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CONSUMPTION AND WEALTH: NEW EVIDENCE FROM ITALY

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Abstract

This paper estimates a consumption function for Italy. In addition to permanent income, housing wealth, the interest rate on household loans and an index of credit conditions, our model introduces household net worth split into liquid and illiquid assets. The consumption dynamics are examined by using financial accounts and real national accounts in a Vector Error Correction Model (VECM), estimated from 1975 to 2017. The results show that the marginal propensity to consume out of liquid financial assets – mainly deposits and bonds – is positive and statistically significant, and greater than that for illiquid assets (mainly unquoted shares and insurance and pension assets); we also find that housing wealth has a smaller and significant impact on consumption. As expected, permanent income accounts for a large fraction of consumption, while the effect of the interest rate is negative.

JEL Classification: E21, E32, E44, E51.
Keywords: Liquid and illiquid assets, permanent income, credit conditions.
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Contents

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1 Introduction[1]

There is widespread disagreement about the influence of wealth on households' consumption (Buiter, 2010; De Bonis and Silvestrini, 2012; Cooper and Dynan, 2016). Even the recent pandemic of COVID-19 has renewed the debate on the channels through which wealth, housing as opposed to financial assets, may affect consumer spending. Across the world, non standard monetary policies and credit conditions influenced asset values. A reduction of interest rates and a relaxing of credit constraints may increase household consumption relative to their income (Jappelli and Pagano, 1994) and reduce downpayments (Balta and Ruscher, 2011; Liberati and Vacca, 2016). Moreover, lower interest rates increase the value of collateral-backed loans for households that already own the collateral (Poterba and Manchester, 1989; Miles, 1992). So, as suggested by Muellbauer (2020), a macro model having the object to capture the functioning of different economic channels taking into account different risks and behaviours, should include a detailed household-housing sub-system.

In this paper we estimate a consumption function for Italy, splitting up household net worth into liquid and illiquid components and taking into account the role of housing assets, permanent income and credit conditions. The Italian economy historically shows higher wealth accumulation and saving rates than other countries (De Bonis and Marinucci, 2017) although with a convergence in the latest years. There is no general consensus about housing wealth effects on consumption in Italy. These are positive and rather small according to Catte et al. (2004) and Guiso, Paiella, and Visco (2006), and sometimes even negative (Boone and Girouard, 2002; Slacalek, 2009). On the other hand, financial wealth effects are stronger and statistically significant than housing ones (Bassanetti and Zollino, 2010).

Many macro models consider net worth as a single variable having a unique effect on consumption, but its components have different degree of liquidability. Obtaining liquidity from housing wealth (using mortgage equity withdrawal; MEW onward) needs more time than exploiting financial assets; moreover MEW does not exist in many countries. Notwithstanding the presence of booms and bursts, the ratio of Italian household financial wealth to GDP increased in the last 40 years. However, housing wealth remains the main asset for Italian households: this is common to many advanced economies (see Caprara, De Bonis, and Infante, 2020).

Propensity to consume is typically heterogeneous in households' population due to income shocks, liquidity constraints and different assets' taxation. A recent strand of literature accounts for this feature focusing on the role of tax rebats (see e.g. Kaplan and Violante, 2014; Huntley and Michelangeli, 2014). Of course, policy implications may arise: Neri, Rondinelli, and Scoccianti (2017) estimate that Italian households receiving recent tax rebates introduced in 2014 increase their consumption: the rise is larger for families with low liquid wealth or low income. Differently from previous cited works, due the nature of the model and of the data, we abstract from heterogeneity issues.

[1]The views expressed herein are solely those of the authors and do not necessarily reflect the views of the Bank of Italy. We wish to thank G. Albareto, A. Brandolini, F. Lilla, G. Marinelli, A. Silvestrini, R. Zizza and F. Zollino for their useful suggestions. 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 $60th$ Annual Scientific Conference held at Palermo University, 24-26 October 2019, for their useful comments.

Different types of income shocks could lead to substantial differences in consumption responses. According to the simplest version of the permanent income hypothesis, only unanticipated permanent income shocks should induce substantial changes in consumption. On the contrary, expected or temporary income shocks should not alter consumption significantly. Then, effects on consumption depend on the households' perception about the transitoriness/permanence of the shocks (Jappelli and Scognamiglio, 2016). For this reason, in our analysis we control for the impact of the main recessions experienced by the Italian economy that could be recognized as shocks to the permanent income (Miniaci and Weber, 1999; Rodano and Rondinelli, 2014; Brandolini, Gambacorta, and Rosolia, 2018).[2]

Credit market conditions could affect aggregate consumption through different channels, as interest rates (i.e. mortgage or bond rates), credit limits (i.e. loan-to-value ceilings), debt renegotiations. The recent drop in interest rates following the Great Recession produced two opposite effects: it reduced mortgage payments and lowered financial assets returns. Thus, the effect of credit market conditions on consumption depends on the reaction of households balance sheets. In our analysis the Italian credit conditions for the private sectors are proxied by the ratio of the used credit lines to the granted ones, a measure that tracks very well the credit conditions over the two recent crises.

Our contribution to the literature is threefold. First, we estimate the effects of liquid and illiquid financial wealth on consumption using the Italian financial accounts, based on the assumption that the higher the degree of liquidability of an asset the higher is expected to be its effect on expenditure. Second, to estimate long run effects, we reconstruct the quarterly relevant variables for financial wealth back to 1975. Third, we set out for the Italian economy the Friedman-Ando-Modigliani basic aggregate life-cycle/permanent income consumption model where the consumption function depends not only on past wealth, but also on an estimate of real permanent income, as in Aron and Muellbauer (2013); this is a more robust assumption than the Euler equation that assumes households to be continuously and efficiently trading off between consuming now and consuming in the next period. In our setting, a household wishing to sustain consumption will realize that not all of its assets can be spent now without damaging future consumption, and that future income has a bearing on sustainable consumption.

Aron et al. (2012) and Aron and Muellbauer (2013) employ a modified Ando and Modigliani (1963) consumption function which incorporates permanent income, income uncertainty, housing collateral and other credit effects:[3] estimation of consumption for U.S. highlights positive and significant wealth effects for all the components of the net worth, including the housing one. Following a similar approach Muellbauer, St-Amant, and Williams (2015) find that housing collateral effects on consumption are absent in Canada. For the euro area, Sousa (2010) estimates large significant financial wealth effects and nil and not significant effects of the MEW. Similar results are obtained by Slacalek (2009) who stresses that wealth effects - in particular the housing ones - are greater in countries with more developed mortgage markets. Consistent with this view Andersen and Leth-Petersen (2019) show how housing wealth gains in Denmark are related to the efficient working of the mortgage market. There

[2]See also Grant, Miniaci, and Weber (2002) and Bassanetti et al. (2009).

[3]In countries where credit conditions indexes are not available a "latent interactive variable equation system" (LIVES) is employed. For more details see Duca and Muellbauer (2014).

is a consensus on the fact that mortgage equity withdrawal is important in the U.S., the U.K. and the Netherlands while is absent or smaller in other European countries.[4]

In a nutshell, we find that in our model income plays the lion's share in explaining household consumption attitudes over the past forty years and that, as expected, the effect of the interest rate is negative in most of our regressions. About the wealth effects, we contribute to fill the gap that plagues the existing literature about the marginal propensity to consume out of financial wealth. An increase in liquid financial wealth, like deposits and bonds, rises household consumption by about 7.5 per cent; this effect is twice as large as that for illiquid financial wealth. Housing assets are also associated with an increase in consumption, but due to their lower degree of liquidability, the estimated effect is smaller (less than 2 per cent). All in all, these results are broadly consistent with Bassanetti and Zollino (2010), who found that the marginal propensity to consume out of housing and non-housing wealth is in the range of, respectively, 1.5-2 and 4-6 per cent.

The paper is structured as follows. Section 2 describes the data. Section 3 explains the theoretical and empirical frameworks. The main results on the role of the different of assets in explaining the households' consumption dynamics are reported in Section 4. Robustness estimations about the roles of the permanent income and interest rates are reported in Section 5. Finally, Section 6 concludes. In the appendices we report more detailed statistics and the model used to estimate the permanent income equation.

[4]See Barrell, Costantini, and Meco (2015) for a comparison of the housing and financial wealth effects between the U.K. and Italy.

2 Data and descriptive statistics

We use quarterly data from 1975:q1 to 2017:q4 (see Tables A.1 and A.2). Our dataset mainly relies on the Italian financial accounts.[5] We split up financial assets into liquid and illiquid ones. Net liquid financial assets (NLA) is the difference between gross liquid assets (GLA), which include deposits, bonds, mutual funds and quoted shares, and total liabilities ($LOANS$). Illiquid financial wealth is the sum of unquoted shares and other equity plus holdings of insurance and pension fund instruments (IFA).[6] In 2017 household net financial wealth was twice Italian GDP against 70 per cent in 1975 (figure 1). Until the late 1990s, this dynamics follows that of the net liquid component, while in recent years the illiquid component gained more importance, as Italian households increased their holdings of insurance and pension fund instruments. Housing and land wealth (HA)[7] are the main assets for Italian households: in 2017 they were more than 3 times GDP (figure 2). The dynamics of this ratio is strongly correlated with housing prices.[8]

Between 1975 and 2017 the ratio between consumption and GDP remained quite constant; additionally, there was a decline of total household disposable income, due the very low growth of the Italian economy in the last 25 years (figure 3).

In our analysis we include the real mortgage rate. This is a proxy of the cost of credit for households. Finally, credit conditions can be approximated by the ratio between the used credit lines and the granted ones based on the Bank of Italy's Central Credit Register: a decrease of the ratio indicates a credit easing and viceversa; indeed this measure tracks the last two crises quite well. Figure 4 shows the goodness of our credit index when compared with the dynamics of the real mortgage interest rate. For the sake of simplicity, when we run the model, we define a general credit conditions index ($GCCI$) as the opposite of the previous measure so that an increase of the index is interpretable as a credit easing.

[5]Before 1995 Italian financial accounts are only available at annual frequency (see Bonci and Coletta, 2008): backward estimation is obtained by using temporal disaggregation methods available by the authors upon request (see also Bruno, 2008).

[6]The inclusion of stock values is consistent with a large strand of literature on this topic (see e.g. Edison and Sløk, 2002; Ludwig and Sløk, 2004; Case, Quigley, and Shiller, 2005).

[7]From 2001 official annual data on Italian housing and land wealth are provided by Istat. Before this period we use annual estimation by Cannari, D'Alessio, and Vecchi (2017). Then quarterly data are obtained by using temporal disaggregation methods available by the authors upon request.

[8]The apparent no correlation between the real assets and the housing prices before 2000s depends on the use of the GDP to scale real assets. If we compare the nominal housing and land assets and nominal prices the unconsistency disappears (see figure A.1). Figure A.2 reports the logarithmic change of the consumption and the wealth components: in recent years housing assets showed a less variation than financial assets.

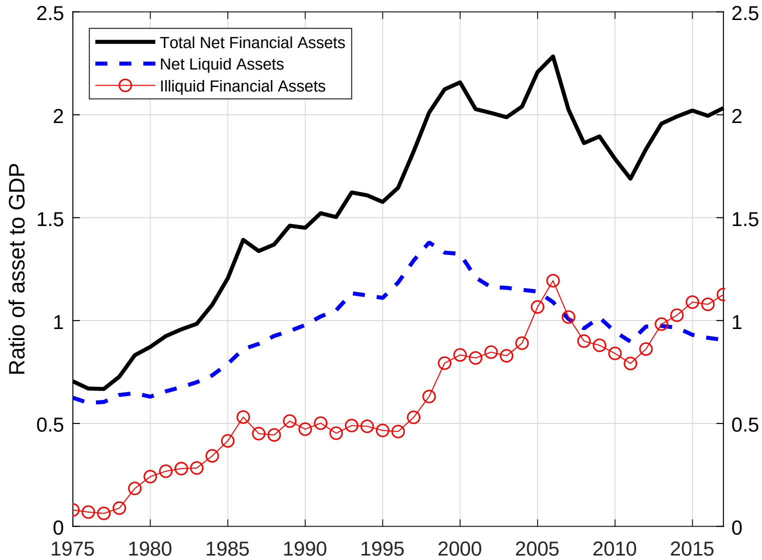

Figure 1: **Italian financial assets as ratios to GDP.**

Source: Bonci and Coletta (2008), Financial Accounts and Istat.

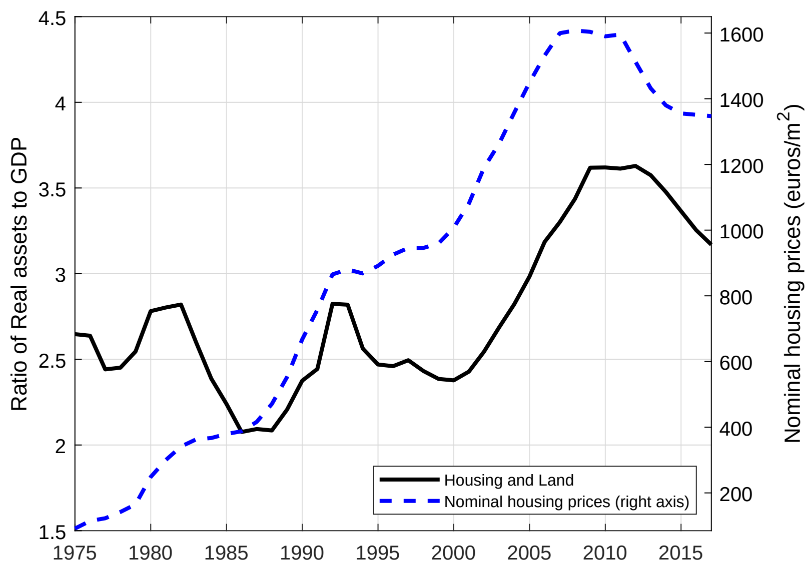

Figure 2: **Household real assets as ratio to GDP.**

Source: Cannari, D'Alessio, and Vecchi (2017), Istat, Bank of Italy, BIS and Revenue Agency.

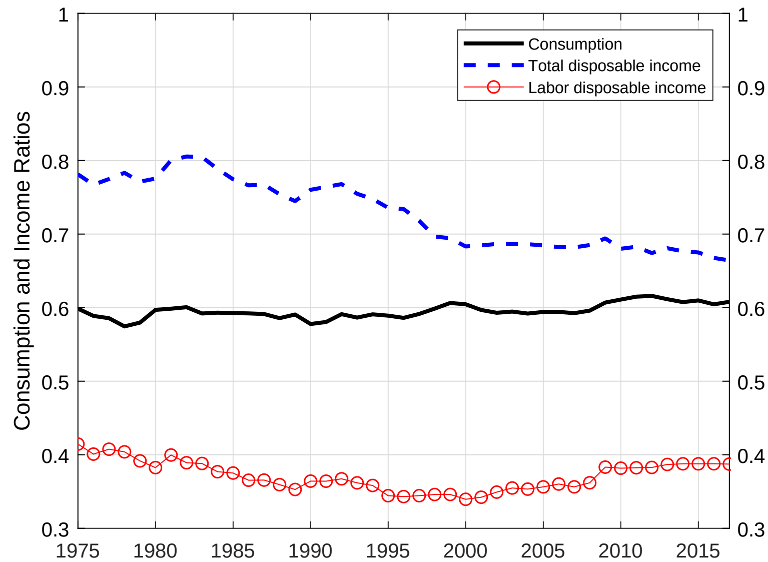

Figure 3: **Consumption and Incomes as ratio of GDP.**

Source: Istat and Brandolini, Gambacorta, and Rosolia (2018).

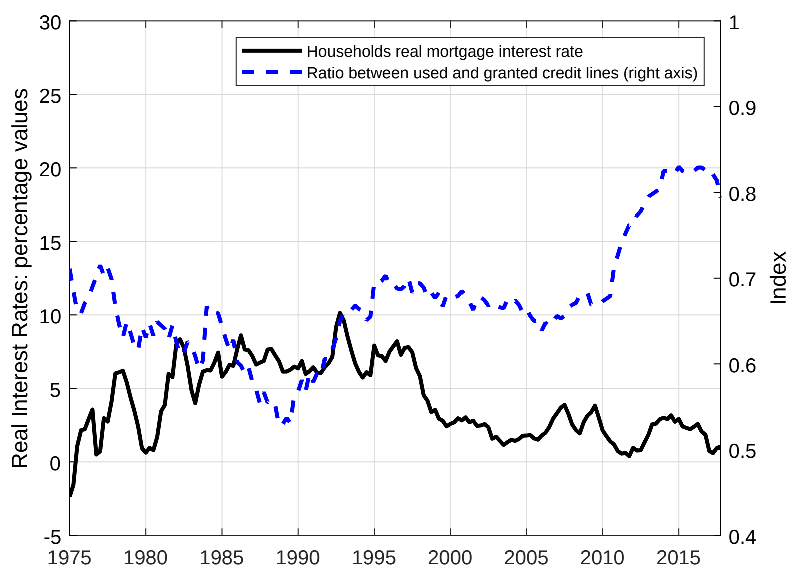

Figure 4: **Credit conditions in Italy.**

Source: Bank of Italy.

3 The Model

3.1 The theoretical framework

In this section we build up a permanent income consumption function in line with the contribution by Aron et al. (2012) who follows the insights of the Friedman-Ando-Modigliani formulation. In the basic life-cycle model, the aggregate consumption function with permanent income, when the real interest rate is not taken into account, assumes the following form:

$$c_t = \gamma^* A_{t-1} + \omega^* y_t^P \tag{1}$$

where c_t is the real per capita consumption, y_t^P is the permanent real per capita income and A_{t-1} is the real per capita level of net wealth of the previous period. Equation (1) requires a forecasting estimate of y_t^P.

Since consumption and income tend to grow exponentially, formulating the consumption function in logs has advantages. By defining y_t the current real per capita income, after some manipulations, the log approximation of equation (1) can be written as:[9]

$$\ln c_t = \alpha_0 + \ln y_t + \gamma \frac{A_{t-1}}{y_t} + \ln \left(\frac{y_t^P}{y_t}\right) \tag{2}$$

where $\gamma = \gamma^*/\omega^*$ and $\alpha_0 = \ln \omega^*$.

The log ratio of permanent to current income, which reflects expectations of income growth, must be estimated. To do it, in line with (Campbell, 1987), we initially compute a measure of the permanent income defined as a moving average of forward-looking real per capita income, and successively, we regress this value on relevant economic and demographic variables. Hence, at the first stage, the log ratio of permanent to current income can be proxied by functions of forecasted income growth rates as follows:

$$\ln \left(\frac{y_t^P}{y_t}\right) \approx \ln \left(y_t^P\right) - \ln \left(y_t\right) = \frac{\sum_{s=1}^{k} \delta^{s-1} \mathbb{E}_t \ln \left(y_{t+s}\right)}{\sum_{s=1}^{k} \delta^{s-1}} - \ln \left(y_t\right) \tag{3}$$

where δ is a discount factor equal to 0.95; in line with Chauvin and Muellbauer (2018) that use this length for France, k is set to be equal to 40, i.e. a time horizon of 10 years.

We initially proxy the disposable non-property income by using the disposable labor one (y^{lab}). Furthermore, we believe that other components of income may affect household consumption choices. Then, we test the possibility that labor income might be mis-measured mainly due to problems of the estimation and allocation of the self-employment and mixed incomes (see Blinder and Deaton, 1985) by using different combinations of the labor and total disposable incomes. Formally, we compute the permanent income as follows:

$$y^P = f\left(y\right) \quad \text{where} \quad y = \mu y^{lab} + \left(1 - \mu\right) y^{tot} \tag{4}$$

Textbook stories usually assume $\mu = 0$ or $\mu = 1$. In this work we calibrate $\mu = 0.5$, i.e. a value consistent with the average share of the Italian households disposable labor income

[9] See appendix A.2 for a complete description of the model.

to the total disposable one.[10] End-of-sample problems due to the discount of future income are overcome by assuming a quarterly growth rate equal to that of the previous period. Shorter time horizons are also suggested by a large strand of literature when households anticipate future credit constraints, according to the buffer-stock theory of saving (Deaton, 1991). Precautionary behavior also generates buffer-stock saving, as in Carrol (2001a,b), where it is argued that plausible calibrations of micro-behavior can give a practical income forecasting horizon of about three years. This horizon was originally suggested by Friedman (1957, 1963) in his application of the permanent income hypothesis to aggregate consumption data. In Section 5 we check our results by using $k = 12$ and $k = 20$, i.e. an horizon of 3 and 5 years, respectively.

We believe that a correct definition of the permanent income must take into account economic and demographic variables. Following Chauvin and Muellbauer (2018), we regress the the log ratio between permanent and current incomes computed in equation (3) on a set of economic and demographic variables, introducing a double split trend corresponding to the 1992-1993 and 2007-2009 crises. We assume that the post 1993 and the post 2009 slowdowns were not foreseen: economic agents learn gradually about the recessions (the details are reported in appendix A.3).The fitted values of this estimation are those used in our model.

After some arrangements, when real interest rates are variable, the change of log consumption can be approximated as:

$$\Delta \ln c_t \approx \lambda \left[\alpha_0 + \alpha_1 r_t + (\ln y_t - \ln c_{t-1}) + \alpha_2 \ln \left(\frac{y_t^P}{y_t} \right) + \gamma \frac{A_{t-1}}{y_t} \right] + \varepsilon_t \tag{5}$$

where λ measures the speed of adjustment, i.e. the speed at which a dependent variable returns to long-run equilibrium after modifications of the other variables.

The previous formulation can be improved along different directions. First, we can split up net wealth into three categories based on the degree of liquidity. Second, we can test the existence of a shift of the consumption-income ratio due to credit conditions. Third, we can introduce inside the cointegration space permanent shocks to control for the possibility of level shifts in the long-run equilibrium relationship. Finally, it is possible to impose short-run effects and dummies to take into account the effects of special events and temporary shocks. Accordingly, equation (5) can be “augmented” in the following way:

$$\begin{aligned} \Delta \ln c_t \approx \lambda \Bigg[& \alpha_{0t} + \alpha_1 r_t + (\ln y_t - \ln c_{t-1}) + \alpha_2 \ln \left(\frac{y_t^P}{y_t} \right) \\ & + \gamma_{1,gla} \frac{GLA_{t-1}}{y_t} + \gamma_{1,loans} \frac{LOANS_{t-1}}{y_t} + \gamma_2 \frac{IFA_{t-1}}{y_t} \\ & + \gamma_3 \frac{HA_{t-1}}{y_t} + \gamma_4 \ln \frac{HP_{t-1}}{y_t} + \tau_1 d_{92} + \tau_2 d_{95} + \tau_3 d_{11} \Bigg] \\ & + \beta_1 \Delta \ln y_t + \beta_2 \Delta \ln c_{t-4} + \beta_3 \Delta \ln c_t^P + \beta_4 \Delta_4 GCCI_{t-1} + \varepsilon_t \end{aligned} \tag{6}$$

where $\alpha_{0t} = \alpha_0 + \alpha_{0c} GCCI_t$ and $GCCI_t$ is a general credit conditions index. GLA stands for gross liquid assets whereas $LOANS$ represents the total liabilities (when we use net liquid

[10] The definition of households disposable labor income includes employments'compensation, social benefits (mainly pension transfers) and taxes and social contributions (see table 5.1 in Bank of Italy, 2018).

assets – NLA – the associoted coefficient is γ_1); IFA is the illiquid financial wealth. HA is the housing and land wealth whereas HP means real house prices. As explained by Aron and Muellbauer (2013), splitting wealth into housing and financial ones or based on the degree of liquidability is an important issue for different reasons: i) houses represent asset and utility values at the same time whereas the financial wealth owns only the first property; ii) in presence of credit constraints, land and housing wealth (HA) acts as collateral; iii) usually, illiquid financial assets (IFA) suffers more asset volatility than the liquid one (NLA). In our model, better credit conditions (captured by an increase of $GCCI$) may raise the intercept of equation (6) determining a higher level of the consumption-income ratio, mainly because of reduced saving for a housing downpayment which can be affetcted by the house prices dynamics (HP).

Among the short-run effects we consider the change of the labor disposable income ($\Delta \ln y_t$), the change of the real per capital public spending (Δc_t^P) and the 4-quarter variation of the lagged credit condition index $\Delta_4 GCCI_{t-1}$. We include public spending to test for the possible crowding-out effects of private consumption.[11] We consider the quarterly change of 4-quarters lagged consumption to handle the residual autocorrelation of the model.

Recessions could be permenantly affect income and consumption profiles. Sovereign debt crisis (10 quarters from 2011:q3 to 2013:q4) lasted longer than previous slumps, in particular with respect to the 1992 currency crisis (6 quarters from 1992:q3 to 1993:q4). Then we test if the length of the crises affects our estimates by introducing step dummies rather than a dummy variable indicator for the events' occourences. Therefore, d_{92} and d_{11} are equal to 1 if the Italian economy is in recession, according to the official dating by Istat, and 0 otherwise.[12]

In August 1995 the so-called Dini's reform radically changes the Italian pension system, mainly by introducing the actuarial approach in order to determine pension benefits (see e.g. Hamann, 1997).[13] Actually, the imposition of 2-quarter lagged dummy implies a better fit than the use of 1995:q3 dummy: as argued by Franco (2002), the reform generated uncertainty and induced elderly workers to retire from the workforce as soon as they were allowed to, for fear of possible cuts in their benefits. Moreover, starting from 1995:q1, pure quarterly data of financial accounts are evaluated at the fair value (ESA 1995 standards); before this date, annual data are available only at the nominal value (ESA 1979 rules). This change particularly affected the dynamics of the value of bonds and shares.[14] Then, we test if these events affect our findings by introducing a shift dummy, d_{95}, equal to 1 until the 1994:q4 and equal to 0 after this period.

With these ingredients, we test the existence of different marginal propensities to consume for liquid and illiquid assets, and for housing (γ_1 and γ_2 versus γ_3), and the possible presence

[11] The public spending has been deflated with the deflator of public consumption.

[12] We also test for the presence of a dummy variable taking into account the global financial crises which is never significant.

[13] The reform process began in 1992 by the so-called Amato's reform. Then, after the main and major intervention of the 1995, other minor measures were taken in subsequent years: the Maroni's reform in 2005 and the Fornero's reform in 2011.

[14] Among the changes of the European System of Accounts during our sample period (1979, 1995 and 2010), the implementation of $ESA95$ was the most relevant: it introduced in national accounts the market value and accrual basis criteria to evaluate assets. See Bonci and Coletta (2008) for more details.

of intercept shift stemming from changes in credit facilities.

3.2 The empirical model and the cointegration analysis

In this Section we estimate a Vector Error Correction Model (VECM) to infer the long run effects of wealth on consumption based on equation (6). Our VEC model abstracts from deterministic components outside the cointegration relationship and can be represented by the following formula:

$$\Delta Z_t = \underbrace{\lambda}_{n\cdot r}\left[\underbrace{\Gamma'}_{r\cdot n}\underbrace{Z_{t-1}}_{n\cdot 1}\right] + \sum_{i=1}^{p-1}\beta_i\Delta Z_{t-i} + \varepsilon_t \tag{7}$$

where β_i for $i = 1, \ldots, p-1$ are the short run effects and ε_t is a zero mean i.i.d. shock. In the square brackets we focus on the cointegrated space: vector $\Gamma' Z_{t-1}$ contains the long-run cointegrating relations among variables, while matrix λ is the speed of adjustment to the equilibrium; $\sum_{i=1}^{p-1}\beta_i\Delta Z_{t-i}$ takes into account the short-run effects; furthermore, exogenous variables and dummy controls can be added. Application of the model represented in equation (7) requires to test empirically the presence of one or more cointegration vectors (equal to r) among variables of vector Z_t.

In order to reconcile the theoretical model represented in equation (6) with the updated version of the econometric formulation (7), Z_t must contain the one period update of the following endogenous variables: the logarithm of the real per capita consumption, the logarithm of the real disposable per capita income, the logarithm of the ratio between permanent and current real per capita incomes, and ratios between the one-period lagged real per capita assets and the current level of the real disposable per capita income as well as the logarithm of the ratio between real house prices and the real per capita income, the general credit conditions index and the real mortgage rate. Someone could argue that the use of y_t – rather than y_{t-1} – to deflate assets may generate a possible endogeneity issue. To avoid this case, in Section 5 we run a robustness check in which we use fiited values for y_t obtained from a parsimonious instrumenting equation which uses y_{t-1} among regressors.[15]

Since cointegration tests require that time series must be non-stationary, we first implement univariate unit root tests to assess the presence of integration. Based on the statistical significance of the intercept and linear time trend, we specify the ADF (Dickey and Fuller, 1979) and KPSS (Kwiatkowski, Phillips, Schmidt, and Shin, 1992) test regressions. All variables appear to be integrated of order one.[16] We initially estimate an unrestricted VAR(p). To select the optimal VAR lag length, we employ several information criteria for different specifications of our model (including or not $GCCI_t$). Following the SC and HQ tests we choose $p = 2$: this choice allows us to maintain a parsimonious model in terms of parameters as well as to be consistent with the economic intuition. Then we run the Johansen (1991, 1995) cointegrations analysis in order to verify if among integrated time series cointegration relationships arise. From the maximum eigenvalue cointegration rank test we obtain a

[15] See also the appendix A.4.

[16] Since the ADF test is often criticised for its low power in rejecting the null hypothesis, especially when the sample size is small, we also implement stationarity tests such as the KPSS test. Results of the unit root tests are available by the authors upon request.

unique cointegrating vector ($r = 1$) which links the log ratio of consumption to income, real interest rate, the three asset-to-income ratios and $GCCI_t$.[17] Previous information suggests that our consumption function can be estimated by using a VECM(1).[18]

4 Main results

Tables 1 and 2 show our main findings about the long-run and short-run effects of wealth on consumption. We follow a step by step approach, adding progressively new variables in the regressions.

We begin by estimating a simple version of the model described by equation (6), where we do not distinguish between financial and real assets ($\gamma = \gamma_1 = \gamma_2 = \gamma_3$) and do not include interest rates ($\alpha_1 = 0$) neither the general credit condition index ($\alpha_{0c} = 0$ and $\beta_4 = 0$). Column [1] shows that both coefficients of the estimated permanent income to the current one (α_2) and the total net worth (γ) have the expected positive signs and are statistically significant. The estimated long-run marginal propensity to consume (*mpc*) out of total net worth is approximately 2.2 per cent.

Column [2] adds the real mortgage interest rate ($\alpha_1 \neq 0$) which is initially not statistically significant. Anyway, the introduction of r_t does not change the sign and the statistical significance of the other coefficients.

In column [3] we split up total net worth into the financial component and the housing one. This model can be represented by equation (6) when $\alpha_{0c} = \beta_4 = 0$, and $\gamma_1 = \gamma_2$. Our results show that the *mpc* of the financial assets is greater than that of housing one: a unit increase in financial and housing wealth out of the current income would be associated with a yearly increase of 5 and 0.2 per cent in total households' consumer spending, respectively. The coefficient of the interest rate becomes negative as expected.

Column [4] relaxes the textbook model, by allowing the ratio to income of net liquid assets (liquid assets minus households' debt) to have a different coefficient from illiquid assets and housing wealth. With respect to the previous model we remove the constraint $\gamma_1 = \gamma_2$. This is a novelty for Italy. The estimated *mpc* out of net liquid assets (γ_1) is 7.6 per cent and greater than the *mpc* of illiquid financial wealth ($\gamma_2 = 3.4$ per cent) and of the *mpc* total net worth (γ) highlighted in column [1]. Coefficients related to net liquid and illiquid financial assets turn out statistically significant whilst the effect of the housing wealth becomes significant ($\gamma_3 = 1.3$ per cent).

In column [5] we introduce the general credit conditions index which should shift the intercept of the model. In this version of the model, we do not find the expected shift due to the introduction of credit conditions: the associated parameter (α_{0c}) is statistically insignificant. However, we confirm all the previous findings: the coefficient related to the permanent income to the current one is positive and statistically significant; the *mpc* out

[17] When we run the trace test some evidence in favor of $r > 1$. Nevertheless, given our economic a priori based on our theoretical setup we set the rank r=1. Moreover, by using fitted values obtained by the instrumenting equation described in the appendix A.4, and by taking into account that $\Delta \ln y_t$ is exogenous with respect the cointegration space, we find only one cointegration vector, both by using the trace and maximum eigenvalue cointegration rank tests.

[18] Results are available upon request.

of net liquid assets (7.5 per cent) is greater than the *mpc* of illiquid financial (3.3 per cent) and housing (1.7 per cent) ones. The real mortgage interest rate has a negative influence on consumption whereas the estimated permanent income to the current one (α_2) accounts for almost 45% percent of the variation of per capita real consumption (36% without $GCCI$).

Column [6] reports estimations by adding the log ratio between house prices and per capita income to proxy the saving need for a downpayment.[19] To appreciate this effect we initially exclude the impact of $GCCI$. On the one hand, since in Italy houses are very often acquired through inheritance, when housing prices go up we can expect positive effects on consumption from housing assets. On the other hand, given the relatively undeveloped mortgage market, people who do not have their own house must save to buy it: for them we could expect a negative sign from the saving for a downpayment. Overall, with respect previous regressions, we find a larger statistically significant housing wealth effect (γ_3) which is offset by the negative significant downpayment one (γ_4=-0.107).

Column [7] summarizes results of previous regression when we reintroduce the general credit condition index: on the one hand the R^2 and speed of adjustment increase; on the other hand the coefficient γ_4 becomes slightly more negative (γ_4=-0.117). Interestingly, in this specification of the model we also find that α_{0c} is positive and statistically significant: hence, credit conditions raise the intercept α_0, determining a higher level of ln c_t/y_t. As argued by Hendry and Muellbauer (2018), the omission of shifts in credit conditions have been considered a misspecification of the model.

Column [8] provides estimates for the specification in which we separate gross liquid assets ($\gamma_{1,gla}$) and debts ($\gamma_{1,loans}$). All our previous results hold and we do not observe large improvements in the contribution of liquid assets: the *mpc* out of gross liquid wealth is statistically significant around 7.6 per cent whereas credit debts show a negative statistical significant coefficient.

In presence of better credit conditions, the coefficient measuring the sensitivity of downpayment requirements to house prices relative to income, γ_4, should become less negative; nevertheless this effect is more than offset by a possible increase of the share of households subject to the downpayment constraint. Column [9] reports estimates of the previous model ending the sample period before the start of the global financial crisis: in this case, despite the more favorable credit conditions, the coefficient on the log house price to income ratio turns out more negative (γ_4=-0.232).[20]

Short-run and exogenous variables' effects highlight that public spending has a negative and significant influence on consumption while the effect of the credit condition index is positive.[21] On the other hand, the variation of the current income has very low significant values. Then, in column [10] we estimate again regression [8] by imposing $\beta_1 = 0$. In this model, with respect column [8], the coefficient used to proxy downpayment constraints is

[19]Other measures of the downpayment constraint relate to the degree of liquidity of housing assets: Iacoviello and Neri (2010) argue that the increase in liquidity of housing is captured in the higher LtV; Garrida and Hedlund (2020) estimate a micro-founded model suggesting that housing liquidity and collateral effects, transmitted to consumption via balance sheets, have a central role in explaining aggregate dynamics.

[20]The introduction of the adult proportion in pre-retirement age group (45-60 years) and/or the share of the working age population are never statistically significant.

[21]Among the short-run effects we also tested for the presence of the change in nominal and real interest rates that are never statistical significant.

γ_4=-0.159. The magnitude of the latter coefficient is larger than the values experienced by other European countries as France (γ_4 ranges between -0.062 and -0.081 based on the sample period; see Chauvin and Muellbauer, 2018) and Germany (γ_4 ranges between -0.057 and -0.070 based on the sample period and the definition of illiquid financial assets; see Muellbauer, Geiger, and Rupprecht, 2016). It is useful to note as in this specification the R^2 (residual standard errors) slightly increases (decreases). Finally, the speed of convergence to the long-run equilibrium (λ) is equal to 0.32 and it is statistically significant in all models.

In all regressions from column [1] to column [10] we include three dummies inside the cointegration space to control for (i) the 1992 recession, (ii) the 1995 pension reform and (iii) the debt sovereign crisis in Italy.[22] Only the dummy for 1992 is always statistically significant in all specifications, consistently with the sharp decrease of the total disposable income highlighted in figure 3 and the beginning of the associated social security reform.[23] Moreover, in the benchmark regression [10], du_{95} imlplies a significant effect on consumption due the 1995 pension reform and the change in accounting rules whereas the dummy related the sovereign debt crisis increases its relevance but the coefficient remains not statistically significant.

Regression [10] is the most complete and therefore we assume it as the benchmark for the robustness checks in Section 5.

[22] Outside the cointegrating relation, we also test the significance of additional dummy variables to control for other specific events, e.g. the official join of Italy to the single currency area in 1998, the profit tax increases of funds and insurances relative to the other financial instruments in 2012 and 2014 or the start of the 2007 global financial crisis: no statistically significant effects arise.

[23] Rossi and Visco (1995) estimate that the impact of the 1992 social security reform in the long run leads the private saving ratio to rise by 3 percentage points.

Table 1: **Italian Consumption Function Estimates, 1975-2017 (1).**

Dep. Var. = $\Delta \ln c_t$	Symbol	(1)	(2)	(3)	(4)	(5)
			Long-run effects			
Speed of adjustment	λ	0.094	0.090	0.234	0.256	0.261
		(0.017)	(0.016)	(0.031)	(0.032)	(0.033)
		[5.601]	[5.810]	[7.465]	[7.917]	[7.837]
$(\ln y_t - \ln c_{t-1})$	-	1,000	1,000	1,000	1,000	1,000
Constant	α_0	-0.064	-0.107	-0.058	-0.142	-0.119
		(0.094)	(0.104)	(0.038)	(0.049)	(0.059)
		[-0.675]	[-1.025]	[-1.537]	[-2.900]	[-2.016]
$GCCI_t$	α_{0c}					0.0542
						(0.074)
						[0.730]
r_t	α_1		0.004	-0.002	-0.003	-0.003
			(0.004)	(0.001)	(0.001)	(0.001)
			[1.215]	[-1.371]	[-2.007]	[-2.292]
$\ln y_t^P/y_t$	α_2	0.492	0.434	0.318	0.361	0.452
		(0.277)	(0.284)	(0.104)	(0.096)	(0.129)
		[1.779]	[1.528]	[3.053]	[3.764]	[3.505]
A_{t-1}/y_t	$\gamma = \gamma_1 = \gamma_2 = \gamma_3$	0.022	0.025			
		(0.010)	(0.011)			
		[2.216]	[2.354]			
TFA_{t-1}/y_t	$\gamma_1 = \gamma_2$			0.051		
				(0.006)		
				[8.549]		
NLA_{t-1}/y_t	γ_1				0.076	0.075
					(0.012)	(0.012)
					[6.222]	[6.462]
IFA_{t-1}/y_t	γ_2				0.034	0.033
					(0.009)	(0.009)
					[3.762]	[3.714]
HA_{t-1}/y_t	γ_3			0.002	0.013	0.017
				(0.005)	(0.006)	(0.007)
				[0.465]	[2.088]	[2.463]

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Dep. Var. = $\Delta \ln c_t$	Symbol	(1)	(2)	(3)	(4)	(5)
d_{92}	τ_1	-0.071	-0.109	-0.040	-0.055	-0.037
		(0.039)	(0.042)	(0.015)	(0.016)	(0.015)
		[-1.813]	[-2.594]	[-2.621]	[-3.501]	[-2.423]
d_{95}	τ_2	-0.037	-0.029	-0.015	0.001	-0.008
		(0.030)	(0.030)	(0.012)	(0.013)	(0.014)
		[-1.261]	[-0.961]	[-1.235]	[0.082]	[-0.554]
d_{11}	τ_3	-0.091	-0.080	-0.005	-0.010	-0.009
		(0.032)	(0.033)	(0.013)	(0.012)	(0.011)
		[-2.814]	[-2.405]	[-0.365]	[-0.821]	[-0.751]
Short-run effects						
$\Delta \ln y_t$	β_1	0.121	0.125	0.027	0.029	0.014
		(0.048)	(0.048)	(0.050)	(0.049)	(0.047)
		[2.538]	[2.627]	[0.553]	[0.589]	[0.303]
$\Delta \ln c_{t-4}$	β_2	0.309	0.323	0.309	0.318	0.328
		(0.065)	(0.064)	(0.061)	(0.060)	(0.058)
		[4.760]	[5.007]	[5.059]	[5.306]	[5.687]
$\Delta \ln c_t^P$	β_3	-0.053	-0.051	-0.051	-0.045	-0.043
		(0.019)	(0.019)	(0.018)	(0.018)	(0.017)
		[-2.775]	[-2.656]	[-2.785]	[-2.495]	[-2.517]
$\Delta_4 GCCI_{t-1}$	β_4					0.070
						(0.023)
						[3.058]
Adj. R^2		35.136	34.937	41.177	42.667	47.121
Eq. standard errors		0.00832	0.00833	0.00792	0.00782	0.00751
Eq. log likelihood		561.44	561.19	569.55	571.68	578.91
Eq. Schwarz Criterion		-6.64	-6.64	-6.74	-6.76	-6.82

Notes: standard errors in parenthesis; t-stats in square brackets.

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Table 2: **Italian Consumption Function Estimates, 1975-2017 (2).**

Dep. Var. = $\Delta \ln c_t$	Symbol	(6)	(7)	(8)	(9)	(10)
		Long-run effects				
Speed of adjustment	λ	0.289	0.331	0.316	0.338	0.318
		(0.034)	(0.038)	(0.035)	(0.043)	(0.030)
		[8.445]	[8.660]	[9.105]	[7.893]	[10.552]
$(\ln y_t - \ln c_{t-1})$	-	1,000	1,000	1,000	1,000	1,000
Constant	α_0	1.202	1.400	1.901	2.662	1.894
		(0.471)	(0.416)	(0.534)	(-1.012)	(0.524)
		[2.549]	[3.362]	[3.558]	[2.631]	[3.616]
$GCCI_t$	α_{0c}		0.168	0.117	0.011	0.118
			(0.058)	(0.065)	(0.117)	(0.064)
			[2.905]	[1.808]	[0.097]	[1.845]
r_t	α_1	-0.003	-0.003	-0.004	-0.003	-0.004
		(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
		[-2.741]	[-3.749]	[-3.992]	[-2.482]	[-4.048]
$\ln y_t^P / y_t$	α_2	0.215	0.408	0.282	0.358	0.284
		(0.093)	(0.102)	(0.129)	(0.172)	(0.125)
		[2.306]	[4.016]	[2.183]	[2.085]	[2.271]
NLA_{t-1}/y_t	γ_1	0.077	0.074			
		(0.011)	(0.009)			
		[7.238]	[8.226]			
GLA_{t-1}/y_t	$\gamma_{1,gla}$			0.076	0.073	0.076
				(0.009)	(0.010)	(0.009)
				[8.055]	[6.990]	[8.197]
$LOANS_{t-1}/y_t$	$\gamma_{1,loans}$			-0.137	-0.107	-0.136
				(0.056)	(0.080)	(0.054)
				[-2.447]	[-1.328]	[-2.549]
IFA_{t-1}/y_t	γ_2	0.016	0.011	0.017	0.016	0.017
		(0.010)	(0.009)	(0.009)	(0.020)	(0.009)
		[1.575]	[1.291]	[1.811]	[0.832]	[1.832]
HA_{t-1}/y_t	γ_3	0.025	0.037	0.044	0.062	0.044
		(0.007)	(0.007)	(0.012)	(0.021)	(0.011)
		[3.551]	[5.356]	[3.805]	[2.940]	[3.929]
$\ln HP_{t-1}/y_t$	γ_4	-0.107	-0.117	-0.160	-0.232	-0.160
		(0.038)	(0.033)	(0.044)	(0.086)	(0.043)
		[-2.838]	[-3.585]	[-3.638]	[-2.708]	[-3.706]

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Dep. Var. = $\Delta\ln c_t$	Symbol	(6)	(7)	(8)	(9)	(10)
d_{92}	τ_1	-0.043	-0.020	-0.022	-0.032	-0.022
		(0.014)	(0.012)	(0.012)	(0.015)	(0.012)
		[-3.187]	[-1.655]	[-1.790]	[-2.167]	[-1.798]
d_{95}	τ_2	-0.004	-0.022	-0.023	-0.010	-0.023
		(0.011)	(0.011)	(0.011)	(0.014)	(0.011)
		[-0.296]	[-2.024]	[-2.025]	[-0.697]	[-2.053]
d_{11}	τ_3	-0.012	-0.010	-0.010		-0.010
		(0.010)	(0.009)	(0.009)		(0.009)
		[-1.120]	[-1.105]	[-1.034]		[-1.033]
Short-run effects						
$\Delta\ln y_t$	β_1	0.005	-0.020	-0.006	0.001	
		(0.048)	(0.047)	(0.046)	(0.055)	
		[0.111]	[-0.422]	[-0.137]	[0.012]	
$\Delta\ln c_{t-4}$	β_2	0.300	0.326	0.287	0.307	0.286
		(0.059)	(0.056)	(0.056)	(0.067)	(0.056)
		[5.101]	[5.860]	[5.093]	[4.574]	[5.147]
$\Delta\ln c_t^P$	β_3	-0.039	-0.035	-0.041	-0.037	-0.042
		(0.017)	(0.017)	(0.017)	(0.020)	(0.016)
		[-2.232]	[-2.126]	[-2.489]	[-1.878]	[-2.529]
$\Delta_4 GCCI_{t-1}$	β_4		0.061	0.048	0.058	0.048
			(0.022)	(0.023)	(0.028)	(0.023)
			[2.740]	[2.104]	[2.058]	[2.124]
Adj. R^2		45.526	50.499	50.939	42.983	51.241
Eq. standard errors		0.00763	0.00727	0.00724	0.00799	0.00722
Eq. log likelihood		575.93	584.39	585.13	425.48	585.13
Eq. Schwarz Criterion		-6.82	-6.89	-6.89	-6.67	-6,93

Notes: standard errors in parenthesis; t-stats in square brackets.

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5 Robustness Analysis

In this Section we run some robustness checks, using a different (i) horizons for the computation of the permanent income and (ii) a wider set of interest rates. For comparison, in table 3 we report in the first column regression results [10] taken from table 2.

As discussed in Section 3.1 a strand of literature suggests to use a shorter horizon to calculate the permanent income ($y^P_{k=12}$): by setting 12 quarters, i.e. 3 years, we find $\alpha_2 =$ around 0.37. With reference to wealth effects, both gross financial and housing assets remain statistically significant whereas the statistical significance of the illiquid financial wealth slightly lowers (table 3). We can observe similar results when we consider a 5 years horizon ($y^P_{k=20}$). Indeed, by comparing models with different definitions of the permanent income, the benchmark specification highlights the largest adjusted R^2 and likelihood and the lowest value of the Bayesian criteria.

Then we analyse the effects of interest rates on bonds and deposits on consumption. On the one hand, as already seen, mortgage rates influence consumer spending; on the other hand, one may envisage that higher returns on deposits and bonds might increase saving and depress consumption. Columns [r_{dep}] and [r_{bond}] show the estimates in the case of the real deposits rate and the real benchmark bond rate of the General Government with an average duration equals to 6.5 years rather than the real mortgage one. We get negative and significant effects of the interest rates on deposits and bonds on consumption.[24] Nevertheless, by comparing models with different interest rates, the specification that uses the mortgage interest rate highlights the best fit.[25] All other previous results are confirmed.[26]

As explained in Section 3.2 our assets are deflated by y_t and in the cointegration space appears $\ln y_t$ rather than $\ln y_{t-1}$. Hence, to avoid any possible endogenous issue, we check our results by instrumenting the current income by using fitted values for $\ln y_t$ (and its exponential, y_t, to deflate assets) from a parsimonious instrumenting equation which uses y_{t-1} among regressors: estimates of column [$\hat{y}$] confirm all our long-run results. Among the short-run effects the variation of the real per capita income (by using the fitted values) bceomes statistical significant ($\beta_1 = 0.19$).

[24]Elmendorf (1996) shows as economists' understanding of the response of household saving and consumption to changes in interest rates is quite limited and it is not possible to provide a precise estimate of the interest rate elasticity of saving with any confidence. Moreover, Cloyne, Ferreira, and Surico (2016) highlights that households heterogeneity matters for the monetary policy transmission: after an interest rate change, households with mortgages to pay adjust their consumption significantly whereas renters and outright home-owners are far less sensitive.

[25]The preference for the mortgage rate holds also by using different definitions of permanent income.

[26]A stability analysis for different subperiods shows that as the number of observations increases the relevance of financial (housing) assets increases (decreases). Moreover, preliminary results on a boom/bust analysis provide evidence that wealth effects during booms are larger than those during bust/stagnation periods.

Table 3: **Italian Consumption Function Estimates, 1975-2017.**

Dep. Var. = $\Delta \ln c_t$	Symbol	(10), Tab. 2	$y^P_{k=12}$	$y^P_{k=20}$	r_{dep}	r_{bond}	$\hat{y}$
		Long-run effects					
Speed of adjustment	λ	0.318	0.307	0.312	0.326	0.299	0.316
		(0.030)	(0.030)	(0.030)	(0.031)	(0.029)	(0.033)
		[10.552]	[10.326]	[10.429]	[10.520]	[10.288]	[9.630]
$(\ln y_t - \ln c_{t-1})$	-	1,000	1,000	1,000	1,000	1,000	1,000
Constant	α_0	1.894	2.212	2.048	2.006	1.725	1.541
		(0.524)	(0.510)	(0.514)	(0.527)	(0.559)	(0.490)
		[3.616]	[4.341]	[3.981]	[3.807]	[3.086]	[3.147]
$GCCI_t$	α_{0c}	0.118	0.088	0.103	0.151	0.146	0.157
		(0.064)	(0.062)	(0.062)	(0.063)	(0.070)	(0.056)
		[1.845]	[1.422]	[1.674]	[2.407]	[2.081]	[2.822]
r_t	α_1	-0.004	-0.004	-0.004	-0.004	-0.003	-0.004
		(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
		[-4.048]	[-3.337]	[-3.606]	[-4.224]	[-2.930]	[-3.725]
$\ln y^P_t / y_t$	α_2	0.284	0.367	0.312	0.391	0.337	0.402
		(0.125)	(0.211)	(0.149)	(0.121)	(0.134)	(0.086)
		[2.271]	[1.740]	[2.093]	[3.227]	[2.516]	[3.723]
GLA_{t-1}/y_t	$\gamma_{1,gla}$	0.076	0.077	0.076	0.083	0.080	0.076
		(0.009)	(0.010)	(0.009)	(0.009)	(0.011)	(0.010)
		[8.197]	[8.005]	[8.067]	[8.701]	[7.644]	[7.992]
$LOANS_{t-1}/y_t$	$\gamma_{1,loans}$	-0.136	-0.174	-0.155	-0.105	-0.106	-0.078
		(0.054)	(0.050)	(0.052)	(0.051)	(0.056)	(0.050)
		[-2.549]	[-3.447]	[-3.014]	[-2.052]	[-1.877]	[-1.559]
IFA_{t-1}/y_t	γ_2	0.017	0.015	0.015	0.020	0.018	0.014
		(0.009)	(0.010)	(0.010)	(0.009)	(0.010)	(0.010)
		[1.832]	[1.476]	[1.494]	[2.182]	[1.778]	[1.485]
HA_{t-1}/y_t	γ_3	0.044	0.052	0.049	0.044	0.041	0.034
		(0.011)	(0.011)	(0.011)	(0.011)	(0.012)	(0.011)
		[3.929]	[4.666]	[4.424]	[3.935]	[3.431]	[3.026]
$\ln HP_{t-1}/y_t$	γ_4	-0.160	-0.188	-0.173	-0.170	-0.147	-0.128
		(0.043)	(0.041)	(0.042)	(0.043)	(0.046)	(0.040)
		[-3.706]	[-4.537]	[-4.129]	[-3.920]	[-3.188]	[-3.212]

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Dep. Var. = $\Delta\ln c_t$	Symbol	(10), Tab. 2	$y^P_{k=12}$	$y^P_{k=20}$	r_{dep}	r_{bond}	$\hat{y}$
d_{92}	τ_1	-0.022	-0.027	-0.024	-0.025	-0.030	-0.028
		(0.012)	(0.013)	(0.012)	(0.012)	(0.013)	(0.012)
		[-1.798]	[-2.109]	[-1.973]	[-2.141]	[-2.329]	[-2.264]
d_{95}	τ_2	-0.023	-0.022	-0.022	-0.017	-0.016	-0.017
		(0.011)	(0.011)	(0.011)	(0.011)	(0.012)	(0.011)
		[-2.053]	[-1.913]	[-1.965]	[-1.513]	[-1.285]	[-1.496]
d_{11}	τ_3	-0.010	-0.008	-0.009	-0.014	-0.010	-0.019
		(0.009)	(0.010)	(0.009)	(0.009)	(0.010)	(0.009)
		[-1.033]	[-0.847]	[-0.961]	[-1.512]	[-1.004]	[-2.106]
		Short-run effects					
$\Delta\ln y_t$	β_1						0.192
							(0.048)
							[4.017]
$\Delta\ln c_{t-4}$	β_2	0.286	0.280	0.282	0.264	0.291	0.273
		(0.056)	(0.056)	(0.056)	(0.057)	(0.056)	(0.057)
		[5.147]	[4.958]	[5.045]	[4.648]	[5.194]	[4.799]
$\Delta\ln c^P_t$	β_3	-0.042	-0.042	-0.042	-0.044	-0.042	-0.043
		(0.016)	(0.017)	(0.016)	(0.017)	(0.017)	(0.017)
		[-2.529]	[-2.510]	[-2.502]	[-2.635]	[-2.561]	[-2.556]
$\Delta_4 GCCI_{t-1}$	β_4	0.048	0.046	0.047	0.041	0.044	0.0423
		(0.023)	(0.023)	(0.023)	(0.023)	(0.023)	(0.023)
		[2.124]	[2.033]	[2.070]	[1.792]	[1.903]	[1.831]
Adj. R^2		51.241	50.556	51.079	50.842	50.491	50.066
Eq. standard errors		0.00721	0.00726	0.00723	0.00724	0.00727	0.00730
Eq. log likelihood		585.13	583.97	584.85	584.45	583.86	583.66
Eq. Schwarz Criterion		-6.93	-6.91	-6.92	-6.92	-6.91	-6.88

Notes: standard errors in parenthesis; t-stats in square brackets. In the equilibrium correction component of column [$\hat{y}$] we use the instrumented income depending on the previous period income (see equation A.2) instead of ($\ln y_t - \ln c_{t-1}$).

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6 Conclusions

The recent financial and economic crises and their impact on households' wealth have spurred new interest on the relationship between consumption and wealth. This paper studies the long-run effects of housing and financial wealth on household consumption in Italy. Our main contribution is to estimate a VEC model with a disaggregation of financial wealth into a net liquid component (deposits, bonds, quoted shares, and mutual funds net of total debts) and an illiquid one (unquoted shares and insurance technical reserves). Using quarterly data, our analysis covers the time span from 1975 to 2017.

We find a positive and statistically significant effect of financial and housing assets on consumption. The influence of net liquid wealth – about 7.5 per cent – is greater than that of illiquid assets – about 3 per cent – whereas housing effect is positive but smaller (less than 2 per cent). Our results are broadly consistent with Bassanetti and Zollino (2010), who find that the marginal propensity to consume out of housing and total non housing wealth is in the range of, respectively, 1.5-2 and 4-6 per cent. Our results show that permanent income has a positive impact on consumption while the effect of the real interest rate is negative; an index of credit constraints implies a shift of the consumption-income ratio in the long-run.

The econometric results are robust to the use of different methods to estimate permanent income and to the inclusion in the regressions of control variables to take into account pension reforms, recessions, public spending, downpayment constraints and interest rates on deposits and bonds. Permanent shocks and macroeconomic conditions matter: for instance the currency crises of 1992 and 1995 pension reform implied a contraction of consumption.

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Appendix

A.1 Raw data descriptive statistics

In this Section we show the main descriptive statistics of the variables used in the analysis.

Time series (obs. 172 quarters)	Mean	Median	Max	Min	S.D.
Consumption (billions)	147.1	152.5	263.3	11.3	84.0
Consumer Price Index (base year 2010)	0.67	0.75	1.09	0.10	0.31
Disposable income (billions)	171.9	191.0	293.1	13.9	91.6
Public spending (billions)	13.7	12.2	26.3	0.9	8.3
Public spending deflator	0.7	0.8	1.0	0.2	0.3
Total population (thousands)	57,512.9	57,030.0	60,758.0	54,659.0	1,713.4
Annual Gross Domestic Product (billions)	1,364.0	1,424.9	1,690.2	8,503.7	247.2
Housing and land (billions)	2,857.8	2.510.4	5,946.8	192.5	1,970.3
Deposits and currency (billions)	613.8	652.1	1,360.9	38.0	414.4
Bonds (billions)	398.5	431.5	802.1	7.8	261.1
Insurances and pension schemes (billions)	322.1	194.4	995.1	3.8	307.4
Quoted shares (billions)	63.8	69.3	208.8	0.06	48.6
Unquoted shares (billions)	438.7	297.4	1,221.4	0.4	341.9
Mutual funds (billions)	199.7	121.6	532.2	0.2	190.3
Loans (billions)	281.8	177.3	717.3	5.4	267.5
Ratio between used and granted credit lines	0.673	-.667	0.830	0.528	0.070
Mortgage interest rate (percent)	9.97	10.01	23.70	1.97	6.26
Deposit interest rate (percent)	5.71	5.93	15.31	0.30	4.74
Bond interest rate (percent)	8.67	8.97	21.21	0.70	5.40
Nominal house prices (euros/m^2)	885.3	910.4	91.5	1,632.4	509.8

Table A.1: **Descriptive Statistics: 1975-2017.**

Description	Source
National consumption expenditure, households, current prices	Istat, National Accounts
Consumer Price Index	Istat
Total disposable income, households, current prices	Istat, National Accounts and Brandolini, Gambacorta, and Rosolia (2018)
Public spending, current prices	Istat, National Accounts
Public spending deflator	Istat
Total population	Istat
Gross Domestic Product, current prices	Istat, National Accounts
Housing and land, households, current prices	Istat and Cannari, D'Alessio, and Vecchi (2017)
Deposits and currency, households, current prices	Bank of Italy, Financial Accounts and Bonci and Coletta (2008)
Bonds, households, current prices	Bank of Italy, Financial Accounts and Bonci and Coletta (2008)
Insurances and pension schemes, households, current prices	Bank of Italy, Financial Accounts and Bonci and Coletta (2008)
Quoted shares, households, current prices	Bank of Italy, Financial Accounts and Bonci and Coletta (2008)
Unquoted shares, households, current prices	Bank of Italy, Financial Accounts and Bonci and Coletta (2008)
Mutual funds, households, current prices	Bank of Italy, Financial Accounts and Bonci and Coletta (2008)
Ratio between used and granted credit lines, total economy	Bank of Italy
Mortgage interest rate, households, percentage value	Bank of Italy
Deposit interest rate, households, percentage value	Bank of Italy
Bond interest rate, percentage value	General Government
Nominal house prices	Bank of Italy, Bank for International Settlements, and Revenue Agency

Table A.2: **Data: Source and Treatment.**

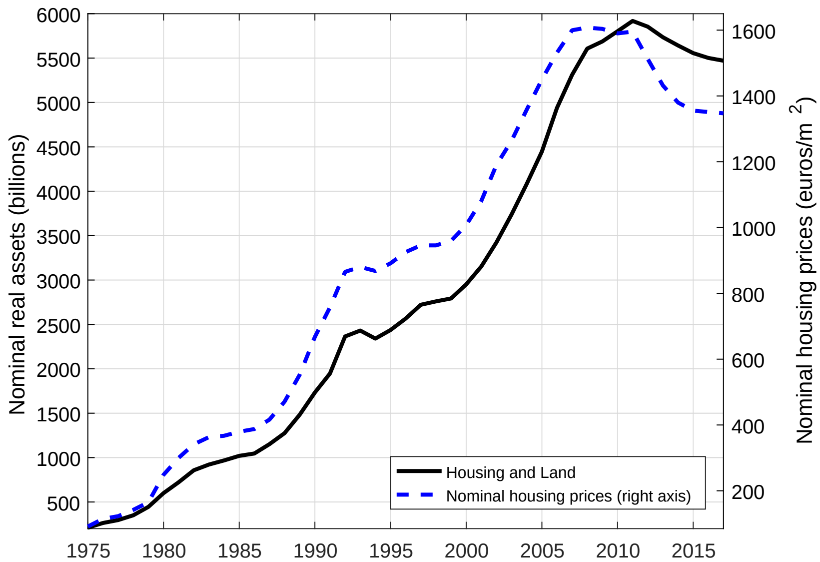

Figure A.1: **Household real assets.**

Source: Cannari, D'Alessio, and Vecchi (2017), Istat, Bank of Italy, BIS and Revenue Agency.

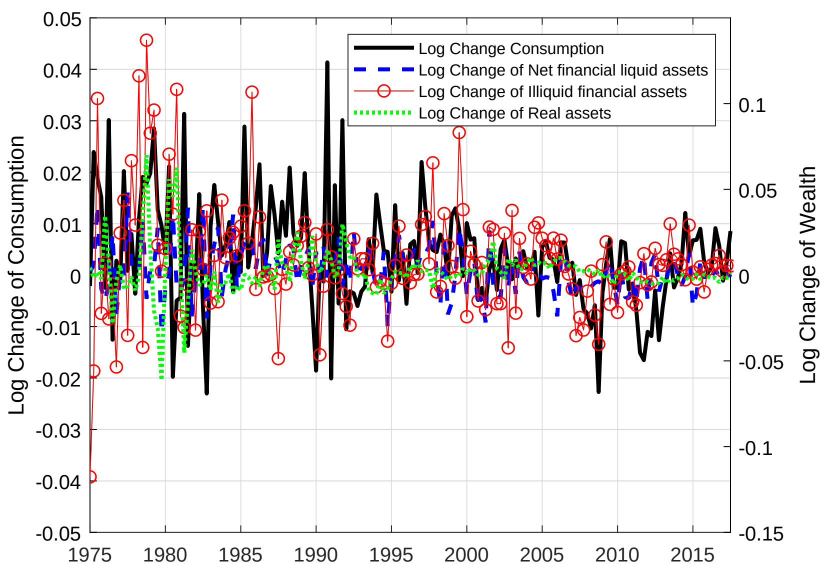

Figure A.2: **Log change of consumption and wealth.**

A.2 The life-cycle permanent income consumption function

Starting from equation (1) we have:

$$c_t = \gamma^* A_{t-1} + \omega^* y_t^P$$

Then:

$$\frac{c_t}{y_t} = \gamma^* \frac{A_{t-1}}{y_t} + \omega^* \frac{y_t^P}{y_t}$$

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

$$\frac{c_t}{y_t} = \omega^* \left[\frac{\gamma^*}{\omega^*} \frac{A_{t-1}}{y_t} + 1 + \frac{y_t^P - y_t}{y_t} \right]$$

As noticed by Aron et al. (2012), the right-hand side of previous equation has the form $1 + x$, where x is usually a fairly small number. Then, when we use the logs we can use the $\ln(1 + x) \approx x$ where $x = \left[\frac{\gamma^*}{\omega^*} \frac{A_{t-1}}{y_t} + \frac{y_t^P - y_t}{y_t} \right]$.

$$\ln \frac{c_t}{y_t} = \ln \omega^* + \ln \left[1 + \frac{\gamma^*}{\omega^*} \frac{A_{t-1}}{y_t} + \frac{y_t^P - y_t}{y_t} \right]$$

$$\ln c_t - \ln y_t = \ln \omega^* + \frac{\gamma^*}{\omega^*} \frac{A_{t-1}}{y_t} + \frac{y_t^P - y_t}{y_t}$$

By using the further approximation $\frac{y_t^P - y_t}{y_t} \approx \ln \frac{y_t^P}{y_t}$ we obtain

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

which is equation (2) and where $\gamma = \gamma^*/\omega^*$ and $\alpha_0 = \ln \omega^*$. Then:

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

$$\Delta \ln c_t = \alpha_0 + (\ln y_t - \ln c_{t-1}) + \gamma \frac{A_{t-1}}{y_t} + \ln \left(\frac{y_t^P}{y_t} \right)$$

Finally, by introducing the role of the interest rate and by allowing for the measurement of effect the permanent income (α_2), the change of log consumption can be approximated as:

$$\Delta \ln c_t \approx \lambda \left[\alpha_0 + \alpha_1 r_t + (\ln y_t - \ln c_{t-1}) + \alpha_2 \ln \left(\frac{y_t^P}{y_t} \right) + \gamma \frac{A_{t-1}}{y_t} \right] + \varepsilon_t$$

which is equation (5) of the text. The long-run equilibrium solution of previous equation is:

$$\ln \left(\frac{c_{t-1}}{y_t} \right) = \alpha_0 + \alpha_1 r_t + \alpha_2 \ln \left(\frac{y_t^P}{y_t} \right) + \gamma \frac{A_{t-1}}{y_t}$$

Updating by one period the previous condition we have:

$$\ln\left(\frac{c_t}{y_{t+1}}\right) = \alpha_0 + \alpha_1 r_{t+1} + \alpha_2 \ln\left(\frac{y^P_{t+1}}{y_{t+1}}\right) + \gamma\frac{A_t}{y_{t+1}}$$

The previous updated variables are those we use in the VECM model.

A.3 The permanent income forecasting model

In the following we report the forecasting model with households' learning (see Chauvin and Muellbauer, 2018). By departing from the permanent income described in Section 3.1 we add some economic and demographic variables:

$$\ln \frac{y_t^P}{y_t} \sim f\left(t, t_{learning}^{1993:q4}, t_{learning}^{2009:q4}, \ln y_{t-1}, (\Delta_4 y_t)_{t-1}, \ln \frac{labor_{t-1}}{pop_{t-1}}, exc_t, comp_{t-5}, r_{ma,t}\right) \quad \text{(A.1)}$$

where t represents the time trend of the regression, $\ln y$ is the log real per capita income, $\ln \dfrac{labor}{pop}$ is the log of the ratio between the labor force and the total population, exc is the real exchange rate, $comp$ is the good market competitiveness index and r_{ma} is the 4-quarters moving average of the real mortgage rate. Finally, $t_{learning}^{1993:q4}$ and $t_{learning}^{2009:q4}$ are dummies used to assume a gradual learning over 2 years from 1993:q4 and 2009:q4, respectively. In particular, $t_{learning}^{1993:q4}$ ($t_{learning}^{2009:q4}$) is computed by using a 2-year declining weighted moving average with quarterly discount factor of 0.95 of a dummy equal to 0 until 1993:q3 (2009:q3), and 1,2,3, and so on, from 1993:q4 (2009:q4). The regression is estimated by OLS methods (see table A.3). Durbin-Wartson implies the acceptation of the no residual autocorrelation hypothesis. The forecasted permanent income is shown in figure A.3. Figure A.4 reports the forecasted income growth. Finally, we use the fitted values of equation A.1 to estimate model represented in equation (6).

Dependent variable: Log real permanent income	Sample: 1975:q1-2017:q4			
	Estimate	Std. Error	t-value	Pr(>|t|)
Constant	-4.145	0.284	-14.61	0.000***
Linear trend	0.002	0.000	6.13	0.000***
Split Trend from 1993:Q4; discounted present value	-0.002	0.000	-5.87	0.000***
Split Trend from 2009:Q4; discounted present value	-0.001	0.000	-4.22	0.000***
Log real per capita income (t-1)	-0.780	0.046	-16.78	0.000***
4-quarter change in log real per capita income (t-1)	0.274	0.054	5.08	0.000***
Log labor force/total population (t-1)	0.544	0.099	5.46	0.000***
Real exchange rate	-0.364	0.005	-7.82	0.000***
Competitiveness Index (t-5)	0.081	0.019	4.15	0.000***
Borrowing real interest rate (4-qts moving average)	-0.004	0.001	-3.84	0.000***
Durbin Watson statistic	1.86			
Adj. R^2	0.92			
Residual standard error	0.013			

Notes: Statistical significance at the 10%, 5% and 1% levels are denoted by *, ** and ***.

Table A.3: **Estimates for the permanent income model.**

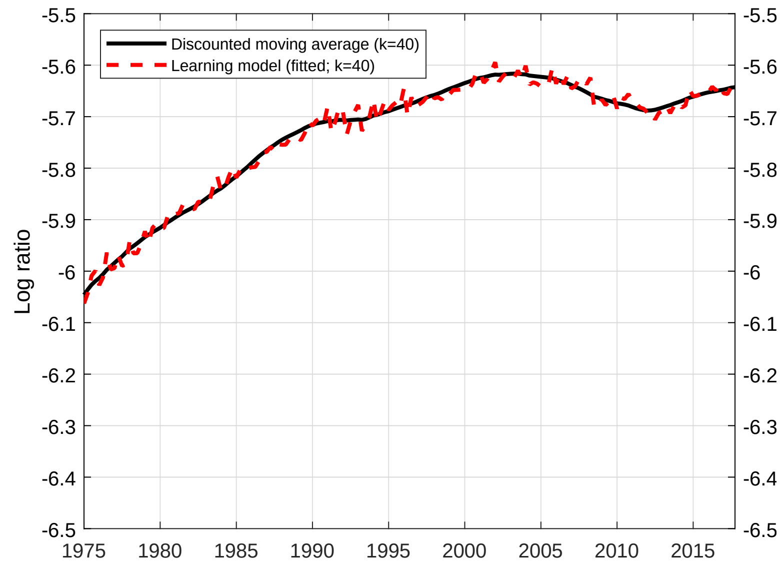

Figure A.3: **Permanent income.**

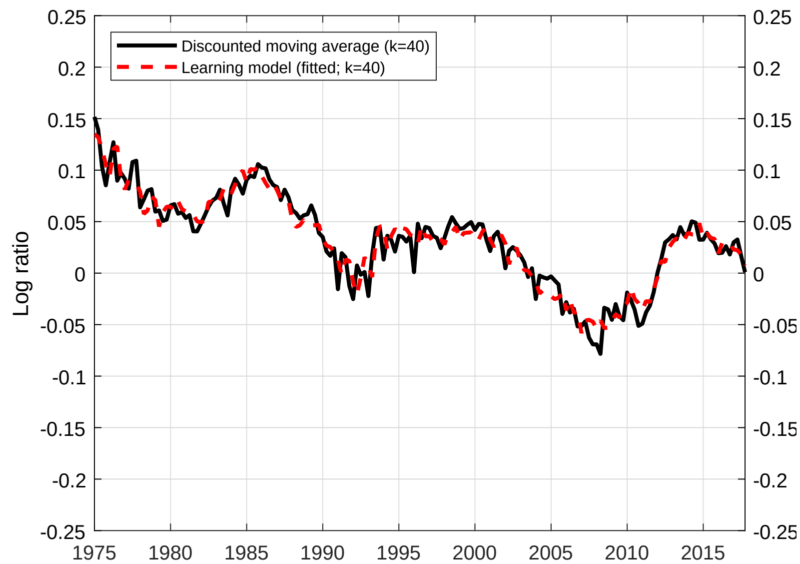

Figure A.4: **Forecasted permanent income growth.**

A.4 The instrumenting equation for the current income

In the following we report the instrumenting equation for the $\ln y_t$:

$$\ln y_t \sim f\left(\ln y_{t-1}, \ln c_{t-1}, \Delta_4 u_{ma,t-1}, \ln \frac{labor_{t-1}}{pop_{t-1}}\right) \tag{A.2}$$

where $\ln y$ and $\ln c$ are the log real per capita income the log real per capita consumption respectively, $\Delta_4 u_{ma,t-1}$ is the 4-quarters variation of the lagged 4-quarters moving average of the unemployment rate and $\ln \frac{labor}{pop}$ is the log of the ratio between the labor force and the total population. Estimates are reported in Table A.4:

Dependent variable: Log real per capita income	Sample: 1975:q1-2017:q4			
	Estimate	Std. Error	t-value	Pr(>|t|)
Constant	-0.403	0.074	-5.44	0.000***
Log real per capita income (t-1)	0.810	0.042	19.10	0.000***
Log real per capita consumption (t-1)	0.091	0.032	2.88	0.005***
4-quarter change in unemployment rate (t-1; 4-qts moving average)	-0.004	0.002	-2.48	0.014**
Log labor force/total population (t-1)	0.192	0.072	2.67	0.008***
Durbin Watson statistic	2.32			
Adj. R^2	0.99			
Residual standard error	0.014			

Notes: Statistical significance at the 10%, 5% and 1% levels are denoted by *, ** and ***.

Table A.4: **Instrumenting equation for ln y_t.**

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