THE EFFECTS OF HOUSING AND FINANCIAL WEALTH ON PERSONAL CONSUMPTION: AGGREGATE EVIDENCE FOR ITALIAN HOUSEHOLDS

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Abstract

In this paper we focus on Italian aggregate quarterly time series covering the sample period 1980-2006 and test the presence and size of wealth effects on consumption, taking separately account of financial assets and residential property on the basis of new stock measures. We find sound evidence in favour of a cointegrating relation, in which both wealth components display the expected, positive effect on households' consumption. In particular, our results point to a lower marginal propensity to consume out of housing than out of non housing net worth, with a respective size laying in the range of 1.5-2 and 4-6 cents. Following the estimate of a vector error correction model, we discuss the role played by transitory and permanent shocks in driving changes in the variables we consider. We have found that both consumption and wealth respond almost exclusively to permanent shocks, which are the sole determinants of the common stochastic trend in our theoretical set-up. As a result, we are confident that our estimates of wealth coefficients in the cointegrating vector match very closely the true long run marginal propensity to consume for Italian households.

JEL classification: E21 C32

Keywords: wealth effects, consumption, housing

Contents

* Bank of Italy, Economic Outlook and Monetary Policy.

1. Introduction

The sizeable rise in house prices that started during the nineties in all main industrial countries – with the important exceptions of Germany and Japan – and in the largest emerging economies (IMF, 2004) has led to fresh interest in the empirics of wealth effects on consumption amid rapid innovation in financial markets. Differently from received theoretical literature, the debate is centred on the different channels through which housing as opposed to financial assets may affect consumer spending and, combined with residential investment, aggregate demand as a whole.

Interest in this issue has gained further momentum recently, given the financial turmoil that show up in August 2007 in the US sub-prime mortgage market that trigged liquidity distress in financial markets at a global level. The wide implications for monetary policy are also revealing, as "*the uncertainty around housing-related monetary transmission mechanism provides one further reason why monetary policy will continue to be an art, albeit one that makes use of science*" (Mishkin, 2007).

Indeed the question is not simple as certified by the controversial results in the available empirical literature, despite the recently prevailing view in favour of a significant and large housing wealth effect whose size, compared with financial assets, seems to change according with country specific factors.

In this paper we focus on the recent experience of the Italian economy, where wealth accumulation seems to outperform international comparison. At the same time, the research effort aimed at the possible link with consumer spending has not been as intense as in other advanced economies.

To the best of our knowledge, recently only Guiso, Paiella and Visco (2005), using the micro data of the *Survey of Households Income and Wealth* run biannually by the Bank of Italy, has focused on this issue. There is instead a substantial lack of research based on macro analysis, where more emphasis has been put on the different contributions coming from pension and non-pension wealth rather that from financial and non-financial assets.

We contribute to the debate by looking at the Italian aggregate time series to estimate the link between consumption, financial and housing wealth over a 27-year sample period (1980-2006), controlling for the role of income and common exogenous drivers, such as the brisk drop in interest rates on the eve of European Monetary Union. For this purpose, we have exploited cointegration analysis to study the long run relationships between the variables involved.

The empirical literature often interprets the coefficients of such relationships as long run elasticities of consumption with respect to income and wealth. Though this practice is quite convenient, these parameters summarize the correlation between the *permanent* movements in the aggregates, because they are based on the existence of a common stochastic trend.

On the contrary, they are not indicative at all of the link between households spending and transitory fluctuations in income and wealth. As a consequence, as first suggested by Lettau and Ludvigson (2004), once we have estimated the long run correlation among the variables of the system, we need to assess whether there exists

(transitory) movements in income and wealth that are not associated with (permanent) changes in consumption. This will allow a correct reading of the cointegrating vector.

Given the increasing relevance of this issue in recent literature, we have adopted the variance decomposition approach put forward by Gonzalo and Ng (2001) to assess the share of quarterly fluctuations which are due, respectively, to permanent and transitory shocks. This can be done easily by exploiting the information directly available from the estimated VECM for Italy. This accomplished, we have all the necessary information to correctly interpret our cointegration outcomes. Among the main results we find a long run marginal propensity to consume in the range, respectively, of 1.5-2 and 4-6 cents out of each Euro increase in housing and financial wealth.

This paper is organized as follows. Section 2 describes some stylized facts about recent developments in the Italian economy, with a particular focus on the household propensity to save and on the housing market. Section 3 delivers a summary of the theoretical predictions about wealth effects, followed in Section 4 by a review of empirical evidence available for Italy. After sketching the theoretical set-up adopted for our analysis (Section 5), we describe the data set and report upon the preliminary analysis (Section 6). Main empirical results are presented in Section 7, followed by a variance decomposition exercise (Section 8). The final section presents a brief conclusion.

2. Searching for macrofacts in Italy

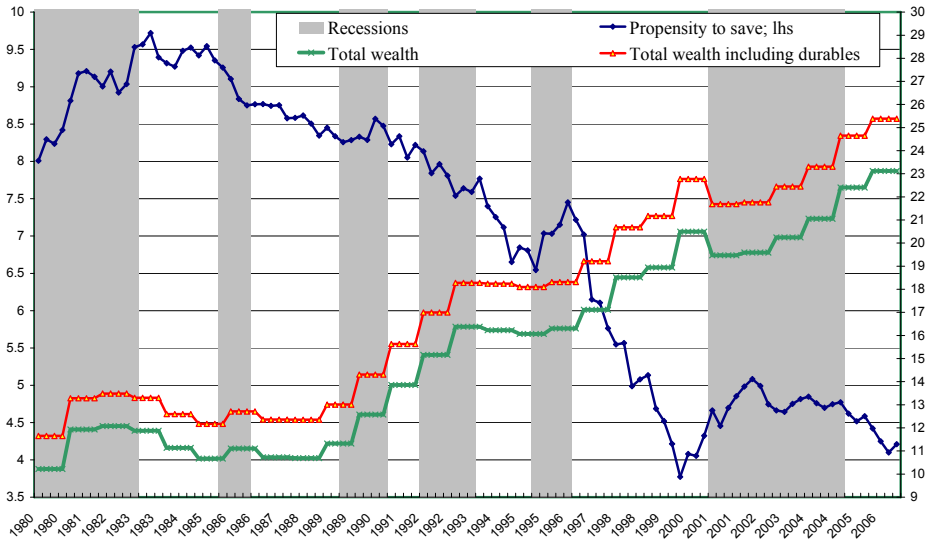
The propensity of Italian households to save had long been particularly high among industrial countries, being still on the rise in the early eighties (Figure 1). Since then, it has shown a brisk deterioration that, apart from a temporary break between 1995 and 1997, has gradually brought the saving rate down to a low of 10 per cent by the end of the nineties. At the beginning of the current decade, private savings first showed a partial recovery, then resumed their negative trend; at the end of last year they were around 10 per cent of disposable income of consumer households.

Looking for explanations of recent developments, a first candidate is a trend reversal in key factors that had previously sustained high personal savings, namely the high productivity performance that drove up income expectations and lagging capital markets which prevented the scope for consumption smoothing (Guiso, Jappelli and Terlizzese, 1994). On one side, since the middle of the nineties the productivity trend has turned negative (Bassanetti *et al*., 2004), plausibly curbing consumer spending in the medium term; on the other, the development of the capital markets accelerated with progress in financial liberalization at domestic level and European integration, peaking at the start of monetary union in 1999, and plausibly alleviated liquidity constraints for Italian households.

These forces might have offset the demand for precautionary saving that probably originated from increased economic uncertainty amid a sequel of reforms in social security and labour markets, that began in the early nineties.

Figure 1

Propensity to save and wealth income ratio



Sources: Elaborations based on data from Istat and the Bank of Italy.

The recent drop in personal saving went in association with a brisk acceleration in real and financial wealth held by consumer households. Following broad stability throughout the eighties, their total wealth to income ratio soared from around 5 at the start of the recession in 1992 up to 7 in 2000, eventually peaking at 7.8 at the end of last year (Figure 1).[1] Based on a tentative exercise in which we have disentangled the main contributors to changes in financial and housing wealth, it seems that since the start, in 2000, of a prolonged economic stagnation the deliberate investment of savings on the part of consumer households has explained an increasing part of the accumulation in financial assets (Figure 2).[2]

This result occurred despite a negative performance in asset prices – which apparently caused a huge devaluation in financial wealth – pointing to the possibility that a precautionary motive gradually gained momentum to drive up equity stock as a buffer against an uncertain long-term perspective. At the same time, the savings channelled to the residential market have kept increasing, even if playing a very minor role in driving very fast growth in housing wealth, as long as the positive trend in house prices markedly accelerates.

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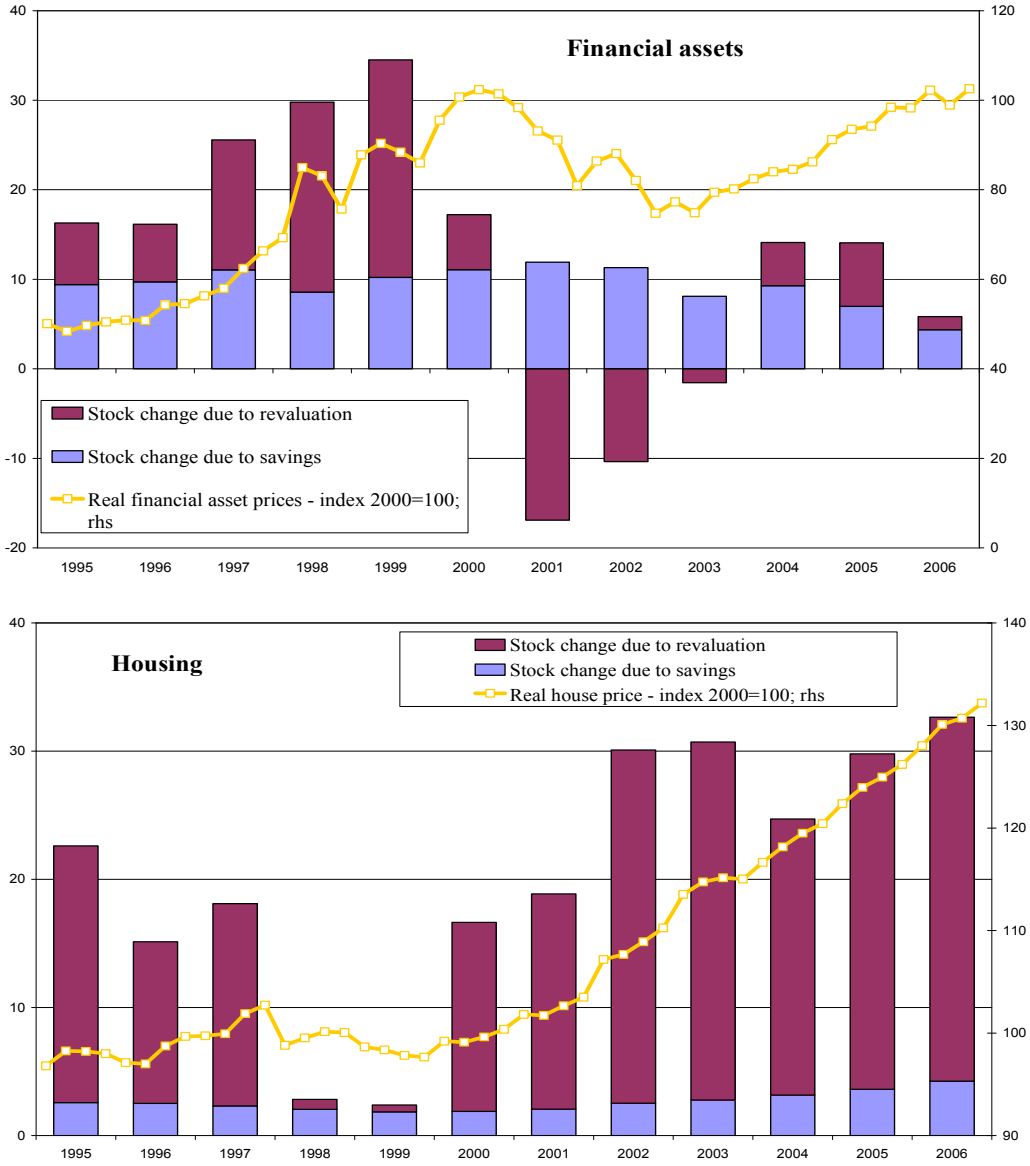
1 If durable stock is included, over the whole period, total wealth climbed from 5.5 to 8.6 as a ratio of disposable income of consumer households.

2 In the exercise, the revaluation of financial assets is estimated by the quarterly difference between changes in stock and the value of flow, as the latter reasonably accounts for sole saving/investment transactions. Alternatively, we have first constructed a weighted price index of equities and bonds held by Italian households (the same which in Figure 1 is shown net of consumption inflation), then we have multiplied its quarterly changes by the stock of one period earlier; the ensuing values seem broadly in line with the previous method, even if a higher volatility in the quarterly frequency causes an imperfect overlapping in the annual measures of revaluation. In the case of housing, we first estimated a measure of invested saving by evaluating at market prices the number of new dwellings sold, after adjusting for their average size in square meters; then we retrieved a revaluation measure as the difference between total change in stock and the estimated saving channelled to housing. As before, we checked that results are in line with those resulting from multiplying the changes in prices with the housing stock one period earlier. Interestingly, the amount of savings we have singled out in the changes of financial and housing wealth closely proxies, as a ratio to disposable income, the true propensity to save.

Figure 2

Changes in stock and real prices
(percentage values of disposable income if not otherwise stated)



Source: Elaborations on data from Istat and Bank of Italy

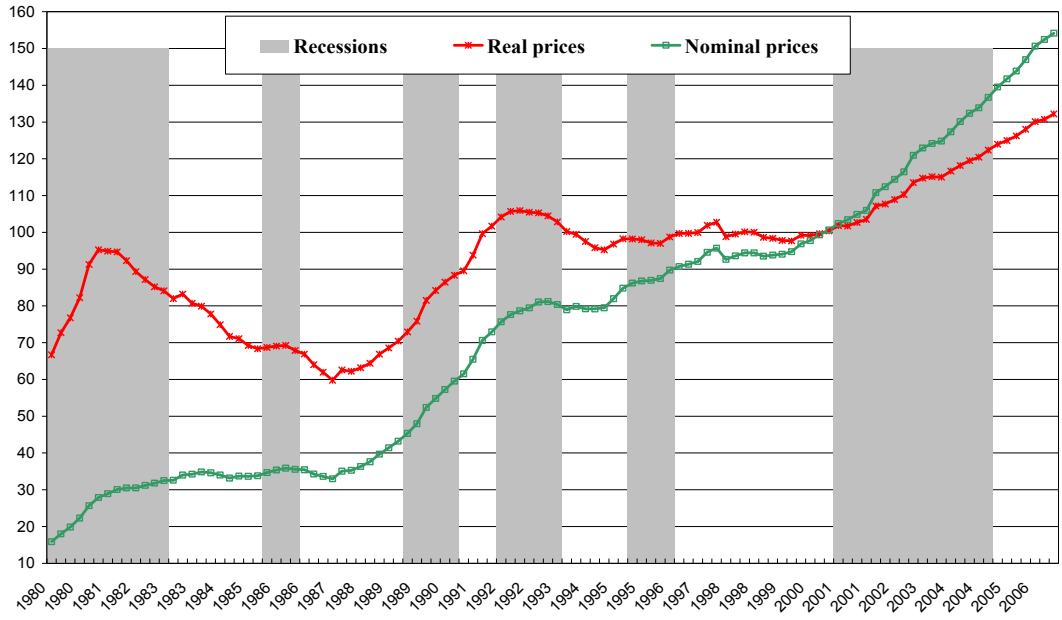
As a whole, these pieces of preliminary evidence seem to imply that the traditional arguments for a positive direct wealth effect may play a less important role in explaining the medium term drop of Italian personal saving than implied by a glance at Figure 1. Anyway the size of wealth effects needs to be tested more carefully against a long lasting period of rapid changes in the market value of assets, with particular reference to real estate.

Since the start of the current decade, the average annual growth in house prices in Italy has been around 5 percent, net of consumer price variation, or by half a percentage point higher than in the euro area as a whole, although lower than in Spain and France (around 11 and 9 percent, respectively). Differently from previous business cycles, in recent years the performance in house prices has not been affected by the state of general economy, with revaluation keeping increasing in momentum while domestic activity

virtually stagnated (Figure 3). Following the peak of a 10 per cent increase reached in 2002, house prices have gradually slowed down, more recently reflecting higher mortgage interest rates.

Figure 3

House prices and business cycles
(index 2000=100)



Sources: Elaborations on data from Istat, Il Consulente Immobiliare and the Bank of Italy

The upward trend in prices has entailed a strong increase in the value of the real estate wealth of Italian households, which at the end of 2006 accounted for almost 60 per cent of total wealth, or 20 percentage points more than financial assets (in 2000 the gap was below 4 points, with housing's share at around 51 per cent of total wealth). The home ownership rate now stands at 72 per cent among Italian households, or some 10 points higher than the estimated value for the euro area as a whole; among the main industrial countries, the Italian lead is below 5 points with respect to UK and USA, around 15 and 30 points with respect to France and Germany, respectively.

3. The theory of wealth effects

According to the life cycle model, households smooth over time their consumption spending on the basis of intertemporal budget constraint, given by the sum of the discounted flow of future expected income and the current endowment in wealth. Smoothing is achieved by borrowing when young against higher expected future income, then repaying the debt when income actually raises and consuming out of accumulated wealth when retired. It turns out that consumption expenditure depends on permanent income and initial wealth, besides life expectancy and time preference. In this framework, an unexpected and permanent increase in wealth entails a roughly equal rise in consumption in every future period that households expect to live. According to the theory, therefore, the marginal propensity to consume out of wealth is significantly positive and increases with age.

Since the standard model is mostly aimed at explaining long run changes in consumption, an additional prediction that has exerted a long lasting impact on macro-modelling is that wealth composition does not matter: households instantaneously adjust their spending by the same amount against either financial or real additional wealth. This is partly related to the key theoretical assumption that capital markets are complete, which rules out liquidity constraints as well as transaction and borrowing costs due to imperfect information. As a consequence most ingredients of the current debate regarding housing prices, housing finance and consumer spending are simply missed in the standard model. Nonetheless, in recent times the challenge of facing the effects of brisk asset price changes on economic and financial stability makes it increasingly important, for monetary policy purposes, to understand how a rise in the different forms of wealth transmits to household consumption.

For this purpose, following Deaton and Muellbauer (1980) and Poterba (1984) we rearrange the intertemporal budget constraint to take account of two different wealth components: housing and financial assets. This allows one to take into account that while financial assets are mainly a liquid store of value whose holding implies negligible costs, dwellings may be considered both a consumption good, specifically providing housing services, and an investment good that requires maintenance costs to keep the structure from depreciating.[3] The budget constraint becomes:

$$C + p(R+\delta)WH = Y^P + R(WFI_0 + pWH_0)$$

where C is non-housing consumption, p is the relative price of housing compared to a non-housing basket of goods, R is the real interest rate and δ is the depreciation rate, Y^P is a measure of permanent real income, WFI_0 and WH_0 are the initial endowments of financial and housing wealth. Thus $p(R+\delta)WH$ is the cost of housing services, with $p(R+\delta)$ being the real user cost.[4] The effect of a permanent increase in relative house prices on non-housing consumption is given by:

$$\frac{\partial C}{\partial p} = RWH_0 - (R+\delta)WH + (R+\delta)\varepsilon WH = RWH_0 - WH(R+\delta)(1+\varepsilon)$$

where ε is the elasticity of housing demand to its own price. Three transmission channels are at work:

- a positive direct wealth effect, since housing endowment is now worth more: RWH_0;[5]
- a negative income effect, due to higher costs of housing services: $(R+\delta)WH$;
- a positive substitution effect, depending on the own price elasticity of housing:

$$(R+\delta)\varepsilon WH.$$

Extending consumer expenditure to include imputed consumption of housing services, namely $CH = (R+\delta)WH$, the effect on total consumption $CT = C + CH$ is given by:

$$\frac{\partial CT}{\partial p} = RWH_0 - (R+\delta)WH$$

with the substitution effect now collapsed compared with the previous result.

[3] These add to insurance costs and fiscal burden, such as property taxes.

[4] This holds under the simplifying assumption of a constant p expected for the future.

[5] This effect is much the same as for financial assets.

It turns out that the size and sign of the residential wealth effect depend mostly on the following two factors:

i) the accurate measurement of the cost of housing services. If consumption of housing services is imputed on the basis of the real cost $(R+\delta)$, the housing wealth effect is likely to be negative on total consumption (CT), turning positive on non-housing consumption (C) since ε is plausibly lower but close to unity. If the real interest is instead omitted, the sign on total consumption would depend on $(R-\delta)$, while confirming positive on non-housing consumption. In any case, however, this standard model predicts that housing wealth effects, proxied at largest by $R-\delta(1+\varepsilon)$, would be lower than financial ones, proxied by R.

ii) the rate of home ownership, as proxied by the distance between WH and WH_0. In fact, while a positive shock to financial wealth has substantially a null effect on consumption of a non-holder (apart from a possible impact on confidence climate and expected future income), a permanent increase in relative house prices may affect both renters' and owners' optimal choice. Both would suffer from a negative income effect, but for owners this would be somewhat offset by a positive wealth effect. At the aggregate level, the balance between the two conflicting impacts is thus expected to turn positive to the extent that the rate of home ownership increases over time.

A third factor, though not modelled in the framework above, refers to differences in the processes that generate prices of financial and housing assets. Indeed, available international evidence points to much less volatility in the latter, with signs of stronger serial correlation (Leamer, 2007; OECD, 2005). This implies that, given an observed common shock in both asset prices, households may assign a greater importance to the higher permanent component for housing than for financial net worth. The argument would dampen the gap between the effects on consumption from the two different components, otherwise expected in a life cycle approach.

Finally, there is another important feature by which wealth effects may differ, namely that dwelling purchases are usually highly leveraged and imply transaction costs, related to the indivisibility and uncertainty regarding the true asset value which are much higher than for investments in financial stocks. As testified by the current debate about the housing finance revolution, this raises important implications for consumer spending that are missed by the standard theory, due to its simplifying assumption of complete capital markets. When information asymmetries are considered, the size of housing wealth effects proves to be increasing with the rate of innovation in mortgages and, more generally, with financial liberalization. As it is stressed in Mishkin (2007) and Muellbauer (2007), in the long run the main reason is that well developed financial markets make housing wealth very liquid, with two important implications: i) reducing the negative income effects on potential first buyers, especially when young, as far as the requirement for down-payment is relaxed and the savings to be channelled in a housing deposit shrink; ii) increasing wealth effects of owners, as far as they are available to downsize their dwellings, especially when old, without incurring high transaction costs. An additional effect, the equity withdrawal from housing or the supply of higher loans to households due to an increased value in collateral, is likely to play a minor role in the long run, when consumption remains anchored to the fundamental of total resources the households command, net of fiscal and financial costs. Anyway equity withdrawal might play an important role in the short run, since it increases the ability of households, otherwise credit constrained under asymmetric information, to smooth consumption over time.

In summary, the predictions of the standard life cycle model extended to include the main features by which housing and financial wealth may differ implies that the first

may exert a lower effect on long run consumption under complete capital markets; the effect would become larger as the rate of home ownership increases and, under a more realistic world of asymmetric information, as financial liberalization deepens. This helps to explain why housing wealth effects may differ across countries and why they outperform financial assets where the housing finance revolution has proceeded most.

These predictions have recently been subject to increasing criticism as to the significance of the wealth effect itself when the adjustment of stocks is made endogenous too, or when we move from a partial to a general equilibrium analysis (Lettau and Ludvingson, 2004). At the same time, wealth might only exert an indirect effect, for example when a strong stock market performance proves to be a leading indicator of better general economic perspectives, which drive up both asset prices and households' expenditure.[6] These arguments enrich the reasons for ambiguities in the theoretical analysis of wealth effects, which are passed to the validation of empirical tests.

4. Empirical evidence: an overview for Italy

We might think that on balance theoretical literature suggests that the wealth effect plays some role in affecting consumption but it is not precise enough to point to either the size of the overall effect or the relative impact of housing versus financial wealth. Many efforts have thus been made on the empirical side within a large variety of countries, starting from the first macro-estimate for the US economy during the sixties. At the applied level results do not deliver a sound solution either, as evidence proves dependent on the estimation strategy, such as variable choice and measurement, the time period covered, the statistical model, the kind of data, the within or cross country approach.

Among the main controversies, recent debate has been increasingly focusing on the role of different components of wealth, since the wave of innovation in housing finance seems to have made the link between sky-rocketing house prices and consumer spending more uncertain and country-specific. Overall, it seems that housing wealth plays quite a strong role in the US, with the marginal propensity to consume (MPC) ranging between 5 and 10 cents out of a one dollar increase in dwellings against 3 to 6 cents for financial stock (Case, Quigley and Shiller, 2005, Carroll, Otsuka and Slacalek, 2006; CBO, 2007). For the rest of industrial countries evidence is more mixed, with available evidence for continental Europe broadly pointing to a more active role for financial assets than housing wealth in the long run (Ludwig and Sløk, 2002; Altissimo *et al.,* 2005).

When focusing on the Italian economy, there is not a long empirical tradition regarding the consumption effect of financial versus housing wealth. In a cointegration analysis covering the long sample between 1951 and 1992 and controlling for pension wealth, Rossi and Visco (1995) finds the MPC out of financial and real wealth as a whole at around 3 cents in the long run. Updated evidence, based on years up to 1998, shows that the same propensity has increased to around 5 cents in the long run, against a lower MPC for pension wealth (at around 2 cents) as important social security reforms in 1992 and 1995 added uncertainty about future benefits (Zollino, 2001).

[6] Further criticism points to non-linearities in the utility function and different mental accounts that may limit the willingness to spend out of saving accumulated for specific reasons, such as long term retirement or bequest motives. For a survey of main recent contributions on these issues, see Ludwig and Sløk (2002) and Belsky and Prakken (2004).

In a set of recent contributions, which confirm a size of total (non-pension) wealth effect of between 3 and 5 cents in the long run despite a shorter time period, financial assets prove to exert a stronger role than housing wealth on long run consumption. An exception is provided in Catte, Girouard and Price (2004), where evidence covering the period 1975-2002 points to MPC as small as 1 cent out of one euro increase in either financial or housing wealth. More generally the coefficient of the latter proves surprisingly negative, showing occasionally no statistical significance (Kennedy and Andersen, 1994). In a cointegration analysis covering the period 1980-1996, Girouard and Blöndal (2001) found a negative MPC out of housing at around 3 cents in the long run, compared with the positive propensity of 5 cents estimated for financial assets.[7] By adopting the new methodology proposed by Carroll, Otsuka and Slacalek (2006) which is mainly base on estimated slugginess in aggregate consumption, Slacalek (2006) finds evidence for Italy of a particularly strong MPC out of financial assets, at around 10 cents or the second highest value among the main industrial countries; on the contrary, the propensity for housing proves the lowest, resulting negative at 1 cent.

At household level, based on repeated cross-sections retrieved from the Bank of Italy's biannual Survey on Household Income and Wealth (SHIW) over the period 1991-2002, Guiso, Paiella and Visco (2005) estimate a marginal propensity to consume out of housing wealth of about 1.6 cents, which is barely one half of what they found for financial assets. By splitting the sample, they find that homeowner propensity to consume out of housing wealth is close to 3.5 cents, in line with evidence for US and UK, with signs of a negative effect on renters' consumption. By adopting almost the same data set, in the full sample Paiella (2004) finds a somewhat lower MPC out of the housing wealth effect and a much higher one for the financial wealth; in the latter case, however, the coefficient risks to be biased upward due to under-reporting on the part of the wealthiest. Grant and Peltonen (2005), based on the panel section in the SHIW over the period 1989-2002, found that the housing wealth effect on consumption is not statistically significant, with a MPC size far lower in comparison with the value of 5-6 cents estimated for financial wealth.

5. The theoretical set up

We restate the intertemporal budget constraint of a representative consumer according to the formulation given by Campbell and Mankiw (1989):

$$W_{t+1} = (1 + R_{W,t+1})(W_t - CT_t) \tag{1}$$

where CT_t is total consumption and W_t is total wealth, made of the human (HU_t) and non-human (A_t) components; $R_{W,t}$ is the net return on aggregate wealth.[8] Solving forward, log-linearizing and defining $r \equiv \log(1+R)$ yields:[9]

$$ct_t - w_t = \sum_{j=1}^{\infty} \rho_w^j (r_{w,t+j} - \Delta ct_{t+j}) \tag{2}$$

[7] By adopting the same data set, time period and methodology but a different deterministic control (in order to take account of main financial innovation events), Boone, Girouard and Wanner (2001) find a similarly negative propensity for housing, but a bit higher positive propensity for financial assets, at around 6 cents in the long run.

[8] Labour income does not appear explicitly because tradable human capital is included in aggregate wealth.

[9] In the equations of this section we will always neglect the constant component of a linearization process.

which is a log-linear version of the infinite horizon budget constraint.[10] Extending the approach of Lettau and Ludvigson (2004), we then disaggregate non-human wealth into two different components, housing (WH_t) and non-housing (WNH_t) wealth, yielding:[11]

$$W_t = HU_t + A_t = HU_t + WNH_t + WH_t \tag{3}$$

with non housing wealth given by the sum of financial (WFI_t) and valuables (WVA_t) net worth. Simple algebraic manipulation lead us to:

$$w_t \cong \gamma_{hu} hu_t + \gamma_{wnh} wnh_t + (1 - \gamma_{hu} - \gamma_{wnh}) wh_t \tag{4}$$

where γ_i is the steady state share of asset i in total wealth. Substituting in (2) gives:

$$ct_t - \gamma_{hu} hu_t - \gamma_{wnh} wnh_t - (1 - \gamma_{hu} - \gamma_{wnh}) wh_t = \sum_{j=1}^{\infty} \rho_w^j (r_{w,t+j} - \Delta ct_{t+j}) \tag{5}$$

The non observability of human capital hu_t prevents the empirical application of this equation. Once again, we follow Lettau and Ludvigson (2001) to face the issue and assume that income Y_t is the annuity value of human wealth: $Y_t = R_{h,t+1} HU_t$. Some additional manipulation leads to:

$$hu_t = y_t + z_t \tag{6}$$

where z_t is a mean zero stationary random variable. Substituting and taking expectations of both sides yields:

$$ct_t - \omega_{hu} y_t - \omega_{wnh} wnh_t - (1 - \omega_{wnh} - \omega_{wh}) wh_t = E_t \sum_{j=1}^{\infty} \rho_w^j (\varpi_{t+j} - \Delta ct_{t+j}) + \omega_{hu} z_t \tag{7}$$

where

$$\varpi_{t+j} = \omega_{hu} r_{uh,t+j} + \omega_{wnh} r_{wnh,t+j} + (1 - \omega_{hu} - \omega_{wnh}) r_{wh,t+j}$$

All the variables on the left hand side of equation (7) are now observable. Moreover, if we assume that those on the right hand side are stationary, then consumption, income, housing and non-housing wealth should be tied together by some cointegrating relation, upon which we next concentrate in the empirics of the second part of the paper.

6. Data and preliminary analysis

In this section we briefly describe the data set and the definition of variables, then we summarize the results of a preliminary analysis and the kind of deterministic control we consider in the VECM estimation.

6.1 The data set

The empirical analysis focuses on the time range 1980q1-2006q4 exploiting several data sources.

Total households' consumption (CT_t) is readily available from quarterly national accounts and includes spending on non durables, durables and services.

[10] $\rho_w \equiv 1\text{-}exp(ct\text{-}w)$; where ct and w are at their steady state value.

[11] Lettau and Ludivigson's main focus is just on human and non-human wealth.

Households' gross disposable income is published within national statistics only at annual frequency; in this paper we resort to a quarterly disaggregation of this series (Y_t) which is regularly estimated by the Bank of Italy.

Households' wealth is taken from a new database provided by the Bank of Italy whose aggregate series are employed for the first time in this paper.[12] The new *housing wealth* series (WH_t) are at annual frequency and span from 1990 onwards. Consistently with the methodology that led to their estimation,[13] we extend the data backwards exploiting, for the decade 1980-1989, the dynamics of time series on new dwellings[14] and on house prices. Once yearly data span the whole sample, quarterly disaggregation is achieved through the dynamics of residential investment published in national accounts. Wealth effects will be estimated also for a measure of housing assets net of mortgages (WH_A_t).

Official data for net *financial wealth* (WFI_t) are supplied within financial accounts from 1995 onwards; backward estimation is obtained resorting to older versions of the same data, although conformed to previous accounting standards. When WH_A_t is used; the value of financial assets is corrected accordingly (i.e. gross of mortgage debt; WFI_A_t).

The stock of valuables (WVA_t) is evaluated on the basis of information gathered within the Survey of Household Income and Wealth (SHIW) run by the Bank of Italy;[15] annual data are disaggregated at quarterly frequency through linear interpolation.

Finally, interest rates (R_t) are given by the returns on Treasury bonds with maturity longer than one year and deflated according to expected inflation, and public consumption (CC_t) is directly available from quarterly national accounts.

6.2 Preliminary analysis

As reported at the end of section 4, equation (7) suggests the possible existence of a cointegrating relation among consumption, income, housing and non housing wealth. According to a variety of tests, we find evidence in favour of a single unit root in the data generating processes of each of these variables (see Appendix I). We also applied the same tests to the proxy of the expected flow of real interest, namely R_t, and the hypothesis of a unit root cannot be rejected.[16] We have done so also for real public consumption (CC_t), with similar findings. The latter results allow us to explore the possibility of testing if R_t and/or CC_t may play as common exogenous drivers for the long-run equilibrium.[17]

Before estimating a VECM and testing the existence of cointegration, it is important to consider the consequences of two major economic events that happened within the sample period. In 1992-93 the Italian economy witnessed a severe currency

[12] The dataset is the outcome of a research project run in the last two years and provides a remedy for the relevant gap in the national statistical information set.

[13] For details, see Cannari and Faiella (2007).

[14] Data on new dwellings are provided by Cresme which is a research institute focused on the Italian construction industry; house prices for the eighties are estimated based on new dwellings in provincial capitals (see Muzzicato, Sabbatini and Zollino, 2002).

[15] For details, see Cannari, D'Alessio and Marchese in this volume.

[16] Note that the nominal interest rate R_t has been turned into real terms through a measure of inflation expectations. Results do not change if deflation is achieved by adopting the households' consumption deflator. As discussed in the text, our aim is not to test the stationarity of variables on the right hand side of equation (7), namely the weighted average return on different kind of wealth, whose calculation is much more cumbersome.

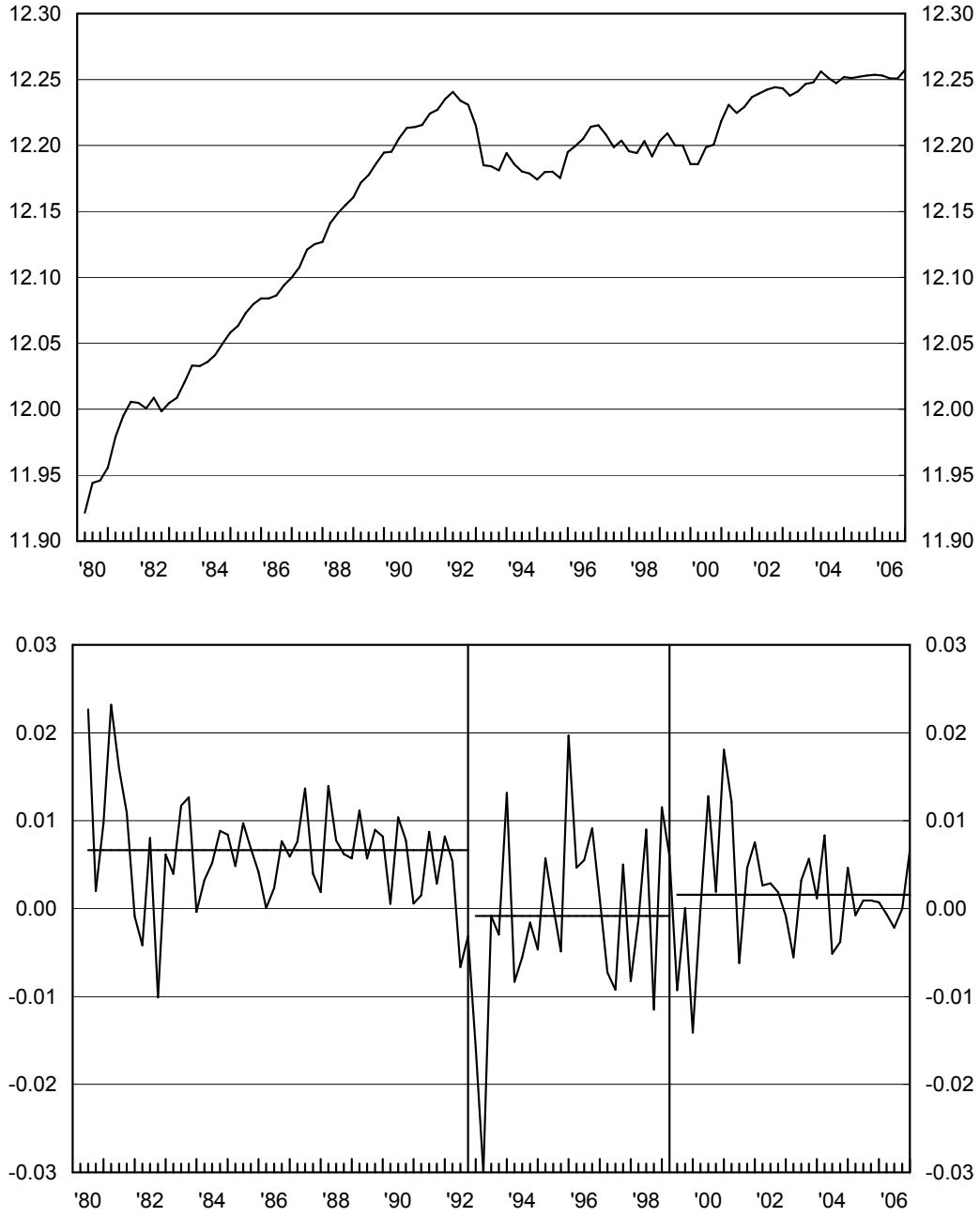
[17] We also tested the effects of other forcing variables, such as public consumption and changes in the unemployment rate, with the cointegration analysis remaining broadly unchanged, though with a worse diagnostic check.

crisis and a deep recession, with GDP falling for six quarters in a row. Following a 4.3 percent drop in 1993 as a whole, households' real disposable income took some six years to recover to the same level as before the recession. The slump was less dramatic for consumer spending, that recovered more rapidly from the annual 3.1 percent fall of 1993. Inspecting the graph helps one appreciate the level shift in the income and consumption series; wealth shows minor changes (Figures 4-7).

Figure 4

Households' gross disposable income

(levels and rates of growth of chain-linked values; in logarithms)

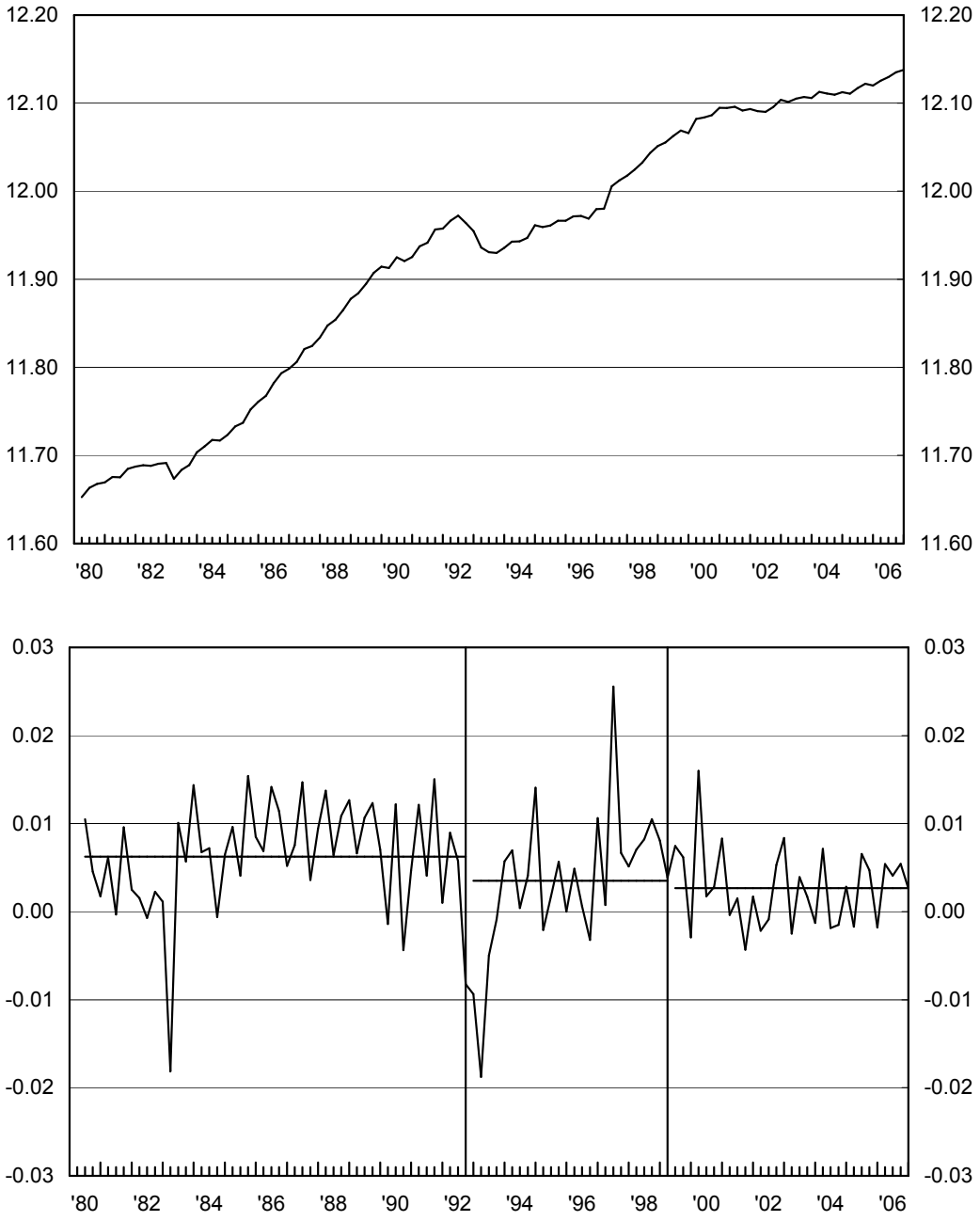


Sources: Elaborations based on data from Istat and the Bank of Italy.

Figure 5

Households' total consumption

(levels and rates of growth of chain-linked values; in logarithms)

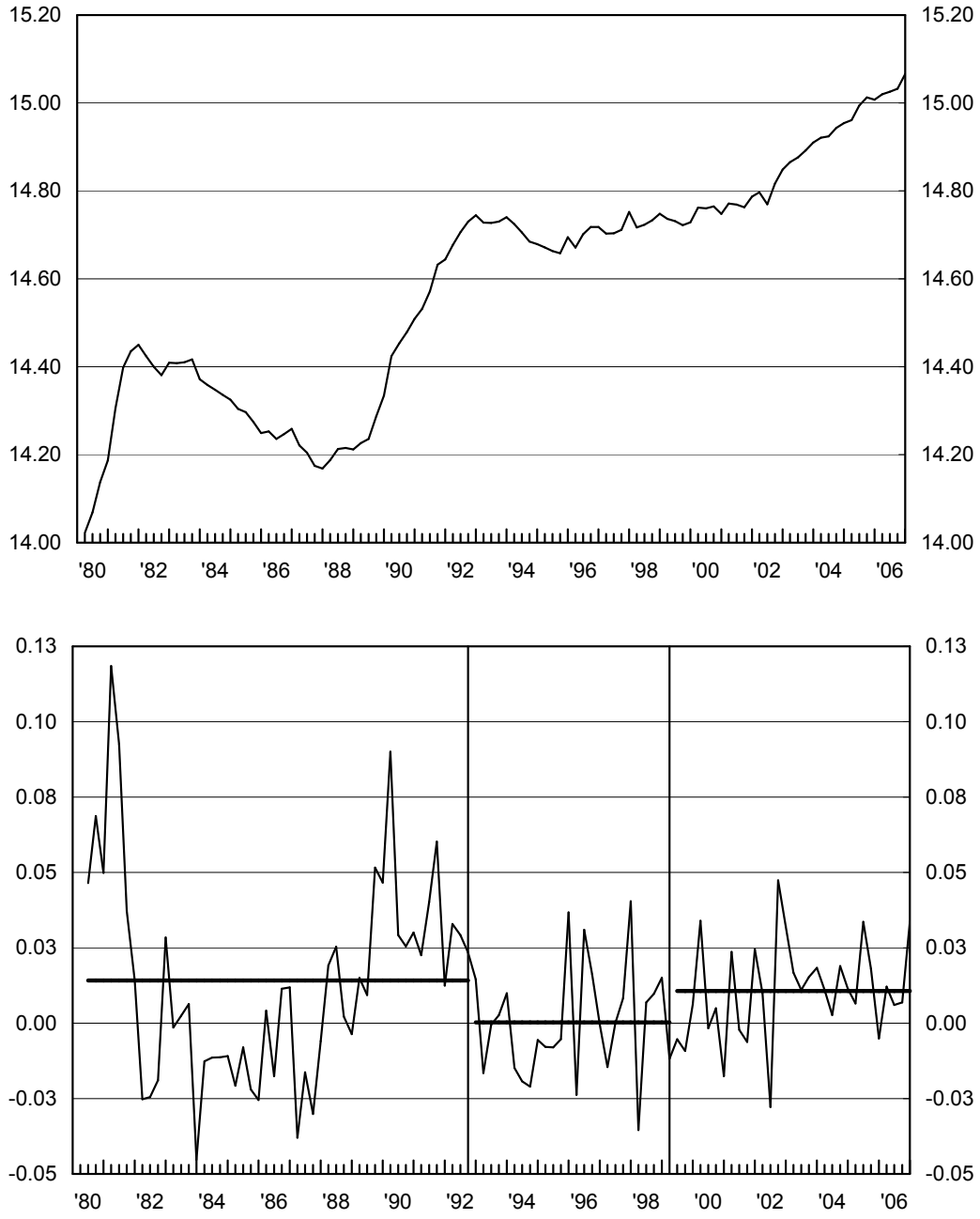


Sources: Elaborations based on data from Istat and the Bank of Italy.

Figure 6

Households' housing wealth

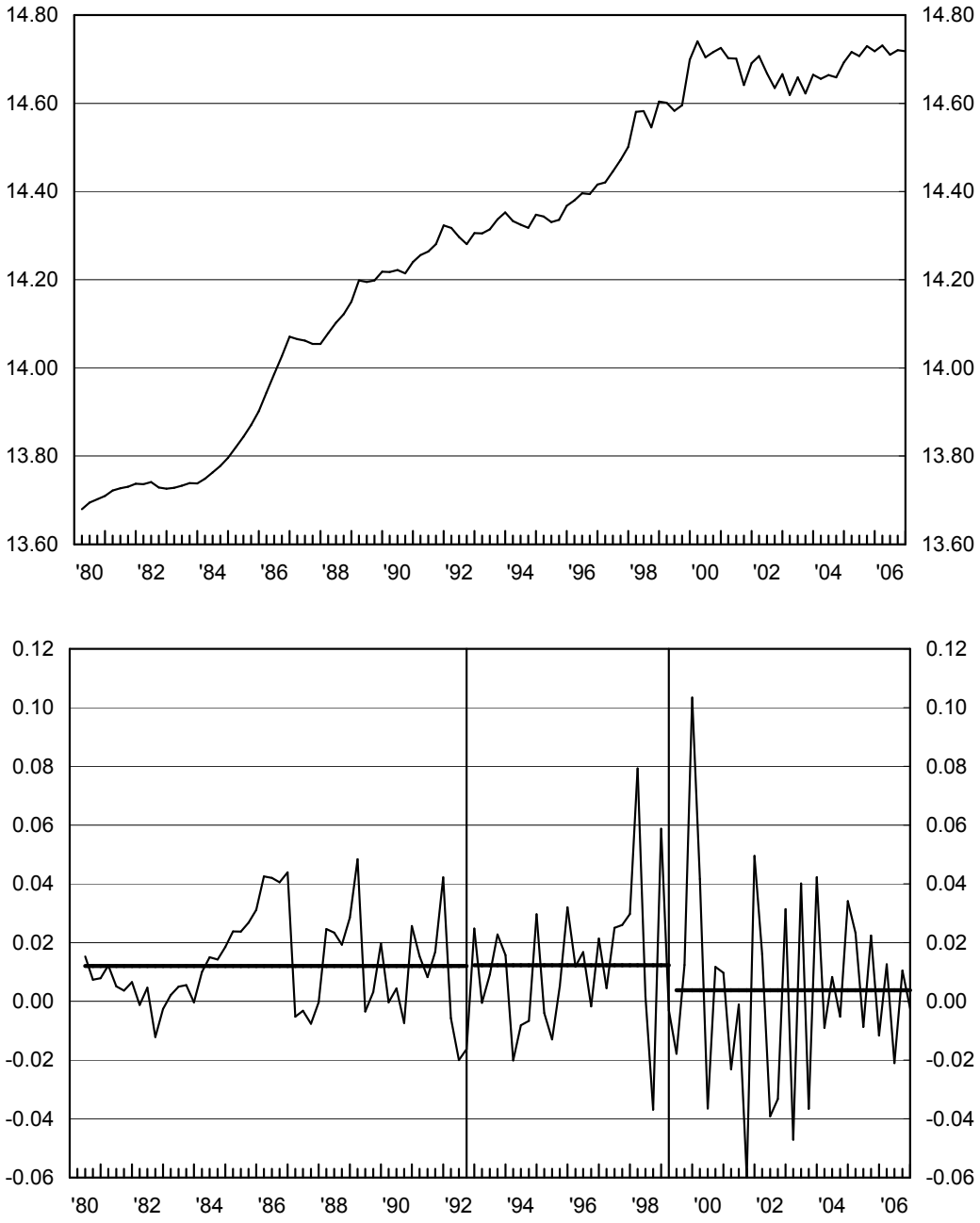
(levels and rates of growth of chain-linked values; in logarithms)



Sources: Elaborations based on data from Istat and the Bank of Italy.

Figure 7

Households' non-housing wealth
(levels and rates of growth of chain-linked values; in logarithms)



Sources: Elaborations based on data from Istat and the Bank of Italy.

The deep nature of these effects put under pressures the productive and institutional setting of our economy as a whole, calling for urgent and permanent revisions in key fields of economic policies; they were actually implemented as for restrictions in social security provisions and public employment, plausibly curbing permanent income on the side of Italian households (Miniaci and Weber, 1999).

The second event happened some years later when, after the huge fiscal effort made in 1997 and 1998 to meet the financial standards set by the European Monetary Union, the Italian economy joined the single currency area. The institutional changes that followed, including new exchange rates and monetary policy regimes, plausibly represent another permanent innovation in the data generating processes. On one side, Italian export performance came under the new pressure of irreversible parity, stressing the need for a more efficient production plans, on the other side the mortgage interest rate almost halved between 1997 and 1999, thus easing the access to the credit market for consumers.

Permanent economic shocks do not disappear in cumulation and, therefore, have effects on stochastic trends. Moreover, they may have varied impacts on the variables of the system and thus their effects may not cancel out in the cointegrating relation. All this suggests the introduction of two dummy variables in 1992q3 and 1999q1, restricted to lying in the cointegrating relationship to control for the possibility of level shift in the long-run equilibrium; their statistical significance will be tested in the next section.[18] Alternatively, we have tested the effects of two forcing variables, namely public consumption and real interest rate, which we think better summarize the economic changes driving level shifts in cointegrating relation.

7. Econometric results

We estimate the VECM $\Delta X_t = \mu+\alpha\beta' X_{t-1}+\Sigma_k \Gamma_k \Delta X_{t-k}+\varepsilon_t$ for the vector $X_t = (ct_t, y_t, wh_t, wnh_t, r_t)'$, where variables are in real terms and lower case letters indicate logs;[19] non housing wealth is $WNH_t = WFI_t + WVA_t$. In the first instance, we do not introduce any deterministic control in the VAR, whose length has been set equal to three.[20]

7.1 Rank of cointegration

Results for the rank test are reported in Table 1; standard critical values appear alongside those to be considered when including the Bartlett correction in the statistics, in order to take into account the relatively small dimension of our sample. The hypothesis of one cointegrating relation ($r=1$) and four common trends ($n-r=4$) is accepted; the diagnostics of the system do not reveal major specification problems, apart from some violation of the normality assumption that will be easily fixed with the introduction of the unrestricted dummies mentioned previously.

The choice $r=1$ is robust across different VAR length specifications and independent from the presence of outlier controls; the same rank is assigned also when the system is reduced to contain a single measure of total wealth ($wtot_t$), instead of the two distinct forms wh_t and wnh_t. Some evidence in favour of $r=2$ arises when we introduce the two deterministic controls for the candidate level shifts, mainly depending

[18] Additionally, three unrestricted (i.e. outside of the cointegrating relation) dummy variables have been introduced to control for data outliers and to improve the diagnostic checks (for example, such an intervention dummy was used in 1997q2, when spending was boosted by fiscal incentives for old car replacement).

[19] Disposable income, housing and non housing wealth have been deflated with the deflator of consumption, while inflation expectations were used for real interest rate, as already outlined in section 6.1. Results remain substantially unchanged when the deflator of consumption is also applied to the interest rate.

[20] Information criteria pointed to a slightly more parsimonious specification, but we preferred to maintain a richer dynamic structure. The main results presented in the following of the section do not depend on the order of the VAR. Details of the lag length determination procedure are available on request.

on the precise timing at which we date the effect of the 1992 recession.[21] Nonetheless, considering the overall outcomes and the economic *a priori* based on our theoretical set-up, we set the rank *r=1*.[22]

Table 1

Cointegration trace test

Common trends *n-r*	Rank *r*	Trace test	Trace test*	Standard critical values	
				p-value	*p*-value*
5	0	84.208	70.808	0.002	0.040
4	1	44.406	37.441	0.101	0.331
3	2	21.485	17.033	0.338	0.646
2	3	8.600	6.849	0.411	0.601
1	4	2.183	1.800	0.140	0.180

Note: Critical values refer to a model with an unrestricted constant term.

(*): Bartlett corrected statistics.

7.2 Results

In the upper panel of Table 2 we report the estimates of the cointegrating vector β for the different model specifications obtained under the control for deterministic shifts in *1992q3* and *1998q1*, that invariantly resulted to be statistically significant, also when we control for the effects of real interest rate. As it results from a comparison with the table reported in Appendix II, the omission of the deterministic controls would mainly imply an implausibly lower coefficient for income, leaving the coefficients of the remaining variables broadly unchanged as well as their robustness and signs. This points to the opportunity to take into account the shocks that likely hit the variables of the system heterogeneously, as was the case particularly with the 1992 crisis, when the income drop was very pronounced and long lasting. Interestingly, as we comment later on, the statistical significance of deterministic controls vanishes at all only when we extend the forcing variables to include real public consumption.

Overall, estimates look statistically significant and economically plausible both in terms of sign and magnitude; further they are quite robust across different VAR specifications. Forward and backward recursive tests do not reject the hypothesis of constancy over the time range under analysis (Appendix III).

While the cointegrating relationship summarizes the *correlation* between the permanent movements in consumption, income and wealth over the sample period, the empirical literature often interprets the estimated parameters in terms of elasticities or MPC. As in Lettau and Ludvigson (2004), this is a reasonable practice as they come out from a regression of consumption on wealth, controlling for income and real interest

[21] Results are available on request. Simulated critical values have been considered with standard ones, in order to take into account the dependency of asymptotic distribution on deterministic components, in particular on structural breaks. The CATS software, which was used for estimations, includes procedures both for simulating critical values for non standard models and for applying the Bartlett small sample correction.

[22] The choice seems corroborated by inspecting the graph of the second possible cointegrating relation, which exhibits signs of persistency and thus suggests that the equilibrium error might be non-stationary. Further, the recursively calculated trace test for r=2 lies under the 5 percent critical value for most of the sample.

rates. The estimated parameters are also super-consistent and thus robust to regressors endogeneity.

Table 2

Cointegration coefficients

(normalized on the coefficient of total consumption ct)

	Model specification					
	(A)	(B)	(C)	(D)	(E)	(F)
	Cointegrating coefficients (β-vectors)					
Total wealth ($wtot_t$)	-0.167 (-3.697)	-0.178 (-3.494)				
Housing wealth (wh_t)			-0.076 (-3.896)	-0.074 (-4.056)		
Non housing wealth (wnh_t)			-0.109 (-2.680)	-0.118 (-2.937)		
Housing wealth net of mortgages (wh_a_t)					-0.076 (-3.723)	-0.047 (-3.879)
Non housing wealth gross of mortgages (wnh_a_t)					-0.093 (-2.004)	-0.191 (-8.463)
Disposable income (y_t)	-0.941 (-12.049)	-1.100 (-9.092)	-0.845 (-8.067)	-0.871 (-8.046)	-0.959 (-7.929)	-0.755 (-9.008)
Interest rate (r_t)		-0.037 (-2.302)		-0.009 (-0.901)	-0.014 (-1.211)	0.045 (6.162)
Public consumption (cc_t)						0.191 (4.181)
Deterministic controls in the cointegrating vector						
1992q3	-0.047 (-3.473)	-0.071 (-4.044)	-0.038 (-3.079)	-0.042 (-3.336)	-0.051 (-3.689)	
1998q1	-0.048 (-4.479)	-0.076 (-3.829)	-0.044 (-3.803)	-0.050 (-3.654)	-0.058 (-3.781)	
	Error correction loadings (α)					
Total consumption (ct_t)	0.033 (1.190)	0.054 (2.283)	0.028 (0.840)	0.020 (0.567)	0.030 (0.968)	-0.160 (-2.427)
Total wealth ($wtot_t$)	0.380 (4.009)	0.265 (3.343)				
Housing wealth (wh_t)			0.641 (4.355)	0.570 (3.681)		
Non housing wealth (wnh_t)			0.176 (1.172)	0.216 (1.374)		
Housing wealth net of mortgages (wh_a_t)					0.482 (3.426)	0.774 (3.064)
Non housing wealth gross of mortgages (wnh_a_t)					0.171 (1.288)	-0.091 (-0.339)
Disposable income (y_t)	0.206 (6.774)	0.178 (6.848)	0.260 (6.731)	0.285 (6.993)	0.255 (7.157)	0.308 (4.055)

Note: *t*-statistics in parenthesis.

Based on our estimates of the cointegrating vector, one percentage point increase in housing (wh_t) and non housing (wnh_t) wealth would be associated with an increase of, respectively, 0.08 and 0.11 percentage points in total consumer spending (column C of Table 2). These elasticities remain statistically unchanged when we correct for mortgage debts in both kinds of wealth (column E). In terms of marginal propensity to consume, the

outcomes state that a one euro rise in housing and non housing wealth would be associated with an increase of about, respectively, 2 and 4 cents in consumption.[23] If we concentrate on a single measure of total net worth ($wtot_t$; column A), the elasticity of consumption amounts to 0.17 and the marginal propensity to consume to 2.6 cents for a one euro wealth increase. Finally the MPC out of disposable income is in the order of 70 cents.

It is worth noting that these results are confirmed when we test for common external drivers leading the system of variables we considered, implying evidence in favour of a direct wealth effect. In this regard, in Table 2 we report estimates obtained by controlling for changes in the long-term real interest rate. This variable proves to be weakly exogenous under the different specifications we tested, and leaves the cointegrating vector substantially unchanged and largely significant.[24]

Additionally, we tested for the role of real public consumption as a proxy for fiscal discipline: alike R_t, CC_t proves weakly exogenous to our cointegrating system. Moreover, comparison of columns E and F shows that when the two variables are jointly modelled as common drivers, the case for deterministic controls drops. The result delivers an interesting economic content of the two regime switches that statistically show up in the dummies 1999q3 and 1998q1. On one side, the stringent curbing of public deficit over the nineties may have provided support to the households' spending plans in the long run much in line with a Ricardian neutrality mechanism; on the other, the fall in real long-term yields was presumably perceived by households as permanent too, and this contributed to shift their plans from saving to consumption. According to this specification, the marginal propensity to consume amounts to about 60 cents out of one euro increase in income, and about 6 and 1.5 cents out of similar changes in financial and housing wealth, respectively.[25]

7.3 Equilibrium correction

In the bottom panel of Table 2 we report the coefficients $\alpha = (\alpha_{ct}, \alpha_y, \alpha_{wh}, \alpha_{wnh})'$ of the error correction term ($\beta' X_{t-1}$) that give a measure of the reaction of each variable, at time t, to the disequilibrium of the system in period $t-1$. Consumption smoothness is confirmed by the fact that α_{ct} is not statistically significant; it turns out that the adjustment process towards the long-run equilibrium is mainly achieved via housing wealth, while non-housing net worth does not show significant error correction movements. This comes as no surprise, given the widespread diffusion of house ownership among Italian households, as shown by the particularly high rate of home ownership. Also disposable income contributes to smooth household spending, although to a much lesser extent.

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[23] Marginal propensities are obtained by dividing the elasticities by the average ratio of the corresponding variable to consumption over the full sample period.

[24] Weak exogeneity implies that the interest rate enters the cointegrating vector, but it does not contribute to the correction towards the long run equilibrium (namely, the loading coefficient α_r is not statistically different from zero). It also turns out that the coefficient in the cointegrating vector is not significant, as shown in the table.

[25] If pension wealth is included in the cointegrating system, its impact on consumption is negligible in the considered period, contrary to the positive effect found over a much longer time horizon (Zollino, 2001). It appears that the repeated changes in social security rules since 1992 have made benefits and retirement dates more uncertain, reducing the responsiveness of consumption plans to pension wealth.

8. Permanent - transitory decomposition

The cointegrating coefficients of the previous section are based on the existence of a common trend that ties together the long run movements of consumption, income and wealth. Thus, as already noted, they reveal the degree of correlation among the permanent components of the variables in the system, whereas they are completely useless for inference regarding the relationship between consumption and transitory changes in income and wealth. If a large degree of quarterly fluctuations of income and wealth was to be actually due to transitory events, then cointegrating coefficients would be poorly informative about their relation with households spending and should be cautiously interpreted. Consequently, as suggested by Lettau and Ludvigson (2004), once we have determined the long run correlation among the variables in the system, we need to assess whether there exist (transitory) movements in income and wealth unassociated to (permanent) changes in consumption.

We apply to our system X_t of I(1) variables the econometric framework proposed by Gonzalo and Ng (2001) to isolate the permanent ($\widetilde{\eta}_t^P$) and transitory ($\widetilde{\eta}_t^T$) shocks, as defined by the following two conditions:

$$\lim_{h\to\infty} \partial E_t(X_{t+h}) / \partial \widetilde{\eta}_t^{P'} \neq 0 \qquad \lim_{h\to\infty} \partial E_t(X_{t+h}) / \partial \widetilde{\eta}_t^{T'} = 0$$

Hinging on Gonzalo and Granger (1995), it can be shown that these shocks can be readily recovered with a two step procedure exploiting the information available in the VECM.

Starting from the multivariate moving average representation $\Delta X_t = \delta + C(L)e_t$,[26] in the first step the vector u_t of unorthogonalized permanent and transitory shocks is given by the following simple transformation of innovations (residuals) e_t:

$$u_t = \begin{bmatrix} u_t^P \\ u_t^T \end{bmatrix}_{r\times 1}^{(n-r)\times 1} = \begin{bmatrix} \alpha_\perp' e_t \\ \beta' e_t \end{bmatrix} = Ge_t \qquad with \qquad G = \begin{bmatrix} \alpha_\perp' \\ \beta' \end{bmatrix}_{r\times n}^{(n-r)\times n}$$

where $\alpha\perp'\alpha = 0$.

In the second step, orthogonality of $\widetilde{\eta}_t$ is achieved exploiting the Choleski decomposition of *cov(u)*: [27]

$$E(u_t u_t^{'}) = HH^{-1}$$
$$\widetilde{\eta}_t = H^{-1}u_t$$

We end up, therefore, with the Wold representation:

$$\Delta X_t = \delta + C(L)e_t = \delta + C(L)G^{-1}HH^{-1}Ge_t = \delta + D(L)HH^{-1}u_t = \delta + \widetilde{D}(L)\widetilde{\eta}_t$$

---

[26] *C(L)* is a distributed lag operator and e_t is a (*nx1*) vector with $E(e_t) = 0$ and $E[e_t e_s^{'}] = \begin{cases} 0 & if \quad t \neq s \\ \Omega & otherwise \end{cases}$

[27] Gonzalo and Ng outline that the Choleski decomposition is one of the many available alternatives for choosing *H*, which need to be lower block triangular. As a consequence, the $\widetilde{\eta}_t$ we found are not unique. Further, the Choleski decomposition, although being quite convenient, entails that the order matters; however it can be easily shown that in this setting a variable X_{it} can react to $\widetilde{\eta}_{jt}^P$ even if $j > i$, thus mitigating the effects of the recursive structure of the system.

that allows one to bring back the growth of each variable in the system to a function of the permanent and transitory shocks. Specifically, on the basis of our results, a cointegration rank equal to one implies the existence of three common stochastic trends (permanent shocks) and one transitory innovation. Gonzalo and Ng's decomposition enables the assessment of their role in the movements of consumption, disposable income, housing and non-housing wealth.

For brevity, in Table 3 we concentrate on a system with deterministic controls and report the share of the variance in the *h*-step forecast error attributable to the two kinds of shock.[28] In the top panel, we used the estimates discussed in the previous section; in the bottom part, the statistically insignificant error correction coefficients, α_c and α_{wnh}, have been set to zero, as suggested by Gonzalo and Ng. In both cases the three permanent shocks dominate the variance of consumption growth; this implies, in accordance with the prediction of the life-cycle model, that households' spending responds exclusively to changes in the permanent component of wealth and income.

Table 3

Forecast error variance decomposition – disaggr. wealth (*)

(orthogonalized VAR residuals)

	Δct_{t+h}-$E_t\Delta ct_{t+h}$		Δy_{t+h}-$E_t\Delta y_{t+h}$		Δwh_{t+h}-$E_t\Delta wh_{t+h}$		Δwnh_{t+h}-$E_t\Delta wnh_{t+h}$	
Horizon	Perm.	Trans.	Perm.	Trans.	Perm.	Trans.	Perm.	Trans.
Setting α_c and α_e to their estimated values								
1.....	0.988	0.012	0.275	0.725	0.696	0.304	0.978	0.022
2.....	0.983	0.017	0.359	0.641	0.778	0.222	0.985	0.015
3.....	0.983	0.017	0.388	0.612	0.855	0.145	0.989	0.011
4.....	0.987	0.013	0.451	0.549	0.896	0.104	0.992	0.008
5.....	0.990	0.010	0.518	0.482	0.921	0.079	0.993	0.007
6.....	0.993	0.007	0.594	0.406	0.936	0.064	0.995	0.005
7.....	0.994	0.006	0.666	0.334	0.947	0.053	0.995	0.005
8.....	0.995	0.005	0.727	0.273	0.954	0.046	0.996	0.004
9.....	0.996	0.004	0.777	0.223	0.960	0.040	0.996	0.004
10.....	0.996	0.004	0.816	0.184	0.964	0.036	0.997	0.003
11.....	0.997	0.003	0.846	0.154	0.967	0.033	0.997	0.003
12.....	0.997	0.003	0.869	0.131	0.970	0.030	0.997	0.003
24.....	0.999	0.001	0.958	0.044	0.984	0.016	0.999	0.001
36.....	0.999	0.001	0.975	0.025	0.989	0.011	0.999	0.001
Setting $\alpha_c = \alpha_{wnh} = 0$								
1.....	1.000	0.000	0.331	0.669	0.704	0.296	1.000	0.000
2.....	0.999	0.001	0.429	0.571	0.775	0.225	1.000	0.000
3.....	0.998	0.002	0.449	0.551	0.848	0.152	0.996	0.004
4.....	0.999	0.001	0.503	0.497	0.890	0.110	0.995	0.005
5.....	0.999	0.001	0.561	0.439	0.917	0.083	0.996	0.004
6.....	0.999	0.001	0.628	0.372	0.934	0.066	0.996	0.004
7.....	0.999	0.001	0.691	0.309	0.945	0.055	0.996	0.004
8.....	0.999	0.001	0.746	0.254	0.953	0.047	0.997	0.003
9.....	0.999	0.001	0.790	0.210	0.959	0.041	0.997	0.003
10.....	0.999	0.001	0.825	0.175	0.964	0.036	0.997	0.003
11.....	1.000	0.000	0.852	0.148	0.967	0.033	0.998	0.002
12.....	1.000	0.000	0.873	0.127	0.970	0.030	0.998	0.002
24.....	1.000	0.000	0.956	0.044	0.985	0.015	0.999	0.001
36.....	1.000	0.000	0.973	0.027	0.990	0.010	0.999	0.001

(*) Shares might not sum to unity because of their rounding.

[28] See the system in column C, Table 2.

Actually from Table 3 it emerges that permanent shocks are responsible for almost the entire variance in non-housing net worth of Italian households at all horizons, the picture being not much different for residential wealth, apart from the very first quarters when $\widetilde{\eta}_t^T$ is not negligible. Transitory innovation effects take time to elapse for disposable income, driving three fifths of its variability on the average of the first year and one fourth after eight quarters, with overwhelming permanent shocks in the longer run.

These overall features hold substantially true also when we focus on a single measure of total wealth, for which transitory innovation effects almost fully disappear after a few quarters (Table 4).

Table 4

Forecast error variance decomposition – total wealth (*)
(orthogonalized VAR residuals)

	Δct_{t+h}-$E_t\Delta ct_{t+h}$		Δy_{t+h}-$E_t\Delta y_{t+h}$		Δy_{t+h}-$E_t\Delta y_{t+h}$	
Horizon	Perm.	Trans.	Perm.	Trans.	Perm.	Trans.
Setting αc and αe to their estimated values						
1	0.974	0.026	0.187	0.813	0.715	0.285
2	0.973	0.027	0.195	0.805	0.792	0.208
3	0.975	0.025	0.224	0.776	0.850	0.150
4	0.978	0.022	0.277	0.723	0.883	0.117
5	0.981	0.019	0.351	0.649	0.904	0.096
6	0.984	0.016	0.441	0.559	0.918	0.082
7	0.987	0.013	0.530	0.470	0.928	0.072
8	0.988	0.012	0.608	0.392	0.936	0.064
9	0.990	0.010	0.672	0.328	0.943	0.057
10	0.991	0.009	0.724	0.276	0.948	0.052
11	0.992	0.008	0.765	0.235	0.952	0.048
12	0.993	0.007	0.797	0.203	0.956	0.044
24	0.997	0.003	0.934	0.066	0.978	0.022
36	0.998	0.002	0.963	0.037	0.985	0.015
Setting αc= 0						
1	1.000	0.000	0.207	0.793	0.713	0.287
2	0.999	0.001	0.235	0.765	0.789	0.211
3	0.997	0.003	0.259	0.741	0.849	0.151
4	0.996	0.004	0.319	0.681	0.885	0.115
5	0.997	0.003	0.397	0.603	0.907	0.093
6	0.997	0.003	0.488	0.512	0.923	0.077
7	0.997	0.003	0.574	0.426	0.934	0.066
8	0.998	0.002	0.646	0.354	0.942	0.058
9	0.998	0.002	0.703	0.297	0.949	0.051
10	0.998	0.002	0.748	0.252	0.954	0.046
11	0.998	0.002	0.783	0.217	0.958	0.042
12	0.998	0.002	0.811	0.189	0.962	0.038
24	0.999	0.001	0.932	0.068	0.981	0.019
36	1.000	0.000	0.959	0.041	0.987	0.013

(*) Shares might not sum to unity because of their rounding.

It turns out that the cointegrating coefficients reported in Tables 2 and 3 are very informative as to the relation between Italian households' spending and wealth, both their movements being soon dominated by the common trend they share.

Some more caution should be used when reading the coefficient of income, whose reaction to transitory innovation is not trivial. Let's consider, for instance, the average

variance decomposition of the first eight quarters: since 51 percent of the movements in income are transitory, then only the remaining 49 percent will be associated with the marginal propensity to consume of 70 cents estimated in section 7.3. In the longer run, however, the overwhelming role played by permanent shocks confirms that we can consider the income MPC as actually informative.

Decomposition results are broadly in line with those already available in the international applied literature, the major difference being the bigger role played by permanent innovations in the dynamics of Italian non-housing net worth. In particular, Lettau and Ludvigson (2004) find that in the US stock market wealth – just a component of the broader financial aggregate we used for Italy – is dominated by transitory shocks, which explain half of the variance of US financial assets also in Kundan Kishor (2006).

According to Pichette and Tremblay (2003), this share declines to about 30 per cent for Canadian stock wealth. The heterogeneity of results can be brought back to the composition of Italian households' financial wealth, in which money and deposits – presumably particularly sensitive to permanent shocks – have traditionally played a relevant role in families' portfolios, while stocks have represented a minor share.

9. Conclusions

Since the strong upsurge of real estate prices in the late nineties, renewed interest has emerged on the linkage between households' consumption, income and wealth. More specifically, analysts and economists frequently ask questions about the significance and magnitude of the marginal propensities to consume out of financial and, above all, housing wealth. Despite the increasing interest, the aggregate evidence on this issue was virtually nil for the Italian economy. Based on fresh estimates on households' wealth, we find sound evidence in favour of the existence of a cointegrating relation between consumer spending and different stock components, with a positive wealth effect in the long run. In the vein of Lettau and Ludvigson (2004), we enriched the research by investigating the role of transitory and permanent shocks in the variables considered. We found that consumption, housing and non housing wealth respond almost exclusively to permanent shocks; the same shocks play an overwhelming role also for disposable income over a long horizon, whereas in the short run the effects of transitory shocks are not negligible. As a result, we estimate a marginal propensity to consume out of housing and non housing wealth in the range of, respectively, 1.5-2 and 4-6 cents, which we may consider to closely match the true values in the long run equilibrium.

APPENDIX I – UNIT ROOT TEST RESULTS

Table I-1

Unit root test results (*)

(variables in logs and in real terms)

	Total consumption (ct_t)	Housing wealth (wh_t)	Non housing wealth (wnh_t)	Housing wealth net of mortgages (wh_a_t)	Non housing wealth gross of mortgages (wnh_a_t)	Total wealth ($wtot_t$)	Disposable income (y_t)	Interest rate (r_t)
	Augmented Dickey-Fuller Test – H_0: series has a unit root							
Variables in level.....	-1.800* (0.38)	-0.638* (0.86)	-1.344* (0.61)	0.095* (0.96)	-1.176* (0.68)	-1.489* (0.536)	-1.7818* (0.39)	-0.958* (0.77)
1st differences.	-3.560 (0.00)	-4.13 (0.00)	-9.023 (0.00)	-0.348 (0.00)	-8.689 (0.00)	-6.410 (0.00)	-3.773 (0.000)	-5.99 (0.00)
	Phillips-Perron Test – H_0: series has a unit root							
Variable in levels.....	-1.456* (0.55)	-1.297* (0.63)	-1.322* (0.62)	-1.435* (0.56)	-1.149* (0.69)	-1.302* (0.63)	-3.763 (0.004)	-0.798* (0.82)
1st differences.	-7.09 (0.00)	-5.232 (0.00)	-9.547 (0.00)	-5.310 (0.00)	-9.485 (0.00)	-6.51 (0.00)	-7.489 (0.00)	-6.05 (0.00)
	ERS(1) Test – H_0: series has a unit root							
Variable in levels.....	1045*	57.47*	351.7*	52.49*	407.0*	477.0*	454.2*	122.40*
1st differences.	1.35	2.16	0.49	2.17	0.49	1.02	2.69	0.68
	KPSS(2) Test – H_0: series is stationary							
Variable in levels.....	1.142	1.065	1.141	1.045	1.148	1.171	0.982	1.142
1st differences.	0.239*	0.072*	0.20*	0.07*	0.16*	0.10*	0.61	0.086*

(*) When reported, *p*-values in brackets. (*): H_0 accepted at 95% confidence level; (1) Elliot, Rothenberg and Stock point optimal test; (2) Kwiatkowski, Phillips, Schmidt and Shin test.

APPENDIX II – COINTEGRATION COEFFICIENTS

Table II-1

Cointegration coefficients in VARs without deterministic controls (*)

(normalized on the coefficient of total consumption ct)

	System specification				
	(A)	(B)	(C)	(D)	(E)
	Cointegrating coefficients (β-vectors)				
Total wealth ($wtot_t$)	-0.360 (-11.245)	-0.227 (-5.850)			
Housing wealth (wh_t)			-0.092 (-5.815)	-0.064 (-5.601)	
Non Housing wealth (wnh_t)			-0.285 (-15.809)	-0.244 (-10.974)	
Housing wealth net of mortgages (wh_a_t)					-0.069 (-4.999)
Non housing wealth gross of mortgages (wnh_a_t)					-0.257 (-9.193)
Interest rate (r_t)		0.037 (3.447)		0.019 (3.024)	0.010 (1.326)
Disposable income (y_t)	-0.746 (-6.993)	-0.729 (-10.036)	-0.429 (-7.855)	-0.444 (-10.930)	-0.452 (-9.223)
	Error correction loadings (α)				
Total consumption (ct_t)	0.015 (0.620)	-0.036 (-1.058)	-0.029 (-0.612)	-0.144 (-2.284)	-0.090 (-1.562)
Total wealth ($wtot_t$)	0.233 (3.351)	0.268 (2.728)			
Housing wealth (wh_t)			0.667 (3.724)	0.825 (3.498)	
Non housing wealth (wnh_t)			0.517 (2.734)	0.401 (1.523)	
Housing wealth net of mortgages (wh_a_t)					0.698 (3.169)
Non housing wealth gross of mortgages (wnh_a_t)					0.457 (2.019)
Disposable income (y_t)	0.130 (5.044)	0.175 (4.722)	0.230 (4.097)	0.276 (3.600)	0.263 (3.812)

(*) *t*-statistics in parenthesis.

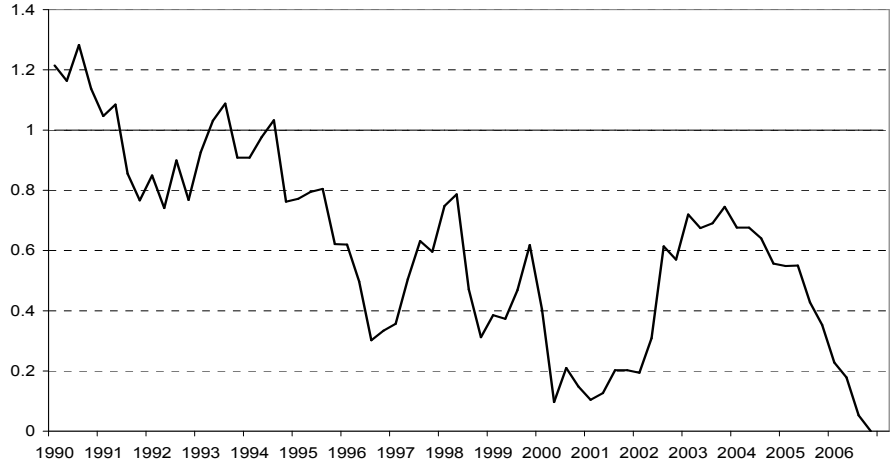
APPENDIX III – RECURSIVE TESTS OF PARAMETER CONSTANCY

The recursive tests are based on the concentrated version of the VECM: $R_{0t} = \alpha\beta' R_{1t} + v_t$, where β is the cointegrating vector, α contains the loadings of the error correction term and the rank of cointegration is set equal to one.[29] The advantage of this approach is that it averages out the short-run structure of the VECM, leaving the sole adjustment towards the long-run equilibrium. It is precisely on the parameters involved by this equilibrium (i.e. α and β) that we focus the testing procedure. We run both forward and backward recursive estimations in order to test the parameters constancy at the end and at the beginning of the sample. As for forward recursions, the baseline sample we chose is 1980q1-1990q1, whereas for the backward exercise it is 1998q4-2006q4.

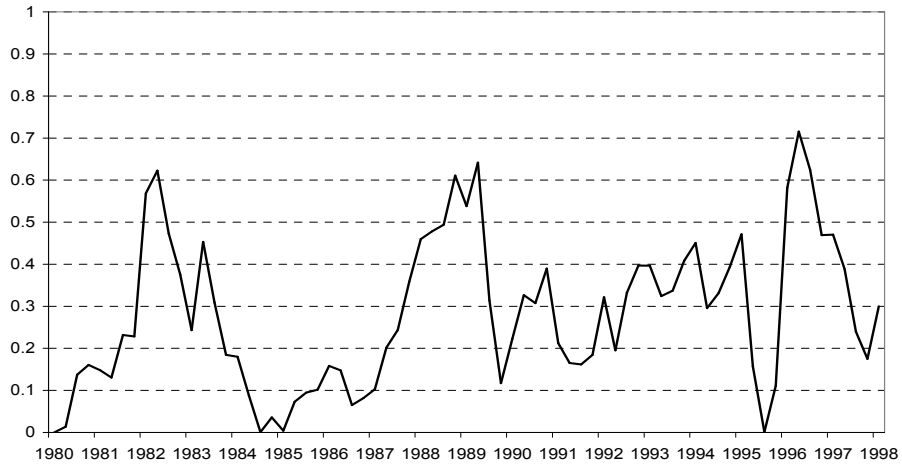
Test for constancy of the log-likelihood

(H_0: constant parameters)

Forward



Backward

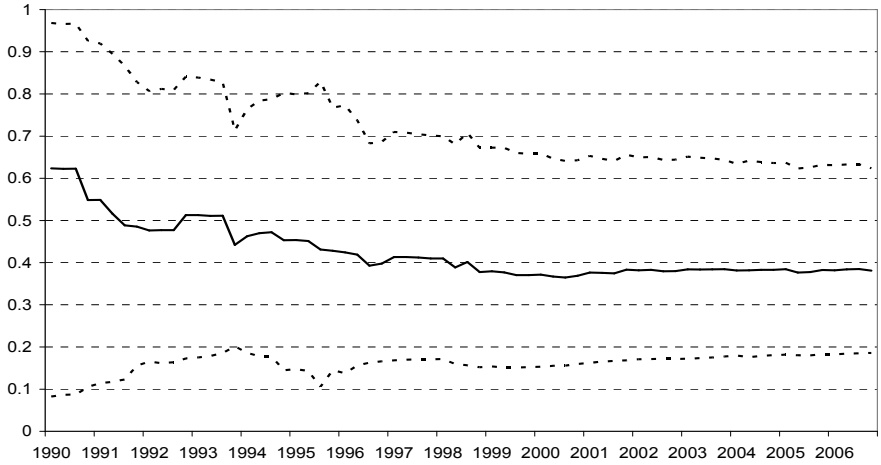


Note: the test statistic has been scaled by the 95% quantile of the appropriate asymptotic distribution; thus a test value less than 1 implies acceptance of H_0.

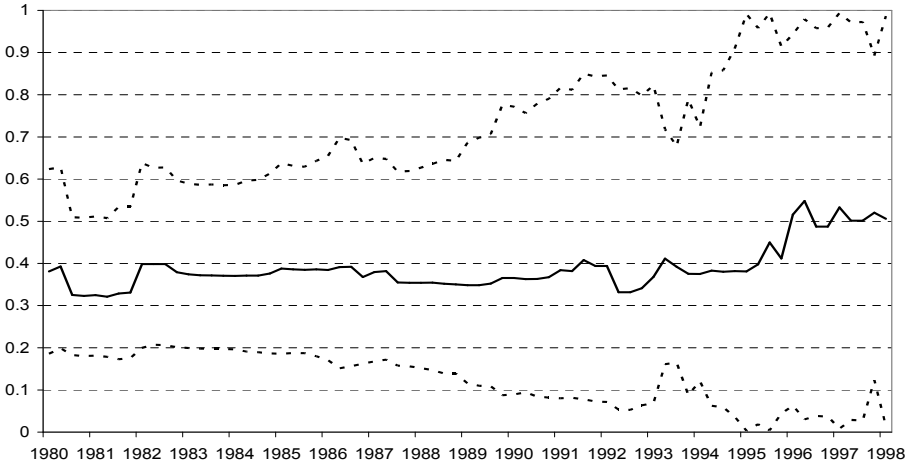
[29] The concentrated version can be obtained starting from the VECM $\Delta X_t = \alpha\beta' X_{t-1} + \Sigma_k \Gamma_k \Delta X_{t-k} + \Phi D_t + \varepsilon_t$ and rewriting it more compactly: $Z_{0t} = \alpha\beta' Z_{1t} + \Psi Z_{2t} + \varepsilon_t$, with $Z_{0t} = \Delta X_t$, $Z_{1t} = X_{t-1}$, $Z_{2t} = [\Delta X_{t-1}, \Delta X_{t-2}, \ldots, \Delta X_{t-k}, D_t]$ and $\Psi = [\Gamma_1, \Gamma_2, \ldots, \Gamma_k, \Phi]$. Then define the auxiliary regressions: $Z_{0t} = B_1' Z_{2t} + R_{0t}$ and $Z_{1t} = B_2' Z_{2t} + R_{1t}$, with $B_1' = M_{02} M_{22}^{-1}$ and $B_2' = M_{12} M_{22}^{-1}$, $M_{ij} = \Sigma_t (Z_{it} Z_{jt}')/T$. It follows that the concentrated model is $R_{0t} = \alpha\beta' R_{1t} + v_t$.

Recursively calculated eigenvalue λ_1

Forward



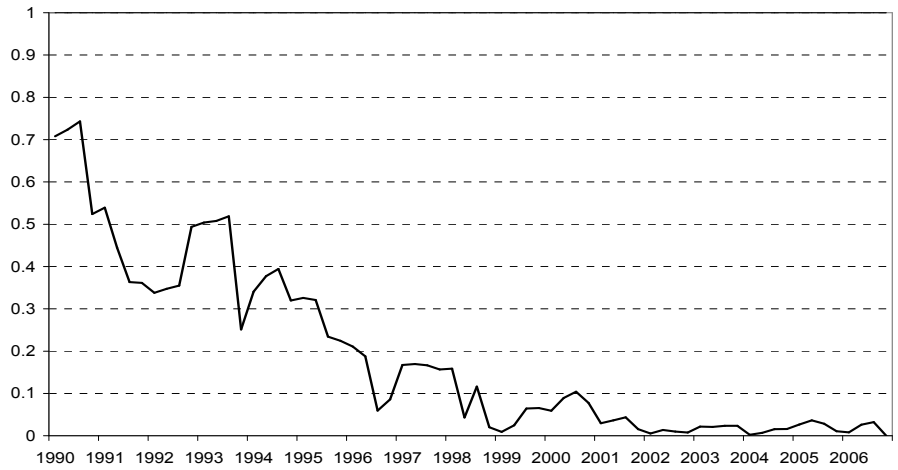
Backward



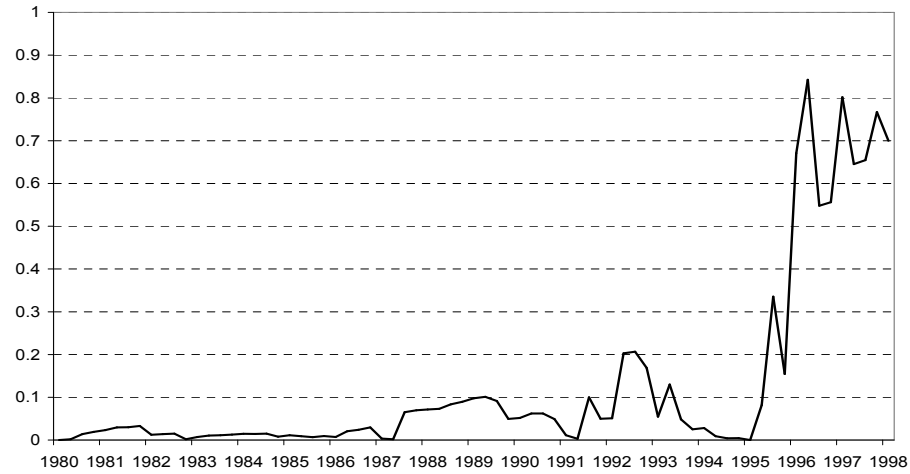
Note: dotted lines correspond to the 95% confidence bands.

Eigenvalue fluctuation test
(H_0: constant eigenvalue λ_1)

Forward



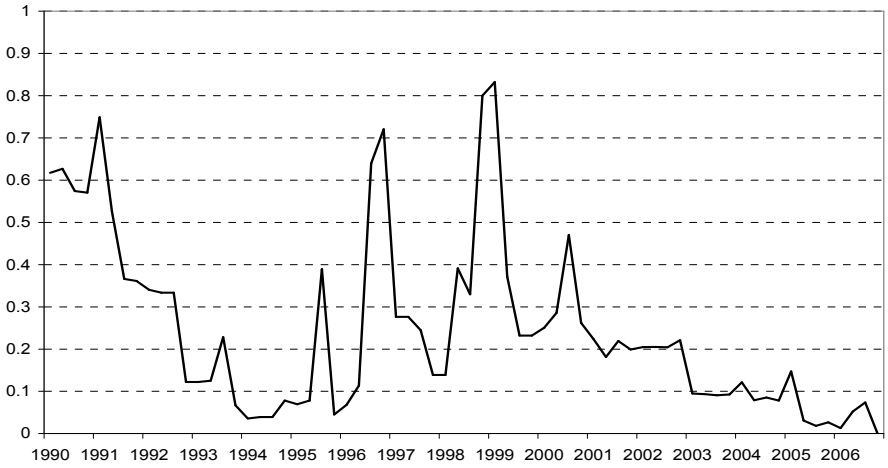
Backward



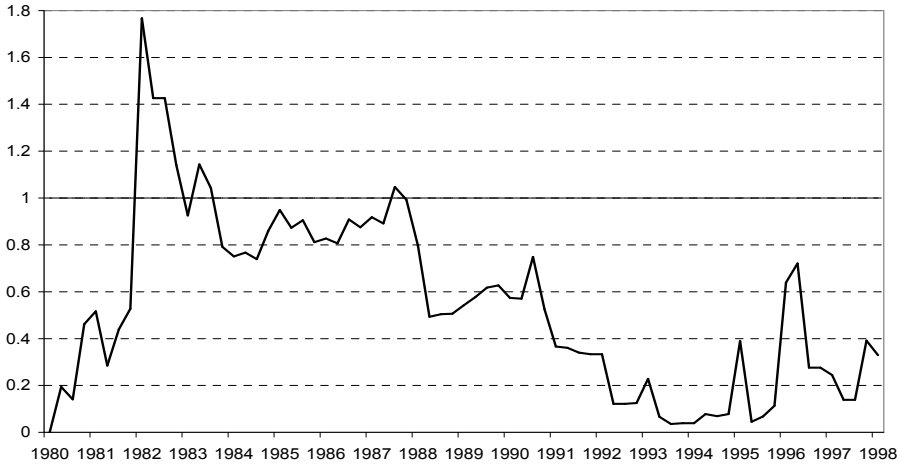
Note: the test statistic has been scaled by the 95% quantile of the appropriate asymptotic distribution; thus a test value less than 1 implies acceptance of H_0. This test can be considered a recursive constancy check of α and β.

Test for β constancy
(H_0: constant cointegrating coefficients)

Forward



Backward



Note: the test statistic has been scaled by the 95% quantile of the appropriate asymptotic distribution; thus a test value less than 1 implies acceptance of H_0.

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