment is much the same. The MPC out of net liquid assets is a little higher at 0.15, that on semi-liquid assets is 0.125 and on illiquid financial assets is 0.059. The estimated increase in the saving rate due to the Fornero reforms is higher at 0.049. The housing effects are quite similar to those for the preferred split of semi-liquid and illiquid financial assets. The post-GFC shift in the relative weights of permanent and current income is confirmed. Fundamentally, the key findings of the paper regarding the importance of separate estimation of the MPC for different groupings of assets, and the need to control for housing affordability when estimating housing wealth effects, are upheld.

A different robustness check is to compare the results for the conventionally defined real interest rate on borrowing with those for the effective real interest rate as defined by the BIS. The estimates and the goodness of fit are very similar, but parameter stability is slightly superior for the effective real rate. While adjustable-rate debt dominates in Italy, as historically has been the case in the UK, Sweden and Finland, the absence of a negative (aggregate) cash-flow effect in Italy by comparison with say the UK, is noteworthy. While in the UK, the change in the nominal interest rate has a significant negative effect on consumer spending (Aron and Muellbauer, 2012 and Cloyne et al. 2016, as originally posited by Jackman and Sutton, 1982), this is not so in Italy. This is almost certainly because household debt is low compared to household liquid assets, in contrast to the higher debt countries mentioned above. When interest rates rise, the cash-flow of borrowers declines while that of liquid asset holders improves. Borrowers tend to respond more strongly to cash-flow than owners of liquid assets (depositors). However, in Italy (unlike in the UK), the far greater size of liquid deposits in the balance sheets by comparison to household debt means that the net cash flow effect is not negative.

Table 3. Italian Consumption Function Estimates, 1980-2019: robustness checks.

$\Delta \ln c_t$	Symbol	Excluding change CCI	No short-run controls	Liquid assets vs loans	Different asset grouping
		(1)	(2)	(3)	(4)
		Long-run effects			
Speed of adjustment	λ	0.307 (0.0278) [11.04]	0.294 (0.0365) [8.08]	0.312 (0.0307) [10.14]	0.289 (0.0292) [9.88]
$\ln y_t - \ln c_{t-1}$	-	1.00	1.00	1.00	1.00
Constant	α_0	0.350 (0.1205) [2.90]	0.373 (0.1638) [2.28]	0.148 (0.1921) [0.77]	0.235 (0.119) [1.97]
1990 seasonal pattern dummy	α_{0S}	0.023 (0.0045) [5.16]	0.029 (0.0065) [4.45]	0.023 (0.0044) [5.21]	0.024 (0.00486) [4.98]
Adjusted r_t	α_1	-0.360 (0.0728) [-4.94]	-0.371 (0.0991) [-3.75]	-0.331 (0.0729) [-4.54]	-0.313 (0.0713) [-4.40]
$\ln (y_t^p / y_t)$	α_2	0.606 (0.0543) [11.17]	0.530 (0.0724) [7.32]	0.660 (0.0703) [9.38]	0.643 (0.0574) [11.2]
2008: q3 GFC dummy x $\ln (y_t^p / y_t)$	$\alpha_{2,int}$	-0.363 (0.1564) [-2.32]	-0.542 (0.2173) [-2.49]	-0.449 (0.1581) [-2.84]	-0.302 (0.1585) [-1.92]
1992 Exchange rate crisis dummy	τ_{1992}	-0.033 (0.0060) [-5.44]	-0.035 (0.0082) [-4.30]	-0.032 (0.0062) [-5.17]	-0.035 (0.00637) [-5.47]
2012 Fornero reforms dummy	τ_{2012}	-0.042 (0.0125) [-3.39]	-0.019 (0.0170) [-1.12]	-0.034 (0.0124) [-2.71]	-0.049 (0.0111) [-4.36]
NLA_{t-1}/y_t	$\gamma_{1,NLA}$	0.132 (0.0084) [15.72]	0.123 (0.0105) [11.79]		0.147 (0.0164) [9.00]
LA_{t-1}/y_t	$\gamma_{1,DEP}$			0.137 (0.0084) [16.25]	
$LOANS_{t-1}/y_t$	$\gamma_{1,LOANS}$			-0.092 (0.0463) [-1.98]	
$BONDS_{t-1}/y_t$	$\gamma_{1,SMLB}$	0.064 (0.0081) [7.85]	0.079 (0.0107) [7.38]	0.060 (0.0081) [7.49]	

$ILFA_{t-1}/y_t$	$\gamma_{1,ILAPEN}$	0.048 (0.0148) [3.23]	0.066 (0.0184) [3.57]	0.024 (0.0172) [1.38]	0.059 (0.00846) [6.97]
$SLFA_{t-1}/y_t$	$\gamma_{1,SEMILA}$				0.125 (0.0131) [9.56]
HA_{t-1}/y_t	γ_2	0.036 (0.0051) [6.99]	0.025 (0.0062) [4.03]	0.030 (0.0110) [2.68]	0.035 (0.00509) [6.95]
$\ln HP_{t-1}/y_{t-1}$	γ_3	-0.103 (0.0222) [-4.64]	-0.100 (0.0298) [-3.37]	-0.066 (0.0374) [-1.77]	-0.086 (0.0220) [-3.90]
			Short-run effects		
$\Delta_2 \ln CCI_{t-2}$	β_0			0.029 (0.0115) [2.48]	0.0296 (0.0112) [2.65021]
$\Delta_2 \ln c_t^P$	β_1	-0.018 (0.0066) [-2.81]		-0.016 (0.0065) [-2.45]	-0.0172 (0.00653) [-2.64]
$\Delta\Delta_4 \ln y_t$	β_2	0.076 (0.0170) [4.47]		0.074 (0.0169) [4.38]	0.0794 (0.0167 [4.76]
$\Delta_2 \ln ur_t$	β_3	-0.290 (0.0858) [-3.37]		-0.306 (0.0840) [-3.65]	-0.284 (0.0830) [-3.42]
$\lvert\hat{\varepsilon}_{t-1}\rvert$	β_4	-0.165 (0.0410) [-4.03]		-0.158 (0.0401) [-3.94]	-0.163 (0.0403) [-4.05]
$d_{1983:q1}$	β_5	-0.021 (0.0034) [-6.25]		-0.022 (0.0033) [-6.46]	-0.022 (0.00336) [-6.43]
$d_{1993:q1}$	β_6	-0.014 (0.0036) [-3.99]		-0.014 (0.0035) [-4.05]	-0.015 (0.00356) [-4.11]
$d_{2013:q1}$	β_7	-0.012 (0.0035) [-3.36]		-0.012 (0.0034) [-3.47]	-0.011 (0.00343) [-3.24]
			Diagnostics		
Standard error x 100		0.327	0.435	0.319	0.321
Adjusted R^2		0.746	0.549	0.758	0.755
LM Het. test (p-value)		0.097	0.136	0.182	0.152
Durbin-Watson		2.164	1.973	2.190	2.194
AR1/MA1 (p-value)		0.290	0.925	0.231	0.214
AR4/MA4(p-value)		0.419	0.158	0.397	0.469
Chow test (p-value)		0.998	0.725	0.828	0.956
RESET2 test (p-value)		0.863	0.104	0.431	0.604

Schwarz Criterion	-648.7	-616.7	-648.7	-649.8
Log likelihood	699.5	649.7	704.5	703.1

Notes: standard errors in parenthesis; t-stats in square brackets. An ogive dummy based on 8 quarters starting from 2017:q1 is not significant. Moving average transformations apply to the 1992 exchange rate crises dummy (MA2). See notes to table 1 for portfolio acronyms for columns 1 to 3. In column 4, net liquid assets *NlA* is defined as deposits minus loans; semi-liquid assets *SLFA* is defined as bills plus mutual funds; illiquid financial assets *ILFA* is defined as bonds plus quoted shares plus pension wealth.

A further robustness test is to estimate the consumption function and the permanent income equation as a two-equation system. Then, instead of using the fitted value of the log permanent income, the log permanent income function is substituted into the jointly estimated consumption equation. The latter fits slightly less well, as the consumption parameters are influenced by the need to satisfy the cross-equation restrictions, but the parameter estimates and even the estimated standard errors are only a whisker away. It is encouraging that joint estimation gives similar estimates to the single equation estimates for the coefficients of the permanent income equation. It suggests that the selected model is capturing reasonably well the average income outlook of households.

A final robustness issue concerns the impact of the income growth projections made in 2019Q4 on the estimates. As consumer confidence data in 2019 show some deterioration relative to 2018, one might wonder whether the *Oxfordeconomics.com income* projections are over-optimistic relative to the actual expectations of Italian households. As a check, the growth rate of the projections was halved, otherwise retaining the cyclical pattern. This has remarkably little impact on the estimates, including of the permanent income equation. The fit of the consumption equation actually improves very slightly, though the change in the log likelihood is only 0.2, and the estimated coefficients are almost the same.[20] Previously reported estimates for the reduction in the consumption to income ratio following the Fornero reforms are reduced slightly, e.g. from 0.035 to 0.033 for our preferred model.

Though the outperformance in fit and parameter stability of the favoured specification (shown in Table 1 (column 6) and repeated in Table 2 (column 5) is only small, there are two, more important, reasons for preferring it to that shown in Table 3, column 4. One is that the estimate of the effect of the Fornero pension reforms of 0.035 is more plausible and consistent with micro-evidence than the estimate of 0.049 in the Table 3, column 4 alternative. Even more important is the fact that the estimate of the MPC out of illiquid financial assets of 0.03 (t=1.8) in the favoured specification is much more in line with estimates for other countries where illiquid financial assets also include directly-held stock market wealth and pension and long-term life insurance assets. The Table 3, column 4 alternative includes bonds in the illiquid asset definition, a far larger item in Italian household portfolios than is the case in other countries, with an MPC of 0.059 (t=7.0). In contrast,

[20] While we could have chosen results tables for this variant, it seemed preferable simply to take the more standard projection as the basis for end-of period permanent income estimates. In regular updating of the model estimates, one would typically be using the most recently available income growth projections, whether from the Oxford source or the 5-year ahead forecasts from the IMF. Alternatively, if this consumption equation were incorporated in a revised version of the Bank of Italy's BIQM, the income forecasts from the full model would produce an approximation of 'model consistent expectations'.

the favoured specification treats bonds as a semi-liquid asset, intermediate between liquid and illiquid assets, with an MPC of 0.062 (t=7.6). As can be deduced from Figures 4 and 5, when bonds are included in illiquid assets, the bond time-profile tends to dominate variations in the time-profile of illiquid assets, and this pushes up the estimated MPC in column 4 on illiquid assets (which includes bonds).

4.3 Comparisons with other countries: MPCs and credit conditions

In principle, it would be highly desirable to extend this study to many countries beyond Italy. However, there is an important reason why this is beyond our current research capacity. The exposition in Section 2 of the credit-augmented generalized consumption function emphasized the role of credit conditions. Throughout our study for Italy, tests for the relevance of generalized credit conditions, covering private sector firms as well as households, as part of the long-run solution showed an insignificant effect, in contrast to the important role played by variations in household credit conditions in France (Chauvin and Muellbauer, 2018), the UK (Aron et al. 2012), the US (Duca and Muellbauer, 2013), South Africa (Aron and Muellbauer, 2025a) and even Germany (Geiger et al. 2016), where credit liberalization went less far. It is noteworthy that household debt relative to income has grown only slowly in Italy and Italy still has one of the lowest ratios of household debt to income among OECD countries. This suggests little change in credit conditions relevant in the long-run for households in Italy, by comparison with many other economies. While the level of credit conditions is not relevant in the Italian consumption function, the *rate of change* of credit conditions does play a role in our models, helping to forecast income growth in the permanent income equation and directly significant in the consumption equation. However, parameter estimates for the long-run coefficients are robust, as we have seen, to the exclusion of the rate of change of credit conditions.

For most other countries the exclusion of controls for changing credit conditions would seriously contaminate any attempt to estimate the consumption function parameters. One illustration of the importance of credit conditions is that in three countries where home equity withdrawal is easily available - the UK, the US and South Africa - housing wealth appears *only* in interaction with mortgage credit conditions, and with high levels of significance. This suggests that in those countries housing wealth has more of a collateral effect than a conventional wealth effect. However, credit conditions are hard to observe directly, which is why the above studies use a latent interactive equation system (LIVES) to deduce shifts in credit conditions for mortgages and non-mortgage debt. Typically, this involves estimating an equation system also including mortgage and non-mortgage debt, house prices, and liquid assets drawing on country by country institutional information to guide empirical identification.[21] While extending such research beyond the countries mentioned above would be desirable, the effort required would be substantial, not least because assembling the necessary quarterly balance sheet estimates for many countries would encounter similar challenges to those surmounted in this paper for Italy.

[21] For the UK, the estimates of mortgage credit conditions were taken from a separate latent variable equation system, using distributional data on high loan to value and loan to income mortgages, as well as data on the aggregate mortgage stock, see Fernandez-Corugedo and Muellbauer (2006).

For the five countries, where broadly comparable estimates are available, Table 4 summarises estimates of the parameters for net liquid assets and for illiquid financial assets, where the latter include the assets of pension and long-term life insurance funds.

Table 4: Estimates of MPCs for net liquid and illiquid financial assets for 6 countries

MPC out of assets		Net liquid assets$_{t-1}$ / y $_t$		Illiquid financial assets$_{t-1}$ / y$_t$	
		Coefficient, γ_1	*t-ratio*	*Coefficient,* γ_2	*t-ratio*
Italy	**1980:1 to 2019:4**	**0.14**	**15.6**	**0.03**	**1.8**
France	1981:2 to 2016:4	0.14	4.4	0.022	3.3
South Africa	1992:1 to 2020:1	0.14	4	0.018	3.1
Germany	1981:3 to 2012:4	0.09	4.1	0.016	2.5
UK	1967:1 to 2005:4	0.11	8	0.022	8
US	1971:4 to 2011:1	0.1	7.6	0.017	8.6

Sources: For Italy: this paper, Table 1, column 6. For France: Chauvin and Muellbauer (2018). For South Africa: Aron and Muellbauer (2025a). For the UK: Aron et al. (2012). For Germany: Geiger et al. (2016). For the US: Duca and Muellbauer (2013).

Although the institutions and the composition of household balance sheets differ amongst the six countries, the parameter estimates are broadly similar. The Italian estimates for γ_1 of 0.14 (t=15.6) and for γ_2 of 0.03 (t=1.8) fit well into the range of estimates for the other five countries, especially of γ_2, the MPC for illiquid financial assets, of 0.017 to 0.022. This is an important reason for preferring the chosen grouping of Italian assets into the different asset categories, compared with the alternative asset grouping in Table 3, column 4, though the latter fits almost as well.

5. Conclusions

In recent years, central banks have evolved away from New Keynesian DSGE models, or early vintages of these models, through the redesign of the main policy model, or in the weights given to different models within a suite of models. Previous vintages of New Keynesian DSGE models were founded on the optimizing behaviour of representative agents facing linear intertemporal budget constraints, and typically omitted household balance sheets, asset prices and credit flows, and imposed strong restrictions on the data being modelled. Most central banks now include semi-structural econometric models in their policy modelling platforms, allowing the data to 'speak' under less strict prior restrictions.[22]

Central to the semi-structural econometric policy models is the formulation of the aggregate consumption function, with consumption being by far the largest component of GDP. Yet many of these formulations are still based on the text-book model of a representative consumer following the life-cycle permanent income hypothesis. In the text-book model, complex household portfolios are summarized in a single net worth measure, treating liquid and illiquid financial assets and housing as if the same marginal propensity to consume applied to each component. Credit constraints or variations in loan standards, and uncertainty, are largely disregarded. That the extreme restrictions of the text-book model should be challenged and tested is also supported by an evolution in the micro foundations of macroeconomics. The modern view of heterogeneous agent behaviour under uncertainty in incomplete markets has monetary policy transmission implications for the household sector that are radically different from the text-book model. For example, rather than operating primarily through intertemporal substitution, monetary transmission via households is now seen to operate more through general equilibrium effects on income and through the implications of changes in interest rates for asset prices and for balance sheets (Pascal, 2024).

This paper has challenged the restrictions embodied in the conventional formulation for an aggregate consumption function for Italy, using quarterly data from 1980 to 2019. Our model disaggregates net worth into sub-components with different liquidity characteristics and housing wealth, it controls for income expectations through modelling a permanent income construct with adjustment for structural breaks through learning, and it controls for a real interest effect. In the short run, the model controls for income uncertainty and the effects of changing credit conditions.

The fully restricted standard net worth formulation is strongly rejected. Progressive relaxation of the restrictions reveals striking improvements in statistical fit and model coherence. It is difficult even to find a significant net worth effect. An alternative formulation where net worth is replaced by net financial wealth, as in the Bank of Italy's semi-structural model BIQM (Bank of Italy Quarterly Model), does a little less badly, but it is still strongly rejected against our preferred specification where net worth is disaggregated into liquid, semi-liquid and illiquid assets and debt, controlling separately for housing affordability and housing wealth.

With heterogeneity such an important issue, and with these consumption models estimated on aggregate data, an element of caution needs to be exercised in their policy application. The most obvious example is if there were to be large shifts in the distribution of assets over different households (e.g. as might have occurred in the pandemic). In this case, the parameters estimated on

[22] These central banks include the US Federal Reserve, the ECB, the Bank of Canada, the Reserve Bank of Australia, and many central banks in Europe, including those of Germany, France, Italy, Spain, the Netherlands and Ireland. These banks are currently being joined by the Bank of England, and the Czech National Bank, where such models are under development (see Frait (2025) and Aron and Muellbauer (2025b) for the background to the CNB's decision).

macro-data may be subject to instability or may shift. In the pandemic, by the end of 2020, there was a surge 'excess savings', see Colabella et al. (2023). Some of this saving would have been seen in increases in total holdings of deposits and other liquid assets. However, data from the Surveys of Household Income and Wealth suggest little change from end-2016 to end-2020 in the distributional pattern by income quintile (see our Appendix A.3). There is, therefore, little *prima facie* evidence for notable changes in Italy's MPCs for aggregate holdings of deposits and other liquid assets, pre versus. post-pandemic.

Expectations of future income play a crucial role in our empirical model. The omission of the permanent income variable in the regression results in a halving of the estimated speed of adjustment, an implausible negative illiquid asset wealth effect, a substantially worse equation fit, parameter instability, and serially correlated residuals. With the higher discount rate applied to future income flows motivated in the paper, our results suggest weights of 43 percent on current household income and 57 percent on permanent income, before the European sovereign debt crisis. This supports the view that current income is important for consumption, and contrasts with the simple life cycle model which puts all weight on permanent income. Since the debt crisis, we find evidence for the weight shifting even further to current income, as households became more uncertain about long-run income prospects. In contrast to the text-book view, our findings imply a substantial fiscal policy multiplier, important for evaluating the effectiveness of fiscal policy

Our results for Italy find that the marginal propensity to consume out of liquid assets (comprising deposits, Treasury bills and bank-managed funds) is around 0.14, the MPC for bonds, more widely-held by households in Italy than in most other countries, is around 0.06, and the MPC for illiquid financial assets (consisting of directly-held quoted shares and pension and long-term life-insurance assets) is around 0.03. To contrast these findings with the Bank of Italy's BIQM, that model reports only the long-run MPC for net financial worth (excluding housing), which is around 0.08. This exaggerates the MPC for quoted shares and pension and life-insurance assets but understates that for liquid assets. In our model, a significant positive effect from housing wealth is partly but not completely offset by the negative effect of housing affordability (measured by the house price to income ratio). Thus, in contrast to the BIQM, which does not include housing, our model has a *net* positive effect from higher house prices on consumption.

A full comparison of model properties for monetary transmission needs the incorporation of our consumption model in a complete econometric model. Nevertheless, several differences stand out as between our model and the BIQM which consider a consumption function for the entire non-financial private sector and is mainly used to produce projections and forecasts of the main macro variables. First, concerning a long-run effect of cuts in the short-term policy rate, our consumption equation shows a significant direct short-term interest rate effect, while BIQM uses the real yield on long-term Italian sovereign debt, which historically (and in BIQM) has been quite detached from the short-term policy rate. Second, our model implies pass-through channels from a cut in the short-term policy rate to the mutual funds and short-term bills component of liquid assets, and to net housing wealth. Third, our model also implies a channel to permanent income through equity prices. BIQM does not include a role for permanent income and therefore omits this channel. Fourth, in the other direction, our model suggests that the rise in debt caused by a lower policy rate, will, in due course, have negative consumption effects. Both models share a long-run general equilibrium transmission through current income. Finally, the speed of adjustment of our consumption model at around 0.3 is higher than that of the BIQM consumption equation, resulting in a faster and more complete adjustment to shocks and transmission of monetary impulses.

Our higher speed of transmission is consistent with recent evidence from Buda et al (2025). Using daily data on consumption and investment from bank transactions, they find large, short-run effects of monetary policy, by contrast with conventional wisdom that such lags are 'long and variable'. It is not surprising that mis-specified models find lags which are both long and variable. They will be long because misspecification tends to show up in downward biases in the adjustment speed. They will be variable because, for example, as the portfolio composition alters, the weights on the different transmission channels shift in a way that the net worth or financial net worth restriction cannot capture.

While some of our findings are specific to Italy, the comparisons with estimated MPCs for assets in other countries, in Section 4, suggests some reassuring similarities, with broadly similar MPCs for different categories of financial assets. Another important similarity, shared with models for Germany, France and South Africa, is that interest rate cuts inducing higher house prices have the opposite effect on the consumption of owners (housing is more valuable) from that of renters (who need to save a bigger deposit and caution with rents more likely to rise). This suggests a serious distributional effect from interest rate changes of which policy-makers should be aware.

However, Italy's credit market architecture is different from that of most OECD countries: for instance, mortgage equity withdrawal does not exist and this contributed to far slower growth in household indebtedness. Any attempt to estimate a similar specification of the consumption function for other countries must include controls for shifts in credit conditions, especially for countries where home equity withdrawal is significant. The full implications for monetary transmission via households need to be explored in a comprehensive model that, *inter alia*, endogenises interest rates on borrowing, household mortgage and non-mortgage debt, liquid assets, house prices and housing investment with separate models for these variables, incorporating controls for credit conditions (see Muellbauer (2022) for the six main housing-related channels). This paper's focus on a well-specified consumption function goes a long way toward improving the understanding of monetary policy transmission via more clearly articulated channels.

References

Albareto, G., A. Cardillo, A. Hamaui, and G. Marinelli. 2020. "Mutual funds' performance: the role of distribution networks and bank affiliation," Bank of Italy Economic working papers, n. 1272.

Ando, A., L. Guiso and I. Visco. 1994. "Saving and the Accumulation of Wealth: Essays on Italian Household and Government Saving Behavior", MIT Press.

Aron, J., J. V. Duca, J. Muellbauer, A. Murphy, and K. Murata. 2012. "Credit, Housing Collateral and Consumption, Evidence from the U.K., Japan and the U.S." *Review of Income and Wealth* 58 (3):397–423.

Aron, J. and J. Muellbauer. 2025a "New models for South African consumption, house prices, and mortgage and non-mortgage debt: Insights for financial stability and monetary transmission", WIDER Working Paper, UNU-WIDER.

Aron, J. and J. Muellbauer. 2025b "Summary of the external review of the CNB's modelling framework" in Proceedings of the Czech National Bank Workshop on Monetary Policy: Inflation Targeting Frameworks Under Review, ed. E. Zamrazilová and J. Frait, p.63-72.

Bank of Italy. 2018. "Annual Report for 2017." May 2018.

Bassanetti, A. and F. Zollino. 2010. *Housing Markets in Europe*, chap. "The Effects of Housing and Financial Wealth on Personal Consumption: Aggregate Evidence for Italian Households". Springer.

Battistini, N., V. Di Nino and J. Gareis (2023) "The consumption impulse from pandemic savings – does the composition matter?", *ECB Economic Bulletin*, Issue 4/2023.

Berger, D., V. Guerrieri, G. Lorenzoni and J. Vavra, (2018), "House Prices and Consumer Spending", *Review of Economic Studies*, 85 (3): 1502-42.

Blinder, A. S. and A. Deaton. 1985. "The Time Series Consumption Function Revisited." *Brookings Papers on Economic Activity* 16 (2):465–521.

Bonci, R. and M. Coletta. 2008. "Italy's financial accounts since 1950", in *Financial Accounts: History, Methods, the Case of Italy and International Comparisons*, Proceedings of the papers presented at the Bank of Italy conference held in Perugia, 1-2 December 2005.

Boone, L. and N. Girouard. 2002. "The stock market, the housing market and consumer behaviour." *OECD Economic Studies 35*, OECD.

Bruno, G. 2008. "Obtaining quarterly series from the annual financial accounts", in *Financial Accounts: History, Methods, the Case of Italy and International Comparisons*, Proceedings of the papers presented at the Bank of Italy conference held in Perugia, 1-2 December 2005.

Buda, G. , V. M. Carvalho, G. Corsetti, J.B. Duarte, S. Hansen, A.S. Moura, A. Ortiz, T. Rodrigo, J.V. Rodriguez Mora, A. Ortiz, and G.A. da Silva. 2025. "The Short Lags of Monetary Policy," Working Papers w202501, Banco de Portugal, Economics and Research Department.

Buiter, W. H. 2010. "Housing Wealth Isn't Wealth." *Economics - The Open-Access, Open-Assessment E-Journal* 4:1–29.

Bulligan, G., F. Busetti, M. Caivano, P. Cova, D. Fantino, A. Locarno, and L. Rodano. 2017. "The Bank of Italy econometric model: an update of the main equations and model elasticities." Temi di discussione (Economic working papers) 1130, Bank of Italy, Economic Research and International Relations Area. URL https://ideas.repec.org/p/bdi/wptemi/td_1130_17.html.

Calomiris, C., S. D. Longhofer and W. Miles. 2009. "The (Mythical?) Housing Wealth Effect," NBER Working Paper 15075.

Campbell, J. Y. 1987. “Does Saving Anticipate Declining Labor Income? An Alternative Test of the Permanent Income Hypothesis.” *Econometrica* 55 (6):1249–1273.

Cannari, L., G. D’Alessio, and G. Vecchi. 2017. “Wealth”, in G. Vecchi (edited by), *Measuring Wellbeing. A History of Italian Living Standards*. Oxford University Press.

Caprara, D., R. De Bonis, and L. Infante. 2020. “What do we know on household wealth?” *Rivista di Storia Economica*, April, 43-85.

Carroll, C. D. 1992. “The Buffer-Stock Theory of Saving: Some Macroeconomic Evidence.” *Brookings Papers on Economic Activity* 23 (2):61–156. URL https://ideas. repec.org/a/bin/bpeajo/v23y1992i1992-2p61-156.html.

Case, K. E., J. M. Quigley, and R. J. Shiller. 2005. “Comparing Wealth Effects: The Stock Market versus the Housing Market.” *Advances in Macroeconomics* 5 (1):1–32.

Catte, P., N. Girouard, R. Price and C. André. 2004. “Housing markets, wealth, and the business cycle.” OECD Economics Department Working Paper 394, OECD.

Chauvin, V. and J. Muellbauer. 2018. “Consumption, household portfolios and the housing market in France.”, *Economics and Statistics*, vol. 58. 157–178.

Cloyne, J., C. Ferreira and P. Surico. 2016. “Monetary Policy when Households have Debt: New Evidence on the Transmission Mechanism”, Bank of England Staff Working Paper No. 589.

Colabella, A., E. Guglielminetti and C. Rondinelli. 2023. “The distribution and use of Italian households' savings after the pandemic” Questioni di Economia e Finanza (Occasional Papers) 797, Bank of Italy, Economic Research and International Relations Area..

Cooper, D. and K. Dynan. 2016. “Wealth Effects and Macroeconomic Dynamics.” *Journal of Economic Surveys* 30 (1):34–55.

Crawley, E., and A. Kuchler. 2023. “Consumption Heterogeneity: Micro Drivers and Macro Implications”, American Economic Journal: Macroeconomics, 15 (1): 314-41.

De Bonis, R. and M. Marinucci. 2023. “A Short Note on Interest Rates and Household Wealth”. *Italian Economic Journal*, July, 617-635.

De Bonis, R. and A. Silvestrini. 2012. “The effects of financial and real wealth on consumption: new evidence from OECD countries.” *Applied Financial Economics,* 22 (5):409–425.

Deaton, A. 1991. “Saving and Liquidity Constraints.” *Econometrica* 59 (5):1221–1248.

Duca, J. V. and J. Muellbauer. 2014. *A Flow-of-Funds Perspective on the Financial Crisis*, chap. “Tobin LIVES: Integrating Evolving Credit Market Architecture into Flow-of-Funds Based Macro-Models”. Palgrave Studies in Economics and Banking. Palgrave Macmillan, London.

Fagereng, A., M. B. Holm and G. J. Natvik. 2021. "MPC Heterogeneity and Household Balance Sheets," *American Economic Journal: Macroeconomics*, 13(4), 1-54.

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

Fornero, E. 2020. “Italy’s pension reform: the twists and turns of an obstacle course”, Revue de l'OFCE, 120 (6) 59-84.

Frait, J. 2025. “Monetary policy in the era of “Great Macrofinancial Volatility” in Proceedings of the Czech National Bank Workshop on Monetary Policy: Inflation Targeting Frameworks Under Review, ed. E. Zamrazilová and J. Frait, p.53-62.

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

———. 1963. *Measurement in Economics*, chap. Windfalls, the ‘Horizon’ and Related Concepts in the Permanent Income Hypothesis. Stanford University Press, Palo Alto, CA.

Garrida, C. and A. Hedlund. 2020. “Mortgage Debt, Consumption, and Illiquid Housing Markets in the Great Recession.” *American Economic Review* 110 (6):1603–1634.

Geiger, F., J. Muellbauer and M. Rupprecht. 2016. "The housing market, household portfolios and the German consumer," Working Paper Series 1904, European Central Bank.

Gertler, M., J. Gali, and R. Clarida. 1999. “The Science of Monetary Policy: A New Keynesian Perspective.” *Journal of Economic Literature* 37 (4):1661–1707.

Godfrey, L. G. (1978). "Testing Against General Autoregressive and Moving Average Error Models when the Regressors Include Lagged Dependent Variables". *Econometrica.* 46 (6): 1293–1301.

Guiso, L., M. Paiella, and I. Visco. 2006. “Do Capital Gains Affect Consumption? Estimates of Wealth Effects from Italian Households’ Behavior.” In *Long-run Growth and Short-run Stabilization*, Chapters, chap. 3. Edward Elgar Publishing.

Guglielminetti, E., and C. Rondinelli. 2024. “Macro and Micro Consumption Drivers in the Wake of the COVID-19 Pandemic.” *International Journal of Central Banking*, 20(1), pages 181-237

Hall, R. E. 1978. "Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence," *Journal of Political Economy*, 86(

Hamann, A. J. 1997. “The Reform of the Pension System in Italy.” IMF Working Papers 97/18, IMF.

Hendry, D., A. Pagan and J.D. Sargan.1984. Dynamic specification, ch. 18, p. 1023-1100 in Z. Griliches and M.D. Intriligator, eds., *Handbook of Econometrics*, vol. 2, Elsevier

Hendry, D. F. 1995. *Dynamic Econometrics.* Oxford: Oxford University Press.

Jackman, R., and J. Sutton. 1982. ‘Imperfect Capital Markets and the Monetarist Black Box: Liquidity Constraints, Inflation and the Asymmetric Effects of Interest Rate Policy’, *The Economic Journal*, **92**, 108–28.

Jordà, Ò.2005. “Estimation and Inference of Impulse Responses by Local Projections.” *American Economic Review*, 95 (1): 161-182**.**

Jump, G. V. “Interest Rates, Inflation Expectations, and Spurious Elements in Measured Real Income and Saving.” *American Economic Review,* 70 (5) (1980): 990–1004.

Kaplan, G., and G. L. Violante. 2014. “A Model of the Consumption Response to Fiscal Stimulus Payments”, *Econometrica*, 82, 1199–239.

Kaplan, G., and G. L. Violante. 2018. “Microeconomic Heterogeneity and Macroeconomic Shocks”, *Journal of Economic Perspectives*, 32(3), 167–94.

Kaplan, G., B. Moll and G. L. Violante. 2018. “Monetary Policy According to HANK”, *American Economic Review*, 108(3), 697–743.

Kaplan, G., G. L. Violante and J. Weidner. 2014. “The Wealthy Hand-to-Mouth”, *Brookings Papers on Economic Activity*, Spring.

Lepinteur, A., A. Clark and C. D'Ambrosio. 2022. “Job insecurity, savings and consumption: an Italian experiment” , paper prepared for the 37th IARIW General Conference.

Muellbauer, J. 2020. “Implications of Household-level Evidence for Policy Models: the Case of Macrofinancial Linkages.” *Oxford Review of Economic Policy* 36 (3).

Muellbauer, J. 2022. “Real estate booms and busts: implications for monetary and macroprudential policy in Europe”, in *Challenges for monetary policy in a rapidly changing world*, proceedings, ECB Forum on Central Banking, Sintra. https://www.ecb.europa.eu/pub/pdf/sintra/ecb.forumcentbank202206~a6bc0541ca.en.pdf

Otsuka, M. (2004), ‘Essays on Household Portfolio and Current Account Dynamics’, PhD Thesis, Johns Hopkins University.

Pascal, P. 2024. "Heterogeneity in macroeconomic models: A review of theory and computation," BCL working papers 185, Central Bank of Luxembourg.

Reis, R. (2006), “Inattentive Consumers, Journal of Monetary Economics,53(8),1761-1800.

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

Data Appendix

A1. Consumption and income data.

Constant price consumption data comes from the Bank of Italy Quarterly Model (Bulligan et al., 2017) based on National Accounts (ISTAT). Basically, it is the sum of constant price expenditure on non-durable goods and services and expenditure on durables. The ratio of the current price sum to the constant price sum defines the consumer expenditure deflator. Per capita data are defined by deflating by quarterly population.

Total household disposable income also comes from the national accounts. Quarterly estimates of labour income based on payroll data plus transfer income data (mainly state pensions) are tax-adjusted using more detailed breakdowns of categories in the annual accounts. We implicitly assume the same tax rate for each form of income. The current price income data are deflated by the above consumer expenditure deflator and converted to a per capita basis.

A2. Adjusting for breaks in the asset data

As explained in Section 2, consistent quarterly portfolio data begin in 1995. Before then, annual data from Bonci and Coletta (2008) are quarternalised; most of the statistics are taken from the Italian financial accounts published by the Bank of Italy. Because of the move from the 1979 System of National Accounts (ESA79) to the 1995 System of National Accounts (ESA95, implemented in 2000 with time series starting in 1995), there are breaks in some series. Two general principles are followed in break-adjustment, the first being to adjust earlier data to match later data. The second is to use the average of immediate pre and post break growth rates to estimate what the value for the last pre-break observation would have been if it had grown at that rate relative to the first post-break observation. That value is then used to splice to the pre-break data. For example, there is a break in the pension data in 1984, when the data jumped from 16.53 at the end of 1983 to 47.56 in at the end of 1984 (Lire converted into Euros). The average of the pre and post break growth rates is 11.6%. Projecting back gives a revision of the end of 1983 observation from 16.53 to 42.63, so that the earlier data are revised up by 42.63/16.53 to be continuous with the post-break data.

Another break is in the loans data in 1989. As Bonci and Coletta (2008) point out, from 1989 to 1994, due to the new standards of ESA95, the sector consisting of households and non-profit institutions serving households was extended to include sole proprietorships with at most 20 workers. The increase in household loans from 93.426 to 202.975 from 1988 to 1989 is clearly not a genuine increase in household loans. The fact that NFC loans actually decrease in the same year against a strongly rising upward trend is a symptom of the reclassification. The average of the pre and post-break growth rates is 18.3% which suggests an upward revision of 78.150 to the end of 1988 loans data from 93.426 to 171.577. As a check, the same methods applied to NFC loans data would imply a closely consistent downward revision of 77.289.

From 1995Q1, consistent quarterly balance sheet data have been published by ISTAT and the Bank of Italy. These data, for the first time, split debt securities into short term – 'bills' and long term – 'bonds', both mainly in the form of government debt. As bills carry very little revaluation risk, one would expect them to be more liquid than bonds, which with Italy's volatile inflation and interest rate history, are far riskier. Internal data from the Bank of Italy on end of year levels of short-term and medium- and long-term securities before 1995 add to the information in Bonci and Coletta (2008).

There are breaks in the quarterly balance sheet data relative to the pre-1995 data from Bonci and Coletta (2008). These were corrected using the principles outlined above. This resulted in upward adjustments of pre-1995 data on deposits, mutual funds and quoted shares and downward adjustments for bills, bonds, unquoted shares and pension assets.

A3. Heterogeneity in asset ownership

Based on the data of the Survey of Household Income and Wealth (Bank of Italy, 2018), at the end of 2016, almost 70 per cent of Italian households owned their main residence; this proportion remained fairly stable compared with 2006. A one percentage point drop, to 52 per cent, in the share of property owners among households headed by someone 45 years or younger was offset by the sharp decline in the percentage of these households out of the total, from 37 to 27 per cent, continuing the trend that began in the early 1990s.

Regarding financial assets, the SHIW suggests that households belonging to the poorest quartile of the income distribution primarily have bank or post office deposits; the share of government securities, private sector bonds and managed investments (mainly investment funds) rises gradually across the income distribution; the richest 25 per cent of households are the ones most likely to directly own stocks and to entrust the management of a significant portion of their financial assets to investment professionals (Figure A1).

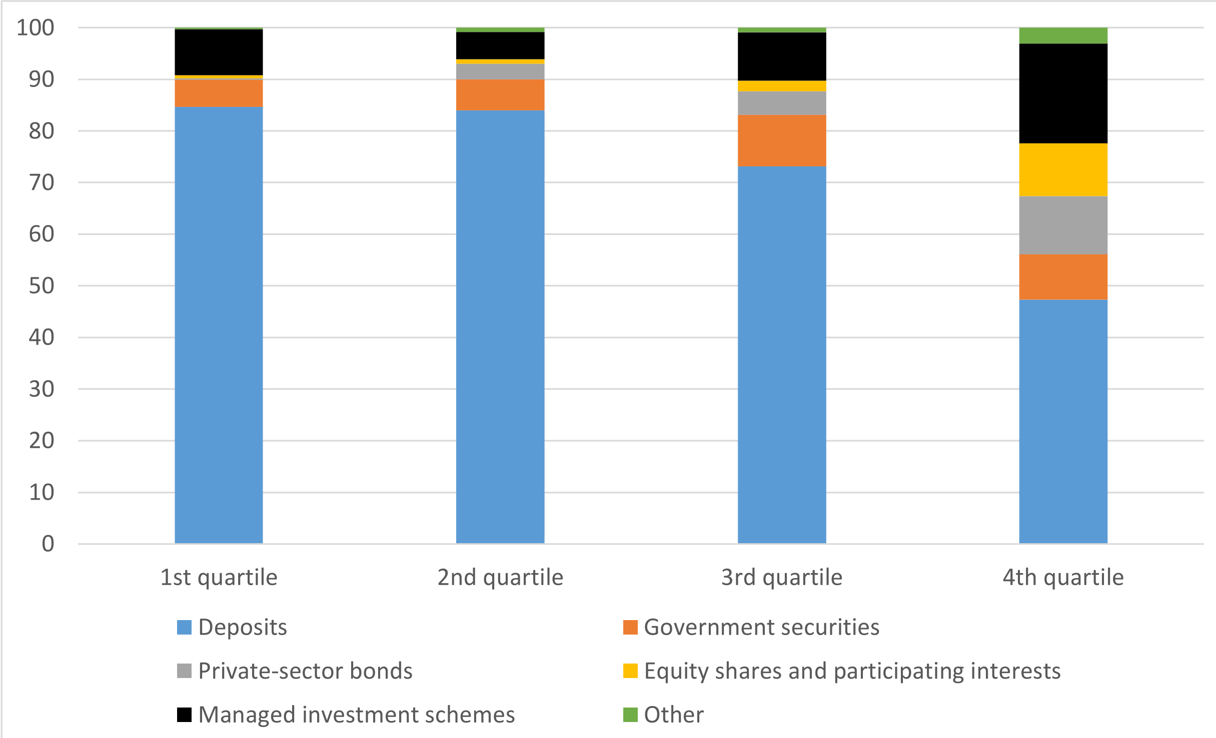

Figure A1: Breakdown of financial assets by income quartile distribution in 2016
Source: Authors' calculation from Survey of Household Income and Wealth (Bank of Italy)

Colabella et al. (2023) estimate that at the beginning of 2023 the amount of financial assets accumulated due to the pandemic exceeded 130 billion euros. While all households, including less affluent ones, managed to accumulate excess savings the majority of these funds (more than 60%) is held by the higher-income households. Figure A2 plots the distribution of financial assets at the end of 2020 (figures are taken from the SHIW in 2020, last wave available; Bank of Italy, 2022) by income quartile. It shows that for more affluent households, there was a substantial rise from end-2016 to end-2020 in the fraction of total financial wealth held in managed investment schemes. For the top income group, this displaced mainly government securities, private sector bonds and equity shares.

For aggregate behaviour, an interesting question is whether, for each asset aggregated across households, there was a notable change in the distribution by income group. For example, suppose that the share of liquid asset holdings by the top income group had increased sharply in the pandemic. Given that the MPC out of liquid assets is almost certainly lower for the top income group than for lower incomes, this would imply a fall in the aggregate MPC for liquid assets. Table A1 provides information from the SHIW for the asset classes used in that survey comparing end-2016 with end-2020. It shows little change in the distribution across income quintiles in liquid deposit holdings. For managed investment schemes and for 'equity shares and participating interests', there were falls in the share of the top quintile compensated by increases in the shares of the 3rd and 4th income quintiles. If anything, this would point to slight increases in the aggregate

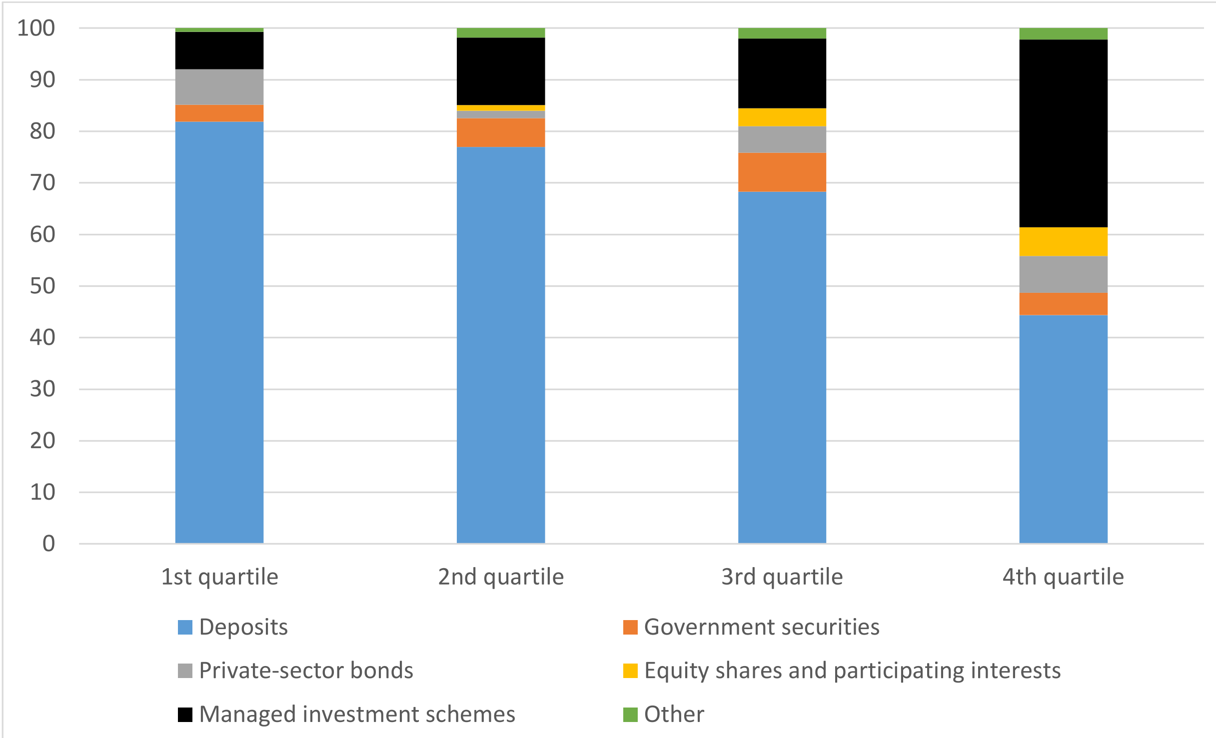

Figure A2: Breakdown of financial assets by income quartile distribution in 2020
Source: Author's calculation from Survey of Household Income and Wealth (Bank of Italy)

MPCs associated with these less liquid assets. It does suggest that there are no reasons to suppose that aggregate MPCs out of all financial assets have fallen as a result of the large changes in the structure of asset holdings due to the pandemic. However, evidence from the SHIW needs to be interpreted cautiously given the differences between balance sheet totals implied by the micro data and the national balance sheets. For example, for liquid deposits in 2016 and 2020, under 40 percent of the national balance sheet totals can be accounted for by the survey totals.[23] Shifts in the representativeness of the SHIW in the pandemic could have affected conclusions about changes in the distributional pattern of asset holdings.

Battistini et al. (2023) investigate potential composition effects on aggregate spending from the savings accumulated during the pandemic, comparing the US and Euro area. In the US, data suggest that the surge in savings was mostly accumulated in liquid assets, while in the Euro area the accumulation in illiquid financial assets was greater than in liquid assets. This would imply, other things being equal, a smaller post-pandemic consumption response in the Euro area than in the US. Battistini et al. (2023) combine aggregate balance sheet and distributional survey data to simulate potential compositional and distributional effects on aggregate consumption. They divide assets into two: deposits and illiquid financial assets minus debt. Our evidence from Italy

[23] Though some of the difference is likely to be due to the inclusion of non-profit institutions serving households in the national balance sheets.

is that grouping of semi-liquid assets and debt with illiquid financial assets would likely have distorted the conclusions of such a study in the case of Italy.

Table A1: **Comparing 2016 and 2020 shares of each asset held by different income quintiles**

	Deposits		**Government securities**		**Private-sector bonds**		**Equity shares and participating interests**		**Managed investment schemes**		**Other**	
WAVE IBF	2016	2020	2016	2020	2016	2020	2016	2020	2016	2020	2016	2020
1st quintile	3.94	3.22	0.52	1.73	0.03	2.99	0.24	0.03	1.91	0.26	0.38	0.53
2nd quintile	7.81	7.12	4.14	5.83	1.09	0.9	0.62	1.13	1.18	2.06	0.07	4.43
3rd quintile	13.99	13.68	8.82	12.17	2.86	4.65	0.51	5.14	3.88	5.99	5.49	11.28
4th quintile	22.07	22.44	21.02	20.27	20.94	15.07	6.21	15.82	11.65	17.57	7.42	9.61
5th quintile	52.19	53.54	65.5	59.99	75.08	76.4	92.41	77.88	81.39	74.12	86.63	74.15
	100	100	100	100	100	100	100	100	100	100	100	100

Source: Author's calculation from Survey of Household Income and Wealth (Bank of Italy)

A4. Measuring credit conditions

We have quarterly data on the ratio of granted to used credit lines for the whole period but regard the ratio of granted credit lines to a moving average of GDP as a better indicator of credit availability. From 1996, we have data on used credit lines, which we fit to data on the stock of bank credit advanced to the private sector at t and the change relative to t-4 indicating a mix of level and change effects. The adjusted R-squared is 0.983. We estimate granted credit lines before 1996 by multiplying the ratio of granted to used credit lines by the fitted value of used credit lines. Our credit indicator is estimated by credit lines divided by the lagged 8-quarter moving average of GDP in current prices. As granted credit lines is more of a stock affected by past economic conditions than a flow, this avoids temporary fluctuations in GDP distorting the measure. Figure A3 compares the ratios of granted credit lines respectively to used credit lines and to a moving average of GDP.

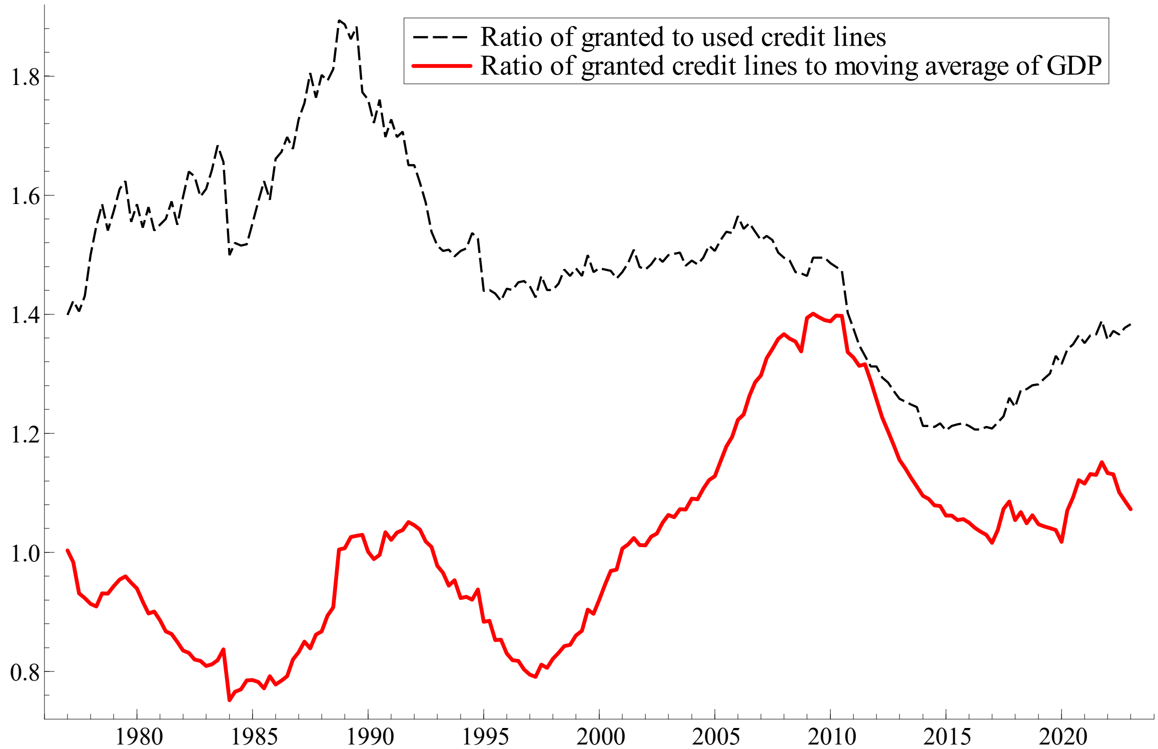

Figure A3: The ratios of granted credit lines to used credit lines and to a moving average of GDP.
Source: Credit register, Bank of Italy and national accounts data from ISTAT

The empirical evidence is that the lagged rate of change in the latter measure helps explain short-term fluctuations in consumption.

A5. Other short-term controls.

Turning to controls for the short-term dynamics, these include the 2-quarter change in the log of the credit conditions measure, in the log of the per capita measure of public consumption – a partial substitute for private consumption, and in the unemployment rate, see Figure A4.

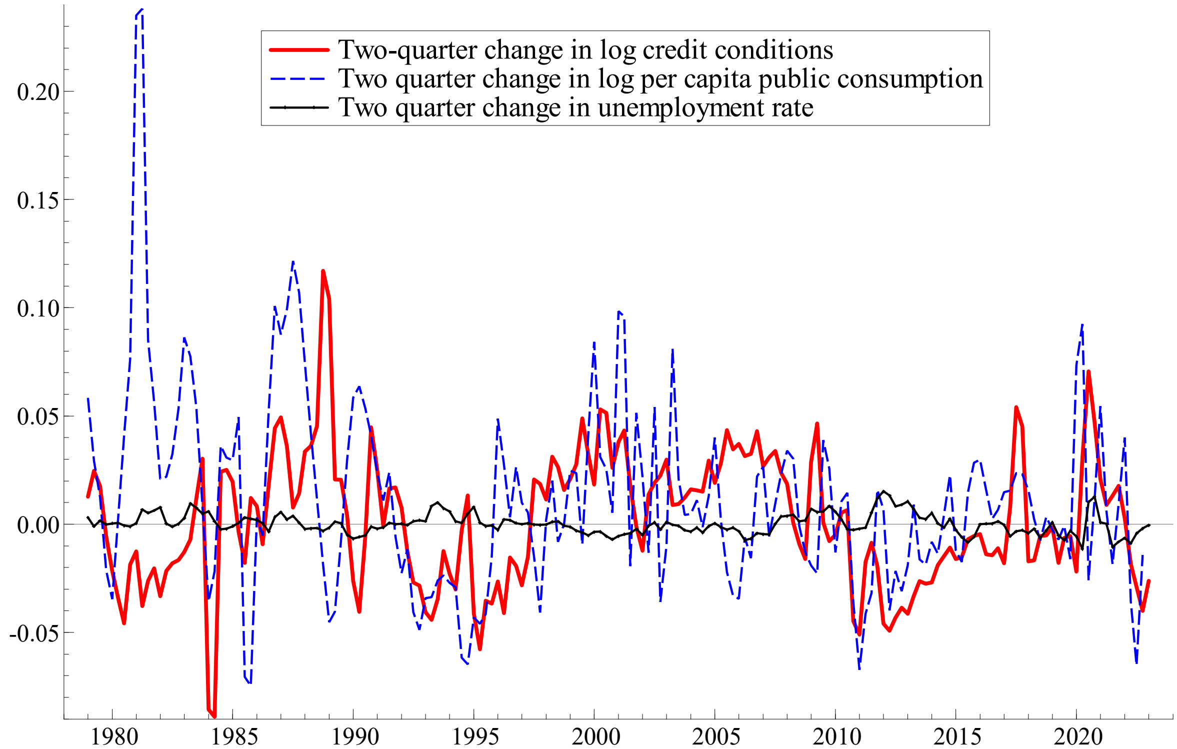

Figure A4: Two-quarter changes in the log of credit conditions, the log of per capita public consumption, and in the unemployment rate.
Source: See Figure A3, and ISTAT.

Other short-term controls include the change in the annual growth rate of income and a measure of income volatility (Figure A5). Income volatility is defined as the absolute value of the residual of the change in log income from an AR2 autoregressive process.

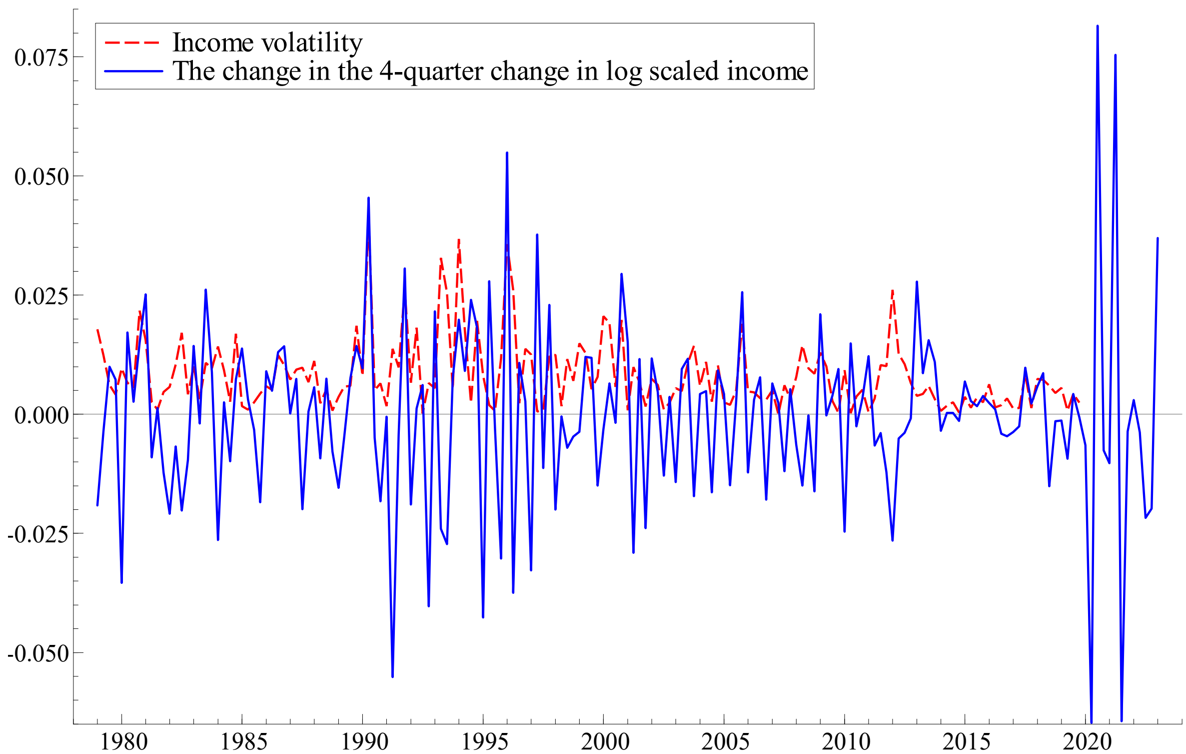

Figure A5: Income volatility and the change in the 4-quarter growth rate of income.
Sources: National Accounts data from ISTAT; income volatility is the absolute value of the residual from an AR8 process for log scaled income.

A6. The model for permanent income

The definition of log permanent income was given in equation 3.3 with the horizon k=40 quarters and the discount factor δ=0.95. A summary of the findings was given in Section 2. The following equation was estimated:

$$\begin{aligned}\ln y_t^p = {} & a_{y0} + a_{yt}t + a_{y08}t_{perm}^{2008:q3} + a_{yd08}dt_{perm}^{2008:q3} + a_y \ln y_{t,ma} \\ & + (1-a_y)\ln\left(\frac{labour_{t-1}}{pop_{t-1}}\right)_{ma} + a_{yMIB}\ln MIB_{t-1} + a_{yCOMP}COMP_{t-4,ma} \\ & + a_{yOIL}\ln OIL_{t-4,ma} + a_{yUR}UR_{t-1} + a_{yCCI4}\Delta_4 \ln CCI_{t-1} + a_{yUR4}\Delta_4 UR_{t-1} \\ & + a_{yOIL4}\Delta_4 \ln OIL_t\end{aligned}$$

The dependent variable is the log of permanent income. In order, the regressors include a time trend, the present discounted values of a trend shift in 2008Q3 and of a step change in the mean of the long-run solution in 2008Q3, both with significant negative coefficients. Next come the 4-quarter moving averages of log real per capita scaled income and of the log ratio of the labour force to the population. The latter coefficient is restricted so that the long-run effect of the latter on log income – combining permanent with current- is unity. This makes economic sense as our income variable is per capita relative to the population but the capacity to produce of the economy is constrained by the labour force, not population. The latter explains a good deal of the slow-down of permanent income growth in the early 1990s, as the figures below will show. The next regressor is the log of the real stock market index, followed by the moving average of a measure of competitiveness of the Italian firms, strongly related to the real exchange rate. It enters at a lag of 4 quarters, implying a slow response of future real household incomes to this factor. The log of real oil prices enters with a similar lag structure. High real oil prices damage the terms of trade of the economy for an oil importing economy and therefore affect living standards. The final term in the long-run solution is the level of the unemployment rate. This has a *positive* coefficient, which is at first sight surprising as one might have expected workers to have weaker bargaining power when unemployment is high and therefore have a lower wage share in GDP. However, GDP is not one of the controls in the equation. What the term suggests is that when the unemployment rate is high, there is a tendency for policies to be enacted that eventually stimulate growth in order to bring unemployment down. Similarly, when the unemployment rate is very low, policies tend to reverse to stabilise the business cycle.[24]

However, in the short-term dynamics, the lagged 4-quarter change in the unemployment rate has a significant negative effect probably because in the short run it results in lower wage growth, which, other things being equal lowers the growth rate of household income in the near future. The other terms in the short-term dynamics are the lagged 4-quarter change in the log of the credit conditions indicator and the 4-quarter change in the log of real oil prices. As large energy price shocks feed into the inflation rate faced by households and there tends to be nominal inertia in wages, near term real household incomes are negatively affected.

[24] It is also possible that endogenous mechanisms such as wage and price dynamics are affected by the unemployment rate and eventually feed into the growth rate of real incomes.

Table A2 Model for permanent income, 1980-2019

Parameter	Estimate	Standard Error	t-statistic
a_{y0}	1.479	0.060	24.59
a_{yt}	0.000	0.000	12.00
a_{y08}	-0.000	0.000	-3.17
a_{yd08}	-0.033	0.004	-7.63
a_{y}	0.206	0.022	9.33
a_{yMIB}	0.011	0.002	6.48
a_{yCOMP}	0.026	0.010	2.77
a_{yOIL}	-0.008	0.002	-3.94
a_{yUR}	0.383	0.047	8.14
a_{yCCI4}	0.053	0.011	4.88
a_{yUR4}	-0.419	0.070	-5.98
a_{yOIL4}	-0.005	0.001	-3.83
Diagnostics			
Standard error x 100	0.433		
Adjusted R^2	0.995		
LM Het. test (p-value)	0.11		
Durbin-Watson	0.22		

In this formulation, it is assumed that the structural break in the income process that occurred in 2008Q3, was fully anticipated and became more and more relevant in permanent income from 1998Q3 as the global financial crisis approached. The estimated model benefits from post crisis hindsight as the structural break is then incorporated in the formulation. To capture what might have been more feasible beforehand, we need to remove the over-optimism represented by the fitted estimated present value of the structural break before 2008Q3 to simulate information of which a sophisticated household proxied by an econometrician might have been aware. However, we then need to posit a learning mechanism. The failure of Lehman Bros in September 2008 and the global financial turmoil that followed was a profound shock that must have made households far less optimistic about future income growth. However, it would have taken some time to arrive at a view of whether this was more of a step change or a change in the underlying growth rate. Both factors are in the equation. We assume that learning was split equally between an instant reaction and gradual learning over the following 16 quarters. In the first 8 quarters before the sovereign debt crisis became serious, learning follows an ogive or S-shaped curve, where learning is initially slow but speeds up near the mid-point of the period and then asymptotes towards completion towards the end. In response to the sovereign debt crisis, from 2010Q3, it follows a new sigmoid curve to 2012Q2. Thus, gradual learning takes into account the two phases of the crisis, the GFC followed by the debt crisis. Learning is implemented by attaching a time-varying weight *w* to the formulation of fitted permanent income from which the present value of the structural break has been removed, and a weight of 1-*w* to the ex-post formulation that includes the structural break. The weight is zero before 2008Q3 and 1 from 2012Q2 and looks like this:

Table A3 Weights for learning adjustment

Quarter	Weight	Quarter	Weight
2008:q2	0.0000	2010:q3	0.7625
2008:q3	0.5125	2010:q4	0.7875
2008:q4	0.5375	2011:q1	0.8250
2009:q1	0.5750	2011:q2	0.8750
2009:q2	0.6250	2011:q3	0.9250
2009:q3	0.6750	2011:q4	0.9625
2009:q4	0.7125	2012:q1	0.9875
2010:q1	0.7375	2012:q2	1.0000
2010:q2	0.7500		

Thus, by 2012Q2, over-optimism has been eliminated and the model has fully incorporated the down-shift in the path of income. Instead of an equal weight on fast and slow learning, variations in the range 0.3 to 0.7 on the relative weight of the two have almost no impact on the estimated parameters of the consumption function and only a marginal effect on the goodness of fit.

The contributions of the different regressors to the fitted, learning adjusted log permanent income are shown in Figures A6 to A8. Figure A6 shows that much of trend movement of log permanent income is accounted for by the time trend, the log of the labour force relative to the population and some by the movement of actual log current income. Much of the abrupt fall following 2008Q3 is accounted for the learning adjusted split trend.

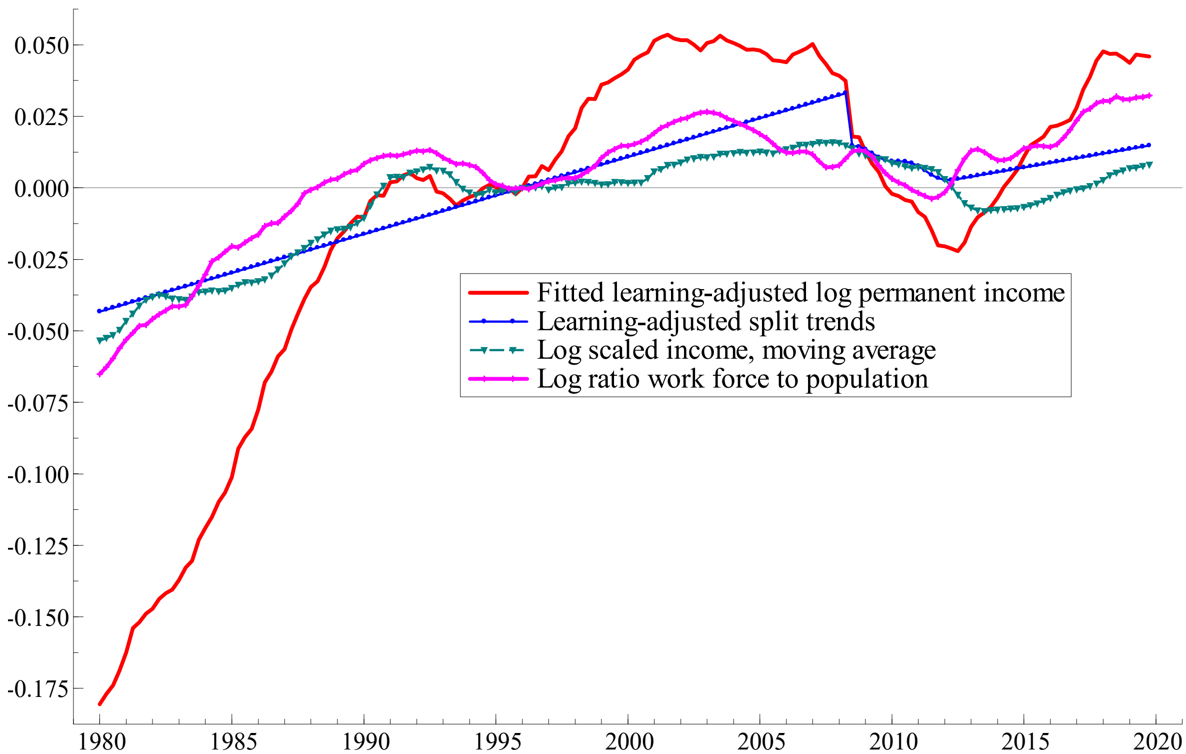

Figure A6: The fitted contributions to log permanent income of shifting trends, log scaled income and the log ratio of work force to population.

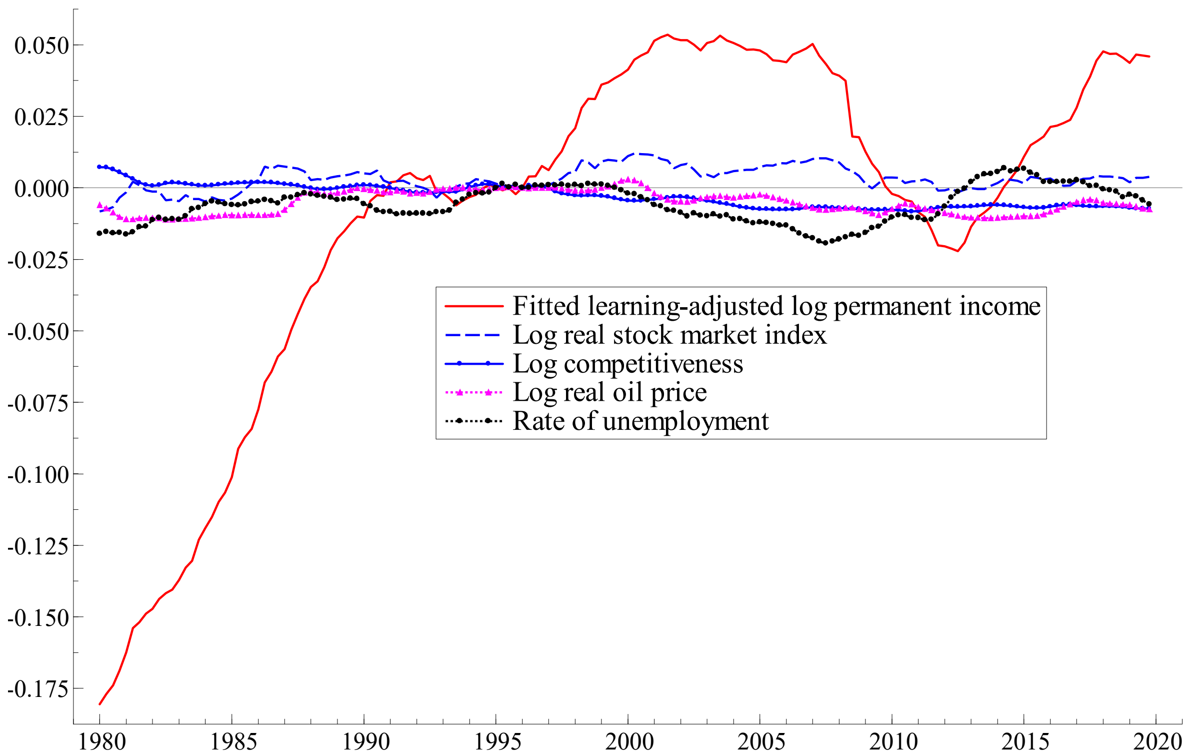

Figure A7: The fitted contributions to log permanent income of the stock market index, competitiveness, the real oil price and the unemployment rate.

Figure A8 shows the fitted contribution of the three short-term regressors: the rate of growth of credit conditions, the change in the unemployment rate and the rate of change of real oil prices.

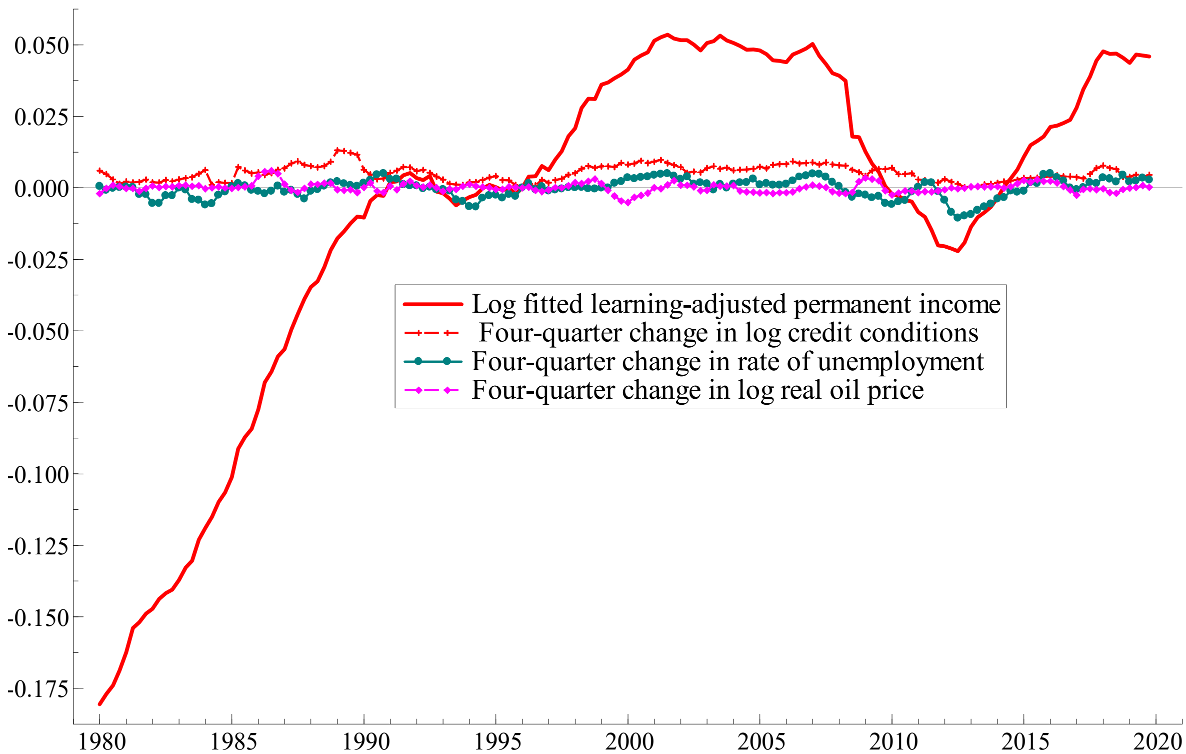

Figure A8: The fitted contributions to log permanent income of three short-term regressors.

A7. Instrumenting current income

The log of current real per capita scaled income was instrumented using a regression on its lag, and lags of log consumption, the unemployment rate, the log of the labour force to population ratio, competitiveness and in the acceleration of inflation. When log current income is instrumented in the consumption function, it is instrumented by its fitted value. Where the level of current income appears, for example, in ratios of assets to income, current income is replaced by the exponential of fitted log income.

$$\ln y_t = a_0 + a_y \ln y_{t-1} + a_c \ln c_{t-1} + a_{UR} UR_{t-1} + a_{LAB} \ln\left(\frac{labor_{t-3}}{pop_{t-3}}\right) + a_{COMP} COMP_{t-6,ma} + a_{\pi 2} \Delta\pi_{t-2} + a_{\pi 3} \Delta\pi_{t-3}$$

Table A4 Model for instrumenting income, 1980-2019

Parameter	Estimate	Standard Error	t-statistic
a_0	0.418	0.102	4.09
a_y	0.781	0.039	20.03
a_c	0.102	0.026	3.86
a_{UR}	-0.205	0.064	-3.20
a_{LAB}	0.344	0.079	4.33
a_{COMP}	0.041	0.011	3.78
$a_{\pi 2}$	-0.226	0.105	-2.15
$a_{\pi 3}$	-0.314	0.105	-3.00
Diagnostics			
Standard error x 100	0.010		
Adjusted R^2	0.986		
LM Het. test (p-value)	0.36		
Durbin-Watson	2.27		
AR1/MA1 (p-value)	0.054		
AR4/MA4(p-value)	0.180		
Chow test (p-value)	0.019		
RESET2 test (p-value)	0.000		
F test (p-value)	0.000		
Schwarz Criterion	-493.3		
Log likelihood	513.6		

A8. Descriptive statistics

Table A5 Descriptive statistics 1980-2019

Variable	Mean	Std. Dev.	Minimum	Maximum
Raw data before transformations (current prices, billions)				
Consumption	170.9	76.2	30.5	272.4
Disposable income	199.0	79.4	39.1	297.4
Labour and transfer income	103.7	47.4	19.0	179.4
Deposits	741.3	398.3	107.6	1448.2
Loans	343.2	267.0	18.5	737.8
Mutual funds	272.1	218.8	0.5	677.3
Bonds	346.8	262.0	5.5	755.8
Bills	66.8	52.1	1.8	182.2
Pensions and life insurance	404.5	330.3	25.2	1119.4
Quoted shares	84.5	39.1	7.6	208.8
Unquoted shares	506.9	343.2	20.4	1242.4
Data for consumption equations				
Log per capita consumption	1.388	0.130	1.092	1.529
Log per capita scaled income	1.291	0.084	1.054	1.393
Change in log per capita consumption	0.0026	0.0065	-0.0158	0.0211
Change in log per capita scaled income	0.0018	0.0114	-0.0337	0.0448
Real effective borrowing rate	0.094	0.017	0.028	0.150
Log learning-adjusted permanent income	1.316	0.061	1.144	1.378
Net liquid assets/scaled income	1.175	0.217	0.791	1.595
Bonds/scaled income	0.473	0.305	0.040	0.932
Illiquid financial assets/scaled income	0.683	0.294	0.276	1.297
Housing assets/scaled income	5.030	1.055	3.431	6.962
Log ratio of house prices to income	5.923	0.121	5.645	6.116
Net worth/scaled income	8.080	1.835	5.219	10.585
Net financial assets/scaled income	3.051	0.929	1.326	4.361
Data for short term controls				
Two-quarter change in log credit conditions	0.0011	0.0311	-0.0889	0.1170
Two-quarter change in log per capita public consumption	0.0121	0.0460	-0.0746	0.2381
Change in 4-quarter change in log scaled income	0.0002	0.0163	-0.0551	0.0549
Two-quarter change in unemployment rate	0.0003	0.0044	-0.0084	0.0152
Income volatility	0.0080	0.0074	0.0000	0.0414
Data for permanent income equation				
Log permanent income	1.31	0.06	1.14	1.38
Log scaled income (4-quarter MA)	1.29	0.09	1.05	1.39
Log labour force/population (4-Quarter MA)	-0.88	0.03	-0.96	-0.84
Log real stock market index	9.83	0.39	8.82	10.64
Competitiveness (4-quarter MA)	1.10	0.14	0.91	1.39
Log real oil price index (4-quarter MA)	3.75	0.52	2.71	4.54

Unemployment rate	0.10	0.02	0.06	0.13
4-quarter change in log credit conditions	0.00	0.05	-0.09	0.15
4-quarter change in unemployment rate	0.001	0.008	-0.012	0.025
4-quarter change in log real oil prices	-0.01	0.30	-1.12	0.96