BANCA D'ITALIA

Temi di discussione

del Servizio Studi

Personal Saving and Social Security in Italy: Fresh Evidence from a Time Series Analysis

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Number 417 - August 2001

PERSONAL SAVING AND SOCIAL SECURITY IN ITALY: FRESH EVIDENCE FROM A TIME SERIES ANALYSIS

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Abstract

The paper provides an econometric analysis of the aggregate saving function of Italian households in the vein of the life cycle theory. Results from an ECM representation based on yearly data for 1951-1998 point to depressive effects on private consumption of recent reforms of social security, actual and expected for next few years. In order to compensate for both reductions in actual pension payments and increased uncertainty about their future claims, households stepped up accumulation of real and financial assets since the beginning of the nineties. First estimates of capital gains do not show a significant impact on consumption demand, in the short and in the long period: their high volatility has likely hindered a fair assessment of their contribution to personal purchasing power on the part of households. Demographic changes, while in the long run not seemingly determined in conjunction with the economic variables we consider, turn out to play a significant role in the evolution of consumption demand.

JEL classification: E21, E62, H31, H55.

Keywords: life cycle, saving, pension wealth, capital gains, demographic changes.

Contents

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1. Introduction[1]

The effect of public transfers, social security in particular, on private wealth accumulation has long been the subject of a large body of literature, both theoretical and empirical. Comprehensive analysis is complex, but the prevailing approach has been centered upon the idea, espoused by the life cycle model since the seminal contribution of Modigliani-Brumberg (1954), that the need to finance consumption after retirement is the main motivation for personal saving. The major implication was that mandatory contributions to social security would fully offset households' voluntary savings, with a potential impact on social welfare[2]. This result is soundly proved in an overlapping generation framework under certainty, rigid factor supply and negligible altruism; it is fraught with difficulties in a more general set-up. In this respect, the main drawbacks stressed in the literature may be summarized in the following points, which are stated under the assumption of a pay-as-you-go social security scheme:

i) if current generations feel altruistic with their offspring, who are eventually called upon to finance the current payouts, the introduction of the social security system may lead to increased private saving in order to augment intergenerational transfers (Barro, 1978);

ii) elastic labor supply would imply the establishment of a social security system to induce earlier retirement, therefore leading people to increase savings in working age in order to finance consumption over a longer retirement period (Feldstein, 1974 and Munnel,

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[1] The opinions expresses in the paper do not engage the Bank of Italy. The author gratefully acknowledges key contributions by Andrea Brandolini and Salvatore Muzzicato in the computation of historic time series of disposable income, real wealth and market price of dwellings. He thanks Sergio Nicoletti and Ivan Faiella for providing micro estimates of social security wealth, Roberto Violi for a capitalization index of financial assets, and two anonymous referees for very helpful comments. Elena Picchio offered valuable editorial assistance. Any errors are the author's alone: zollino.francesco@insedia.interbusiness.it.

[2] An extensive welfare analysis is not in the reach of the paper (for references and discussion, see Blanchard and Fisher, 1989 and Diamond, 1997). It is important to recall, however, that a key role is played by the magnitude of the corresponding change in national saving, for which the funding status of the social security system would matter. The expected sign of this change would be negative under a pay-as-you-go scheme, since the reduction in personal saving would not be offset, *ceteris paribus*, by an increase in public saving; the opposite holds true under a fully funded scheme. But this point is controversial if bequest motives are operative (for a full appraisal, see Seater, 1993).

1974). Moreover, if contributions and benefits are imperfectly linked, the resulting changes in effective tax on labor could affect its supply, which would in turn impact on personal saving (Feldstein-Samwick, 1992);

iii) credit market imperfections reduce the relevance of the life cycle motive for saving, with borrowing constraints limiting the extent to which social security crowds out private savings (Diamond-Hausman, 1984 and Dicks Mireaux-King, 1984). The same holds true when annuity market imperfections prevent a fair assessment of the wealth effects related to social security (Bernheim, 1987);

iv) under uncertain longevity and income, an additional reason why social security may induce lower personal saving is a diminished motive for precautionary saving (Kotlikoff *et* al. 1987 and Hubbard *et* al. 1995). However, this effect would prove smaller if uncertainty surrounded both the financial sustainability of the social security system itself and its impact on the households' economic status in retirement (Carroll-Samwick, 1992 and Bernheim, 1995);

v) lack of economic literacy and mental accounts for different assets may limit the extent of offset between pension and non-pension wealth (Bernheim, 1997 and Thaler, 1994); a similar prediction comes from hyperbolic discounting (Laibson, 1996), and in this case social security would prove to be a commitment technology for workers to raise enough saving for their own retirement. This would reinforce the psychological argument first put forth by Katona (1965), whereby social security may increase personal saving by inducing a higher preference for saving on the part of otherwise very impatient households[3].

In view of the inconclusive results of theoretical models, empirical analysis has gained a key role in predicting the impact of social security on private savings. However, disagreement arises in this field too, with a variety of conclusions coming from different data set and econometric methods since the first contribution by Feldstein (1974). In line with the

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[3] While deviation from rational behavior with foresight augments complexity in theoretical models of social security and saving, some irrationality in households' decisions helps to explain the very existence of a social security system in modern economies (Kotlikoff, 1987). The evidence that a substantial part of households reach retirement with a very low stock of net financial assets (Poterba *et* al., 1994) may provide some support.

theoretical debate, attention has focused on the size of the offset between pension and non-pension wealth in consumption demand. The results have shown great variability, with high sensitivity to measurement errors of social security wealth and to the aggregation level in data (Modigliani, 1986, Castellino-Fornero, 1990 and Engel-Gale, 1999).

With respect to the Italian economy, the empirical literature was started by Brugiavini (1987) and built up over the nineties, in view of an increasing debate on the need for major reform of the domestic social security system. Like the evidence collected in other countries, a lack of consensus has arisen as to the size of the replacement effect between pension and other wealth for Italian households. In Rossi-Visco (1995) and Beltrametti-Croce (1997), aggregate time series analysis shows high values of the offset between pension and non-pension wealth, with estimates around respectively 0.7 and 1.0 for the period 1954-1993. This value does not exceed 0.2 in Brugiavini (1991) and Jappelli (1995) as a result of cross section analysis based on the micro data provided by the Bank of Italy Survey on Household Income and Wealth (SHIW, respectively for the years 1985 and 1989-91). Finally, Favero *et* al. (1994) rejected any significant role of pensions in determining consumers' behaviour in Italy, on the basis of the same aggregate data set as in Rossi-Visco but following a VAR specification of the statistical model.

In this paper we adopt the time series approach followed in Rossi-Visco, with the purpose of providing fresh evidence on the effects on personal savings of the social security reforms passed in 1992 and 1995 (Bank of Italy, 1995). The main motivation is to investigate the extent to which the changed set-up of social security, both actual and expected, explains the prolonged weakness of private consumption, following the financial crisis of the early nineties. In addition, the recent, sizeable switch in Italian households' portfolios from Treasury paper to equities and investment funds provides the opportunity to test the importance of capital gains in affecting aggregate consumption and saving behavior. The intensity of the latter's response to major demographic trends is a further issue addressed in the paper.

In view of the aggregate level of our empirical analysis, a major subject we neglect is the size of distributive effects that may stem from both pension reforms and changes in portfolio allocation, and their impact on consumption demand. Future research to validate our time series evidence on the basis of micro data would be well worthwhile.

The paper is organized in seven remaining sections. In sections 2 and 3 we briefly comment on measurement issues and recent trends in the major variables we consider in the analysis of consumption demand. In the fourth we outline the empirical model, which we estimate in an error-correction representation in the fifth section. In the sixth a plausible interpretation of the empirical findings is suggested and in the following section structural changes in households' expenditure are detected in view of the recent reforms of Italian pension system. Section 8 concludes, summarizing the main results.

2. The data set

Substantial statistical work was required to identify and estimate measures for some key variables covering the whole sample period from 1951 to 1998. In this section we briefly address some measurement issues; more details about computational procedures are reported in the appendix.

The main difficulties involve social security wealth. The measure first computed by Beltrametti (1995) for the years 1951-1993 has been revised and updated from 1989 up to 1998 on the basis of preliminary estimates from a micro-model of Italian economy as in Ando-Nicoletti (1999) and Ando et al. (1999)[4]. The main rationale is that microanalysis enables us to take full account of the diversified effects, across cohorts and occupations, of the major reforms of the pension system in 1992 and 1995. Moreover, it may incorporate the ensuing revision in households' expectations affecting key choices, especially the age of retirement. In this respect, the growing debate on the sustainability of the social security system since the later eighties might have changed Italian households' expectations about future benefits, and then their decision as to labor market participation, before the actual passage of the reforms[5,6]. Indeed, retirement age is considered to be determined by law in the

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[4] In both cases the data set comes from the Bank of Italy survey of household income and wealth (SHIW).

[5] As evidence of the early debate about the need for pension reform, the failure to get it passed is acknowledged by Guido Carli as a major defeat as Minister of the Treasury between 1989 and 1992 (Carli, 1993).

[6] A decline in the expected ratio of benefits to income at retirement was registered in the 1989 and 1991 waves of the SHIW, which explicitly polled Italian households on this subject (Bank of Italy, 1989 and 1991).

computations of Beltrametti, and it may endogenously prove to be higher in the micro model estimates. It is mainly this feature that explains why the latter show a declining trend before 1992, when the so-called "Amato reform" took eventually place. In advocating the need to take the variety of factors working at a micro level into account - especially in the recent institutional set-up, featuring a stratification of rules for computing pension benefits - we nevertheless acknowledge some arbitrariness in the specific measures we computed[7]. In any case, our main claim focuses on the *timing* of the new regime in expected pensions, which, as far as consumption behavior is concerned, may have started prior to the reforms themselves, exerting effects on households' expenditure that might be empirically important.

Capital gains have been estimated separately for dwellings, bonds and equities. For each component, the magnitude of revaluation has been computed on a yearly basis, by applying the changes of a price index each year to the value of the relevant stock at the end of previous year. An alternative measure, given by the difference between total changes in the value of stock and the "out of pocket" saving (Horioka, 1996), was dismissed because of the resulting extreme volatility in capital gains of Italian households, much larger than that which we discuss in the next section.

In respect to the coverage of the household sector and the definition of both private consumption and some components of disposable income, the accounting scheme used is ESA79, since time series consistent with ESA95 were not available when we completed the study. Therefore, we do not take into account the major revisions of some key macro-variables effected in the last few years (Istat, 1998 and Bank of Italy, 1999).

3. Recent trends in some key macro-variables

According to the ESA79 national accounts, Italian households' propensity to save fluctuated around the highest found in all industrial countries until the eighties. Since then, it

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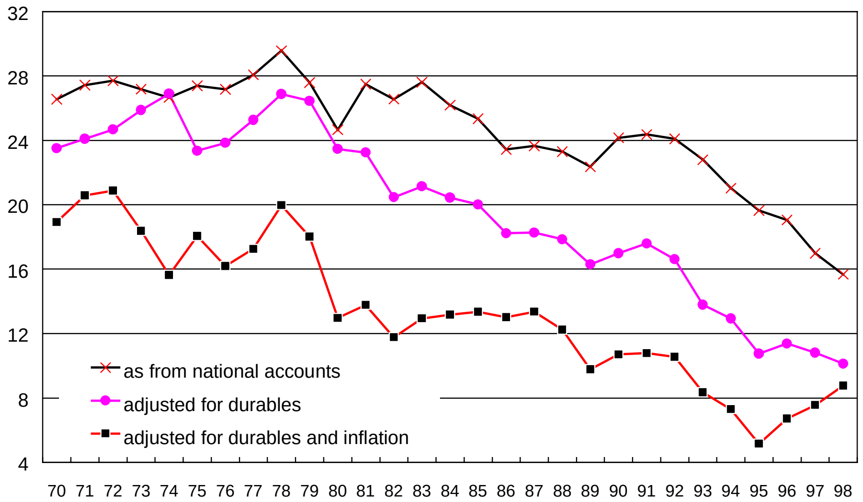
According to Jappelli (1995), between 1989 and 1991 the expected ratio declined on average by four percentage points, to a value around 0.75; interestingly, he also found a negative change in the actual ratio.

[7] We discuss this issue in the appendix. More generally, endemic problems limiting the quality of measures of pension wealth within the empirical analysis of consumption are addressed in Leimer-Lesnoy (1982) and Gale (1998).

has shown a declining trend, more pronounced in the nineties. In fifteen years, it fell by more than 12 percentage points, to under 12 per cent in 1998 (Fig. 1)[8].

Fig. 1

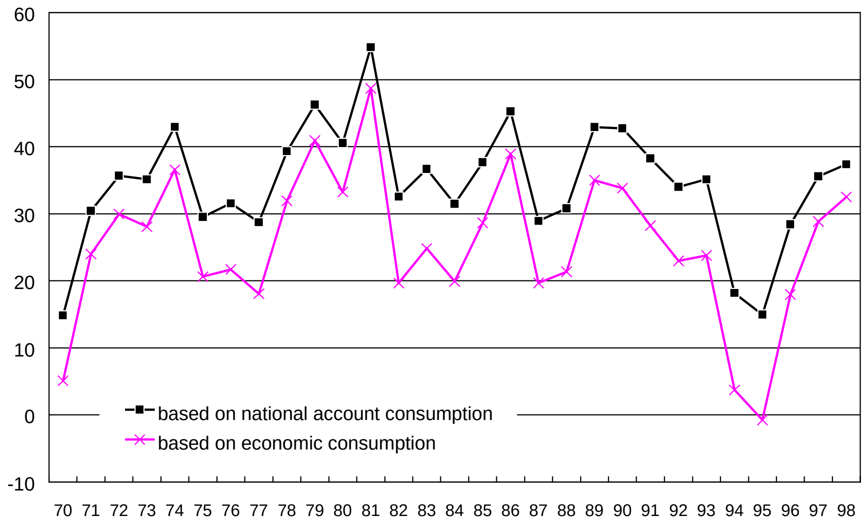
HOUSEHOLD SAVING RATE
(percentage values)



Source: Based on Istat data.

Fig. 2

HOUSEHOLD SAVING RATE
(adjusted for inflation and capital gains; percentage values)



Source: See Appendix.

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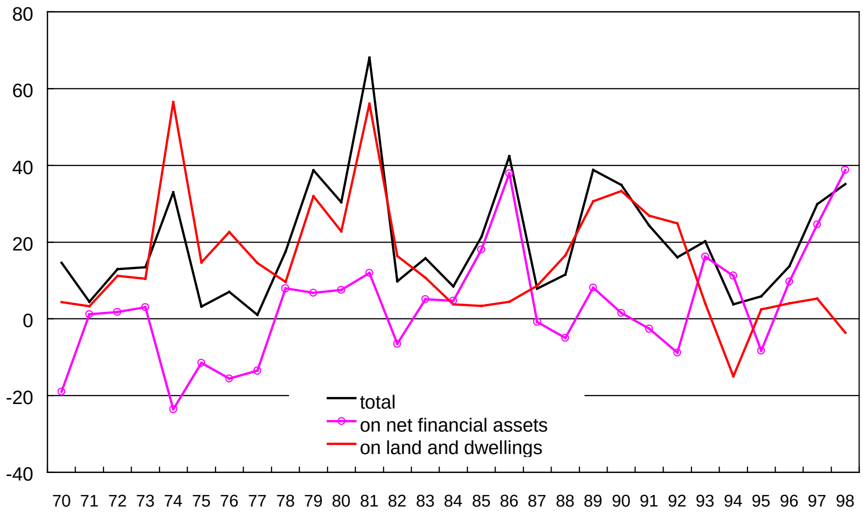
[8] According to preliminary estimates based on ESA95, between 1983 and 1999 households' propensity to save dropped from above 28 per cent to 14.2 per cent (Bank of Italy, 2000a).

Taking account of durables and the loss in purchasing power of net assets due to expected inflation, the decline in the propensity to save began earlier, in the late seventies, and came to a halt in the second half of the nineties: since 1995 there is a clear upward trend.

Fig. 3

RATIO OF HOUSEHOLD CAPITAL GAINS TO DISPOSABLE INCOME
(percentage values)



Source: See Appendix.

In recent years the widening gap between the adjusted and unadjusted measures of the saving rate is due mainly to the rapid decline in inflation and, to a lesser extent, to the fall in long-term interest rates (Fig. 5), which affect the computation of the durable goods' usage costs. Saving rates show a considerable volatility when the measure of disposable income includes capital gains, closely tracking the latter's erratic developments (Fig. 2).

This symmetric movement indicates a rather low propensity to consume out of capital gains, arguably because of the great uncertainty over their magnitude. Between 1994 and 1998 the ratio of capital gains to the national accounts' measure of disposable income[9] climbed by 30 percentage points, after a drop of 10 percentage points in 1993 (40 in the previous five years; Fig. 3).

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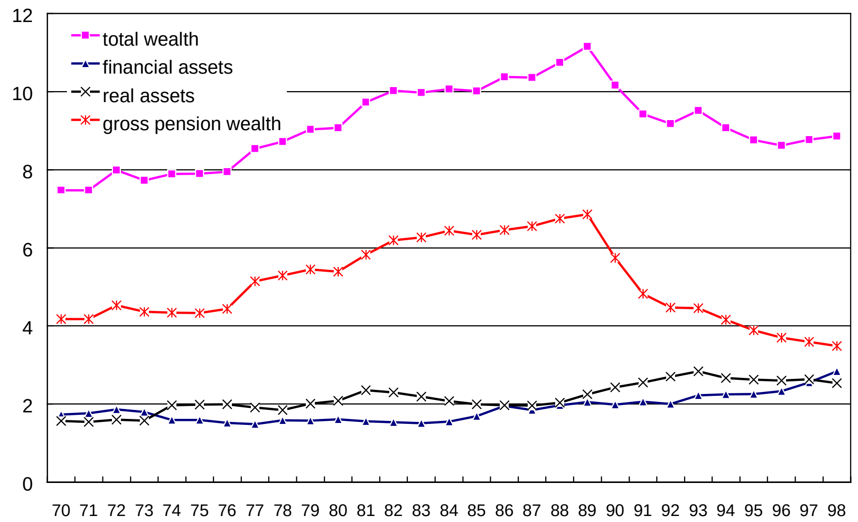
[9] In the old ESA79 capital gains are recorded in the capital accounts, in the subsection of revaluation of net assets, and do not enter in the computation of income. In the new ESA95 system, the share of capital gains coming form capitalization of interest and dividend payments contribute to disposable income.

Moreover the composition of total capital gains changes dramatically: the share of real estate falls virtually to zero in the last five years after accounting for nearly all of them over the previous decade. This may have played a role in increasing uncertainty over capital gains, checking households' propensity to consume out of them[10].

Fig. 4

RATIO OF HOUSEHOLDS' WEALTH TO DISPOSABLE INCOME



Source: See Appendix.

The ratio of total wealth to income rose until the late eighties, mainly driven by the social security component. Since then, the clear decline in this component has reversed the tendency of the overall ratio, offsetting the increased accumulation of net financial assets (Fig. 4). Going by the measures we adopted, social security wealth declined, in levels, by around 25 per cent over the decade of the nineties (Fig. A1 in Appendix). Noticeable, the reduction in the ratio of pension wealth to disposable income between 1989 and 1991 is amplified by the acceleration of the latter variable, which registered average annual growth of around 4 per cent in that period (Fig. 5).

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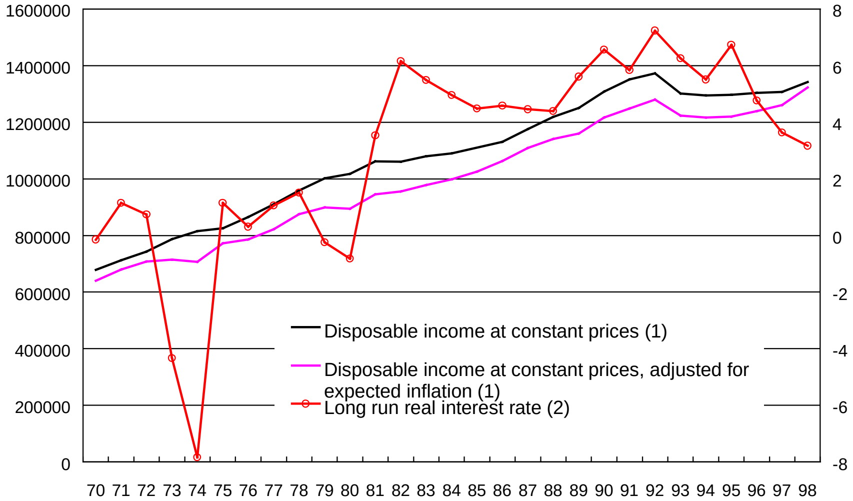
[10] An additional factor limiting the impact of capital gains on aggregate consumption may relate to distributive issues, in view of the greater concentration in holdings of equities and investment funds than of government bonds among Italian households (Bank of Italy, 2000b).

In the presence of positive stock-flow adjustments following the sharp revaluation of equities and bonds, between 1992 and 1998 the ratio of financial assets to disposable income increased by one percentage point, to 3 per cent. In the same period, the incidence of the stock of real assets, which includes durable goods, registered a slight reduction.

Fig. 5

HOUSEHOLD DISPOSABLE INCOME AND REAL INTEREST RATES
(Lira at 1995 prices and percentage values)



Source: See Appendix.

(1) Left hand scale (2) Right hand scale.

Disposable income, adjusted to account for losses in purchasing power of net assets due to expected inflation, did not recover the huge drop registered during the recession of 1992-93 until the late nineties. The average for the decade was thus about the same as in the late eighties.

4. The empirical model

In this section, we sketch the theoretical foundations of the empirical model, which hinges on the standard life cycle approach extended to cover social security effects (Feldstein, 1974). Following the seminal contribution of Modigliani-Brumberg (1954), under conditions of regularity in the preferences of household groups, the aggregate consumption function may be linearized as follows

(1) $$c_t = \alpha y_t^l + \delta w_t ,$$

where c_t is the flow of goods and services consumed in period t, y_t^l is labor income after taxes (as approximation of the annuity coming from human capital) and w_t is the stock of wealth at the end of period. Parameters α and δ generally depend on a variety of factors, notably the age composition of the population and the length of different stages in life (Ando-Modigliani, 1963). In order to discuss the role of social security empirically, total wealth is split into net real and financial assets (w_t^r) and a social security component (w_t^s), namely the expected discounted flow of pension payments accruing to both current and future retirees. In particular, taking account of the possibility of imperfect substitution between the two components (Feldstein, 1974), the total wealth entering the consumption function is given by the sum

(2) $$w_t = w_t^r + \xi w_t^s$$

with ξ measuring the degree of substitutability between pension payments and the real and financial assets in sustaining households' purchasing power. In line with theory (Williamson-Jones, 1983 e Carroll-Samwick, 1992), values for this parameter significantly lower than one may signal uncertain expectations by households as to the future evolution of social security payments. But the value of this parameter is affected by the inclusion of current transfers in the measure of disposable income, which is computed gross of these transfers and net of social contributions according to the national accounts[11]. In order to avoid this problem of *double counting*, here we take social security wealth gross of social contributions and disposable income net of current transfers (y_t^s).

In view of the national accounts' definition of households' disposable income y_t^d as the sum of labor and capital incomes after taxes (including social contributions) and

[11] Whereas current transfers are included also in the computation of social security wealth, it is possible to show that the parameter ξ is expected to be higher than one (Williamson-Jones, 1983). An extensive discussion of the implications of this issue for the right specification of the statistical model is in Favero et al. (1994) and in Rossi-Visco (1995).

transfers, namely $y_t^d = y_t^l + r_t w_t^r + y_t^s \equiv y_t^p + y_t^s$, labor income may be defined by $y_t^l = y_t^p - r_t w_t^r$. From (1), gross saving is then given by the following

(3) $$s_t \equiv y_t^d - c_t = y_t^p + y_t^s - \alpha(y_t^p - r_t w_t^r) - \delta(w_t^r + \xi w_t^s)$$

From this, it results that

(4) $$-\log(c_t / y_t^p) \approx \frac{s_t - y_t^s}{y_t^p} = (1-\alpha) - (\delta - \alpha r_t)\omega_t^r - \delta\xi\omega_t^s,$$

where $\omega = w / y^p$.

Allowing for a flexible representation of the parameter δ as a function of the real interest rate *R* and the growth rate of national income *g* (Modigliani-Brumberg, 1980) and expanding around *steady state* values of ω^s, ω^r, *R* and *g* (Rossi-Visco, 1995), the empirical model reads

(5) $$\log(c_t) = const + \psi_o \log(y_t^p) + \psi_1 g_t + \psi_2 r_t + \psi_3(\omega_t^r + \vartheta\omega_t^s),$$

where $\vartheta(\alpha, \delta, r, g) > \xi$ approximates the coefficient of substitution between gross social security wealth and net real and financial assets.

In this setup the assumption of homogeneity in the *steady state* of consumption plans and total household resources, which is theoretically desirable for dealing with aggregation since it rules out scale factors in the individual household's optimal choices (Modigliani-Brumberg, 1954 e Ando-Modigliani, 1963), is realized with a value of one for the parameter ψ_0. With reference to an interpretation of the coefficients of the two wealth components, they are a combination of the propensities to consume out of the single component and income[12].

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[12] An immediate proof can be derived from a simplified version of the life cycle model $c_t = \alpha W_t + \beta y_t$, which implies $c_t = \beta y_t[1 + (\alpha/\beta)(W_t / y_t)]$. Taking logs, under the reasonable assumption of a small value for the term added to unity in the square bracket, we get a specification quite similar to that considered in the text, $\log c_t \approx \text{costant} + \log y_t + \gamma\omega_t$, with $\gamma = \alpha / \beta$. Accordingly, ruling out measurement errors in life-long income, the intercept would approximate the propensity to consume out of income in a logarithmic specification of households' demand. As to the wealth variables, their normalization with respect to income allows for

5. The estimated consumption function[13]

A preliminary stationarity analysis has been performed for the following variables, valued in real terms: a) logarithm of households' consumption expenditure adjusted for durables (LC); b) disposable income (LYP), net of pension payments and adjusted for losses of purchasing power of net financial assets due to expected inflation; c) ratio of real and financial assets to disposable income (WYP); d) ratio of social security wealth to disposable income (WSSY); e) long-term interest rate (R).

As from Tables A1a-c in Appendix, for variables LC, LYP and WYP, which all show a positive drift, the augmented Dickey-Fuller tests have definitely not rejected the null hypothesis of unit root at a confidence level of 95 per cent. The same result, although to a narrower extent, holds true for variable R, which does not follow a deterministic trend (Table A1d), and for variable WSSY (Table A1e). As to the latter, the Phillips-Perron (1988) test with semi-parametric correction has also been performed, confirming the non-rejection of unit root (table A1f)[14].

In view of the reduced power of the ADF test under structural change (Campos-Ericsson-Hendry, 1996), we have additionally explored non-stationarity of WSSY by controlling for changes in both level and growth rate. According to properly revised critical values (Model C in Perron, 1989), the null hypothesis of unit root cannot be rejected against the alternative of a break in the deterministic component, starting from 1989 (Table A1g).

Following the evidence of non-stationarity, we tested for co-integration among LC, LYP, WYP, WSSY and R to identify, according to the usual interpretation first put forward by Engle-Granger (1987), their relationship in the long run equilibrium. We follow the multivariate, maximum likelihood method of Johansen (1988), which we have applied to the full and restricted samples in order to test robustness of the cointegration rank. Consumption and income are considered in isolation in order not to impose the restriction of their

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controlling the effects of *common trend* and, as is pointed out in Muellbauer-Lattimore (1995), for properly decomposing the total impact of wealth on consumption in the distinct contributions coming from different components of wealth.

[13] The results reported in this section are mostly based on the Microfit 4.0 package (Pesaran-Pesaran, 1997).

[14] The *t*-statistics in this table compare with the critical values for ADF test in previous tables.

homogeneity, which has been tested separately. In line with stationary first differences of the five variables, all of which show a deterministic trend in levels, but the real interest rate R, a time trend in the long-run relationships has been excluded, with an intercept included. In this context, some conditioning variables have been considered, namely the first difference of the logarithm of government consumption (DZ) and the changes in population share of old people (65 and over, DPO)[15]. The former may summarize the effect of fiscal policy stance as in Rossi-Visco (1995), while the latter should take account of the demographic transformation, which has been more pronounced in the last two decades[16]. It is interesting to note that although the role of changes in demographic structure has been well understood in the life cycle model since Modigliani-Brumberg (1954), it is rarely controlled for explicitly in time series analysis of consumption[17].

Table 1

TEST STATISTICS AND CRITERIA FOR SELECTING THE ORDER OF THE VAR MODEL

Based on 44 observations from 1955 to 1998. Order of VAR = 4
List of variables included in the unrestricted VAR: LC LYP WYP WSSY
List of deterministic and/or exogenous variables: DPO DZ

Order	LL	AIC	SBC	LR test	Adjusted LR test
4	279.7819	207.7819	143.5511	------	------
3	266.6148	210.6148	160.6575	CHSQ(16)= 26.3342	15.5611
2	256.7761	216.7761	181.0923	CHSQ(32)= 46.0116	27.1887
1	247.5225	223.5225	202.1123	CHSQ(48)= 64.5188	38.1247
0	-167.3560	-175.3560	-182.4927	CHSQ(64)= 894.2758	528.4357

AIC=Akaike Information Criterion SBC=Schwarz Bayesian Criterion

[15] For both variables in level, the ADF test does not reject the hypothesis of unit root at a confidence level of 95 per cent. The same holds for life expectancy at birth (ASP).

[16] An additional demographic variable has also been considered, namely the first difference in life expectancy at birth (DASP), without any appreciable changes in the results; moreover, the robustness of the cointegration rank proves less sound when this variable is substituted for DPO.

[17] A detailed discussion of parameter instability following demographic change is in Auerbach-Kotlikoff (1983). As to the measures we considered, namely ASP and PO, when they are included in the set of co-integrating variables, either jointly or separately, their coefficients show a very low level of significance. Contrary to the evidence reported in Cigno-Rosati (1996), in our case demographic changes seem not to be determined simultaneously with the economic variables entering the long-run consumption function.

The criteria for selecting the order p of the VAR mostly pointed to a parsimonious specification, with p=1 (Table 1). This option, which is convenient in view of the limited sample size, was validated by diagnostic control on the residuals of each single equation of the VAR(1), which does not signal evidence of misspecification. The co-integration rank proved equal to 1 with a confidence level of 95 per cent for both the trace and the maximal eingenvalue tests (Table 2), with signs of stability over time[18].

Table 2

COINTEGRATION TESTS

(with restricted intercepts and no trends in the VAR

47 observations from 1952 to 1998. Order of VAR = 1.
List of variables included in the cointegrating vector:
LC LYP WYP WSSY Intercept
List of I(0) variables included in the VAR: DPO DZ
List of eigenvalues in descending order:
.70793 .26928 .14747 .032734 0.00

LR Test Based on Maximal Eigenvalue of the Stochastic Matrix

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	57.8455	28.27	25.80
r<= 1	r = 2	14.7452	22.04	19.86
r<= 2	r = 3	7.4987	15.87	13.81

LR Test Based on Trace of the Stochastic Matrix

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	81.6536	53.48	49.95
r<= 1	r>= 2	23.8080	34.87	31.93
r<= 2	r>= 3	9.0629	20.18	17.88

With reference to the values of the coefficients of the sole cointegrating vector, the maximum likelihood estimates have been compared with those deriving from univariate approaches, both in the static version of Engle-Granger (1987) and in the dynamic versions of Stock-Watson (1993) and Pesaran-Shin (1995).

As a common result, the role of the real interest rate in determining the expenditure of Italian households proves negligible in the long run. This is probably because the substitution and income effects almost offset one another, as the latter is bolstered by the

[18] This result is attained by progressively restricting the final year from 1998 to 1990. Robustness extends to the order of the VAR, since the evidence of a single co-integrating vector is confirmed for p=2 and p=3.

traditionally high share of Treasury bonds held by Italian households[19]. In view of this evidence, which is in line with previous analysis (Rossi-Visco, 1995), the variable *R* has not been reported in the tables we present.

One important finding is the sizeable discrepancy in the point estimates for coefficients of WYP and WSSY, which signals a much stronger influence of real and financial assets than of social security wealth on the long-run equilibrium in consumption plans. This evidence, which is confirmed by estimates deriving from the univariate approach (Tables A2a-c in Appendix), is weaker under restricted samples. Taking, for comparability, the same period as in Rossi-Visco, 1952 to 1993, the broad equivalence they found in the effects of the two wealth components is confirmed, although at higher values for the coefficients (Table 3, column B) [20]. This finding is more extensively discussed in sections that follow. At this point, it is also worth mentioning the test for the significance of social security wealth, which has occasionally proven problematic in the literature (Favero et al., 1994). In our case, evidence based on the LR test clearly supports a coefficient significantly different from zero for variable WSSY (Table 4, column B)[21].

Table 3

ML ESTIMATES SUBJECT TO EXACTLY IDENTIFYING RESTRICTION(S)
ESTIMATES OF RESTRICTED COINTEGRATING RELATIONS
(restricted intercepts and no trends in the VAR; S.E.'s in brackets)

Order of VAR = 1, chosen r =1. List of I(0) variables included in the VAR: DPO DZ
List of imposed restriction(s) on cointegrating vectors: a1=-1. 47 observations from 1952 to 1998

	A		B	
LC	-1.0000		-1.0000	
LYP	.85593	(.073940)	.77448	(.054642)
WYP	.06148	(.030992)	.056429	(.023743)
WSSY	.027035	(.010295)	.046393	(.0094723)
Intercept	1.5853	(.83063)	2.5502	(.64317)

A: 47 observations from 1952 to 1998 B: 42 observations from 1952 to 1993

[19] The coefficient of *R* was negative but not significant at both 95 and 90 per cent levels.

[20] In Rossi-Visco the common value of the two coefficients was around 0.022.

[21] A similar result, not reported in the table, applies to coefficient of WYP, too.

With reference to income, the LR test does not reject the assumption of long-run homogeneity between consumption and income at a confidence level of 95 per cent. But the point estimate of the coefficient of disposable income is lower than in previous studies and, under the multivariate approach, shows signs of decline in the nineties. In line with Miniaci-Weber (1998), we can trace this result to set of factors that, in addition to pension reforms, combined to likely reduce households' measure of permanent income since the recession in 1992-93. They include higher taxation of the self-employment income and stricter job and wage control in the public sector.

Table 4

ML ESTIMATES SUBJECT TO OVER-IDENTIFYING RESTRICTION(S)
ESTIMATES OF RESTRICTED COINTEGRATING RELATIONS
(restricted intercepts and no trends in the VAR; S.E.'s in brackets)

Order of VAR = 1, chosen r =1. 47 observations from 1952 to 1998.
List of I(0) variables included in the VAR: DPO DZ

	List of imposed restriction(s) on cointegrating vectors	
	(A)	(B)
	a1=-1; a2=1	a1=-1; a4=0
LC	-1.0000 (*NONE*)	-1.0000 (*NONE*)
LYP	1.0000 (*NONE*)	1.0525 (.062732)
WYP	-.4048E-3 (.032685)	-.019516 (.053893)
WSSY	.011160 (.0072457)	0.00 (*NONE*)
Intercept	.091036 (.18551)	-.43103 (.64679)
	LR Test of Restrictions	
	CSHQ(1)= 2.4308	CHSQ(1)=4.1545

Following the identification of a single co-integrating vector, the error correction representation has been estimated to characterize the adjustment process of consumption demand towards its long run equilibrium (Table 5).

In this context the set of regressors has been enlarged to include, not only the conditioning variables considered in the co-integration analysis, but also the ratio of capital gains to disposable income and a measure of uncertainty[22]. As in Muellbauer (1994), the

[22] The specification of the statistical model proceeds from the general to the specific, by progressive reductions starting from an initial setting in which lags up to second order are considered for each variable.

latter is computed as the absolute value of deviations of the income growth rate from trend at any period, with trend in turn being approximated by the average growth rate in the previous five periods, i.e. $UNC_t = \left| dLYP_t - "trend" \right|$. Finally an exogeneity analysis is performed to check if the likely occurrence of measurement errors and simultaneity bias may have caused a violation of the orthogonality assumption of regressors and disturbances. In particular, the Wu-Hausman procedure has been implemented for variables LYP, R, WSSY and UNC to test for zero coefficients in the consumption function of the residuals derived by regressing each of the four variables on lags of themselves and consumption (Table A3 in Appendix). Since the test supports the assumption of exogeneity, the evidence validates the consistency of the OLS estimates and precludes the need for instrumental variables.

Table 5

OLS ESTIMATION OF ERROR CORRECTION REPRESENTATION
(White heteroscedasticity adjusted S.E.' in brackets)

Dependent variable is DLC; 45 observations used from 1954 to 1998
ECM=-1.0*LC+0.856*LYP+0.061*WYP+0.027*WSSY+1.585

Regressor	Coefficient	S. E.	T-Ratio	Regressor	Coefficient	S. E.	T-Ratio
DLC(-1)	0.376	0.095	3.949	DPO	0.039	0.009	4.309
DLC(-2)	-0.248	0.098	-2.534	DPO(-2)	0.021	0.011	1.934
DLYP	0.318	0.067	4.772	DASP	-0.042	0.011	-3.974
DZ	-0.159	0.052	-3.072	DASP(-2)	0.037	0.015	2.501
DZ(-1)	0.119	0.036	3.257	DWSSY	0.009	0.004	2.182
DR	0.149	0.062	2.410	DWSSY(-1)	0.007	0.004	1.902
DR(-1)	-0.126	0.056	-2.235	ECM(-1)	0.132	0.043	3.051
UNC	-0.002	0.001	-2.908				

R-Squared 0.8998; R-Bar-Squared 0.8531; S.E. of Regression 0.0079; F-stat. F(14, 30) 19.2522
Residual Sum of Squares 0.0019; DW-statistic 1.7364

Test Statistics	LM Version	F Version
Serial Correlation	CHSQ(1)=0.99182	F(1,29)=0.65358
Functional Form	CHSQ(1)=0.72701	F(1,29)= 0.47621
Normality	CHSQ(2)=0.77192	-
Heteroscedasticity	CHSQ(1)=0.98531	F(1,43)= 0.9626
Predictive Failure (*)	CHSQ(3)= 4.6257	F(3, 27)= 1.5419

(*) Chow's second test based on a restricted sample 1954-1995.

6. Interpretation of the empirical results

The estimated error correction model helps to characterize the way a variety of factors considered in the current debate in Italy may impact upon private consumption and wealth accumulation in the short and in the long run.

The first point to arise is the possibility that in the last decade social security changes may have curbed the structural decline in the saving rate of Italian households[23]. The sizeable reductions in pension benefits, first enacted in 1992 and then in 1995, in combination with the ongoing debate over the future need for further restrictive measures, may have lead Italian households to revise their expected sustainability of pensions in financing consumption plans.

Empirically, this is a reasonable interpretation of the signs of decline in the coefficient for social security wealth during the nineties. As mentioned, the maximum-likelihood point estimates of the cointegrating vector show a larger effect for the variable WSSY in the restricted than in the full sample (the coefficient slipping from 0.046 to 0.027 if the final year is moved from 1993 to 1998; Table 3). The opposite result holds for WYP, whose coefficient rises from 0.056 to 0.061. Accordingly, the ratio between the two coefficients, which in the linear model is a proxy for ϑ, is cut almost in half, from 0.82 to 0.44, between the restricted and the full sample. We tend to interpret this as a sign of increased uncertainty of Italian households' over their future pension entitlement given reiterated reforms and the prospect of more to come. Indeed, as recent survey evidence shows for American households, the people with the least confidence in the future of social security exhibit the highest saving rates (Bernheim, 1995).

Moreover, solving the co-integrating vector for the desired ratio of real and financial assets to income, namely WYP*, this variable has risen sharply, much more than the actual WYP ratio (Fig. 6). As a consequence the gap, which was negative in the two previous decades, turned positive in the nineties.

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[23] For the factors explaining such a structural correction, though Italian households' saving rate still remains higher than in the other main industrial countries, see Guiso et al. (1994).

The sluggish growth of private consumption long after the crisis of 1992-93 might result from households' decision to accumulate real and financial assets, partly to compensate for lower social security wealth, partly as precautionary buffer stocks against increased uncertainty over future pension benefits. From this standpoint, the evidence that in the most recent years of the sample the gap between WYP* and WYP, though narrowing, remains significantly positive implies that the adjustment of consumption plans may still be continuing after 1998. Indeed, the estimated low value for the loading coefficient of the cointegrating vector in the error correction model points to quite a slow rate for the transition to the long-run equilibrium.

As to the factors affecting the adjustment process, the coefficient of the measure of cyclical uncertainty proves significant and negative, signaling a depressing effect of income variability on consumption[24]. On the contrary, the impact of capital gains proves negligible, probably because pronounced volatility has hindered sound assessment by households of the ensuing changes in purchasing power. This is confirmed when capital gains on real and on financial assets are considered separately[25], although the standard error related to the latter decreases.

A final remark about demographic changes: despite the difficulty in interpreting, at aggregate level, the signs of the coefficients for the ratio of people over 65 (PO) and life expectancy at birth (ASP), their significance confirms the convenience of taking account of the changing patterns of households' heterogeneity in assessing saving behavior. And we can also put forward a tentative interpretation of the signs of the coefficients. To the extent that PO proxies for the relative frequency of pensioners with respect to workers, its positive correlation with consumption dynamics conforms to the life cycle model. The assessment of life expectancy is more ambiguous. Insofar positive changes proxy better health conditions, they may imply increasing length of both working and retirement stages, so the effect on the

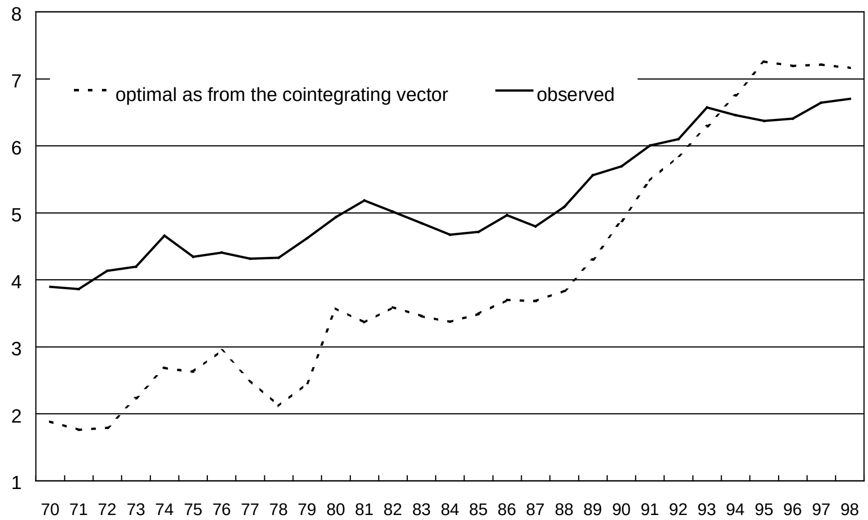
[24] The unemployment rate, an alternative proxy for cyclical uncertainty in the empirical literature (Feldstein, 1974 and Barro-Mac Donald, 1979), was not statistically significant, either in levels or in first differences or as multiplicative factor of changes in income. These results are in line with the evidence for Italy in Boone et al. (1998).

[25] In Boone et al. (1998), variations in the stock exchange index, adjusted for inflation, play a much weaker role in affecting consumption in Italy than in the other main industrial countries, according to evidence based on quarterly data covering the period from 1976-Q1 to 1996-Q2.

ratio between the two, which is the key variable in determining consumption (Modigliani, 1970), remains undetermined. From this standpoint, our empirical analysis does not help to solve the theoretical ambiguity. Nevertheless, the inclusion of demographic variables proves worthwhile even at the aggregate level of our analysis, despite the ensuing difficulty of interpretation, in order to control for the bias that would otherwise affect estimates of the intensity of the offset between pension and non-pension wealth. As is pointed out in Gale (1998), this bias generally arises when the empirical model controls for current income and pensions separately rather than for total lifetime resources, and a correction for the effects of changes in demographic structure would be required.

Fig. 6

RATIO OF HOUSEHOLDS' REAL AND FINANCIAL ASSETS TO DISPOSABLE INCOME



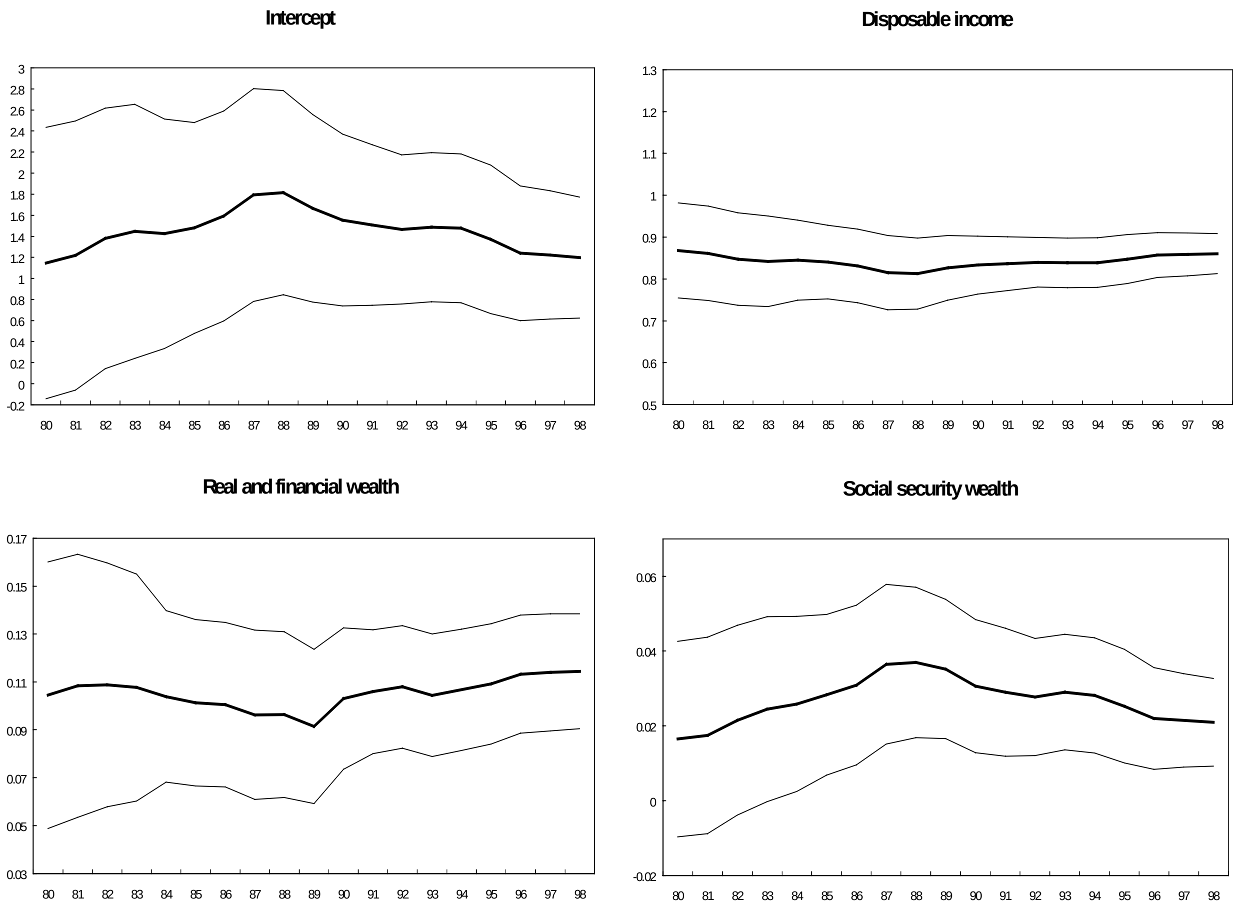
7. Consumption demand and pension reforms: a case for a model shift?

The signs of a weakening impact of social security wealth on households' expenditure in Italy suggest exploring the stability of the empirical model estimated with respect to the reform of the pension system. From an econometric standpoint, a regime shift in the data generating process may affect co-integration analysis mainly by leading conventional tests to accept an erroneously low rank of co-integration and by changing the co-integrating vector

(Gregory-Hansen, 1996, Quintos, 1997 and Inoue, 1999)[26]. In view of the evidence that the rank confirms equal to one in the full and restricted samples, in this section we concentrate on the stability of the cointegrating vector. In particular, we deal with a univariate context, since a multivariate approach would be problematic due to the large size of the co-integrating vector relative to the limited number of observations available in the sample for the years following the candidate structural break.

Fig. 7

COEFFICIENTS AND STANDARD ERRORS' BANDS BASED ON RECURSIVE OLS OF COINTEGRATING REGRESSION



[26] The additional issue of reduced power in ADF test for unit root has been addressed in the stationarity analysis of WSSY (section 5.2).

In first place, the recursive OLS estimates of the co-integrating regression show that some changes in parameters might have actually occurred, even if they are statistically limited in size (Fig. 7). Apart from the temporary rise in the intercept around the end of the eighties, more pronounced changes apparently affected the coefficients of the two wealth components, in opposite directions as we noted.

In second place, we have run the test of parameter stability suggested by Hansen (1992), on the basis of both the fully modified estimator of Phillips-Hansen (1990) and the dynamic OLS of Stock-Watson[27]. In particular, we have computed the test statistics *SupF* by recursively splitting the full sample with a cut point *t* moving from 1984 to 1993, and each time evaluating a *Wald* statistic to test parameter stability; the maximum value in this sequence compares with proper critical values. While we do not impose the date for the structural break to augment the power of the test, we found it reasonable to restrict the possible range for alternative dates[28].

Table 6

HANSEN'S TEST FOR PARAMETER STABILITY

SupF test for the null hypothesis of stability against unknown structural break under p=0 and m_2=3.

Asymptotical critical value at significance level of 5% is *17.2* (Table 1, Hansen 1992)

T	P-H	S-W.	*t*	P-H	S-W.
1984	1.67	0.98	1989	2.58	6.63
1985	0.27	2.82	1990	8.16	1.22
1986	0.19	2.76	1991	6.06	1.29
1987	0.95	3.06	1992	3.13	2.18
1988	1.16	6.49	1993	5.37	6.98
SupF	8.16	6.98			

P-H = Fully Modified Phillips-Hansen estimator; S-W=Stock-Watson dynamic OLS.

[27] This proves convenient to check for robustness to different corrections for serial correlation implied by the two estimators; asymptotic critical values for the stability test prove to be equivalent (Hansen, 1992).

[28] In the lower bound, since we know that the institutional change eventually materialized not earlier than 1992; in the upper bound, due to a degree-of-freedom constraint in latest sub-sample.

The results are univocally against the instability of the co-integrating vector (Table 6), although the high variability of the *Wald* statistic since the nineties suggests that the latest observations are too few to be informative enough that a change in model could prove statistically significant[29]. In this respect, an additional problem is the fact that two shocks came close upon one another (the 1992 and 1995 pension reforms) in the first half of the nineties.

Third, we have tested whether dummy variables controlling for a change in regime since 1990 prove significant in the co-integrating regression. In this case, in order to make a correct inference despite non-stationarity, we first compute the dynamic OLS estimator by including interaction terms between dummies and the variables entering the co-integrating vector. Then we re-scale standard errors by the factor λ_v / s, with s being the usual standard error of the regression and λ_v a consistent estimate of long-run variance of residuals[30].

Evidence confirms our previous results, as both the slope and level dummies are all scarcely significant – although with signs occasionally different than expected (Table 7, column A). Note that the estimation sample terminates in 1996, given the two leads in first differences we include in the dynamic specification, thus eliminating potentially useful observations for a sounder appraisal of the change in regime.

In order to attenuate this problem and gain degrees of freedom in regression, we have turned to standard OLS estimators by including only the relatively more significant dummies, namely DUWSSY and DUWYP, which interact with pension and non-pension wealth respectively (Table 7, column B)[31]. The opposite signs of the coefficients of these variables signal that changes in the consumption function may be under way in line with our interpretation: stronger accumulation of real and financial assets would follow the reduction

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[29] We have replicated the test computing the estimators' long-run variances by different kernel and under (locally) changing bandwidth. (We set the central value of this parameter as low as 2 in view of the evidence of nearly "white" residuals.) The results, not reported, confirm solid evidence against parameter instability.

[30] A detailed description of the computation of the rescaling factor, which is required for achieving efficient estimates uniquely for coefficients of I(1) variables, is in Hayashi (2000). In this case, we found this procedure less cumbersome than a kernel.

[31] These two variables have been identified in the context of dynamic OLS, via a gradual reduction of the general specification reported in table 7.

in expected pension benefits. As noted earlier, a sound statistical identification of the magnitude of these movements is fraught with difficulty, due to the size of our sample and the uncertainty still surrounding the final setup of social security in Italy.

Table 7

COINTEGRATION ANALYSIS UNDER REGIME SHIFT

Dependent variable: LC

A			B		
Dynamic OLS with adjusted S.E.)			Ordinary Least Squares		
43 observations used for estimation (1954-96)			47 observations used for estimation (1952-98)		
List of variables included in the regression $\Sigma_{j=-2}^{+2}\Delta\, lyp_{t-j}$, $\Sigma_{j=-2}^{+2}\Delta\, wssy_{t-j}$, $\Sigma_{j=-2}^{+2}\Delta\, wyp_{t-j}$					
Regressor	Coefficient	S.E.	Regressor	Coefficient	S.E.
INTP	1.8637	0.3892	INTP	1.5935	0.51304
LYP	0.8157	0.0323	LYP	0.8326	0.0427
WYP	0.1026	0.0193	WYP	0.0896	0.0220
WSSY	0.0286	0.0079	WSSY	0.0346	0.0108
DZ	-0.0419	0.1509	DZ	-0.3268	0.1458
DPO	0.0545	0.0309	DPO	0.0368	0.0410
DUINT	9.1026	10.264	DUWYP	0.02813	0.0185
DULYP	-0.5944	0.7121	DUWSSY	-0.0199	0.0136
DUWYP	-0.1229	0.0913			
DUWSSY	-0.0281	0.0237			

DUINT=DT*INTP; DULYP=DT*LYP; DUWSSY=DT*WSSY; DT=0 if t≤1989, 1 otherwise.

8. Conclusions

The paper provides fresh evidence on the aggregate demand of Italian households according to the augmented life cycle model of consumption. Under the caution ordinarily urged in view of the controversial computations of social security wealth, the main findings may be summarized as follows:

- symmetrically with previous evidence on the effects of past expansionary reforms of social security, recent restrictive corrections have contributed significantly to depress the expenditure of Italian households;

- social security wealth proves to exert a much weaker effect on long-run consumption demand than real and financial assets, and the degree of substitutability between the two components of wealth results turns out to be far lower than in previous time series evidence;

• signs of instability in consumption demand have increased during the last decade, although they are not statistically significant, probably because of the paucity of observations subsequent to Italy's social security reforms. Recursive point estimates suggest a declining coefficient of pension wealth in the long-run consumption function and a rising coefficient of other components of wealth since the turn of the nineties;

• in the same period the difference between the observed and the desired ratio of non-pension wealth to income has turned negative, after being positive in the seventies and eighties. Presumably, in an effort to compensate for reductions in actual pension payments and to provide a financial cushion against the uncertainty over future entitlements, consumers stepped up accumulation of real and financial assets;

• demographic trends, such as changes in life expectancy at birth and in the share of the population over 65, play a significant role in affecting consumption demand, although they are not co-determined with it in the long run. The result calls for taking account of the major demographic changes of Italian society in characterizing household expenditure, even at the aggregate level we consider;

• first estimates of capital gains on net financial and real assets held by Italian households show a negligible impact on private consumption, likely due to their high volatility in level and composition.

As a final remark, the interpretation of our evidence requires some caution in view of two issues, which would make good topics for future research:

i) the size and timing of the correction in expected pension benefit of households – which, we have argue, appeared well in advance to the actual reform in 1992 - are subject to revisions of the micro-model underlying the estimates we have adopted; these revisions which are currently under way, as a part of a separate project at the Bank of Italy;

ii) time series analysis of consumers' expenditure notoriously misses important composition effects, which in our case risk being particularly strong due to the great heterogeneity of Italian households with respect to pension entitlements and access to financial markets, in addition to individual variations in age and labor income.

Appendix I: Statistical tables

Table A1a

UNIT ROOT TESTS FOR VARIABLE LYP

DF regressions include intercept and linear trend

43 observations used in the estimation of all ADF regressions (1956-98)

	Test Statistic	LL	AIC	SBC	HQC
DF	-.95061	99.5005	96.5005	93.8587	95.5263
ADF(1)	-.85854	99.6908	95.6908	92.1684	94.3918
ADF(2)	-.87460	99.7168	94.7168	90.3138	93.0931
ADF(3)	-.71166	99.8460	93.8460	88.5624	91.8975
ADF(4)	-.50433	99.9982	92.9982	86.8340	90.7250

95% critical value for the augmented Dickey-Fuller statistic = -3.5162

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table A1b

UNIT ROOT TESTS FOR VARIABLE LC

DF regressions include intercept and linear trend

43 observations used in the estimation of all ADF regressions (1956-98)

	Test Statistic	LL	AIC	SBC	HQC
DF	.86343	119.1179	116.1179	113.4761	115.1437
ADF(1)	-.0076770	121.3640	117.3640	113.8416	116.0651
ADF(2)	.0081516	121.3649	116.3649	111.9619	114.7412
ADF(3)	-.29601	121.8187	115.8187	110.5351	113.8703
ADF(4)	.061977	122.2591	115.2591	109.0949	112.9859

95% critical value for the augmented Dickey-Fuller statistic = -3.5162

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table A1c

UNIT ROOT TESTS FOR VARIABLE LWYP

DF regressions include intercept and linear trend

43 observations used in the estimation of all ADF regressions (1956-98)

	Test Statistic	LL	AIC	SBC	HQC
DF	-2.1379	13.2187	10.2187	7.5769	9.2445
ADF(1)	-2.3465	13.7644	9.7644	6.2420	8.4655
ADF(2)	-2.5244	14.2591	9.2591	4.8561	7.6354
ADF(3)	-1.9905	14.7414	8.7414	3.4578	6.7930
ADF(4)	-1.5896	15.0172	8.0172	1.8530	5.7441

95% critical value for the augmented Dickey-Fuller statistic = -3.5162

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table A1d

UNIT ROOT TESTS FOR VARIABLE R

DFuller regressions include intercept but not trend

43 observations used in the estimation of all ADF regressions (1956-98)

	Test Statistic	LL	AIC	SBC	HQC
DF	-2.3076	107.4149	105.4149	103.6537	104.7654
ADF(1)	-2.3149	107.5419	104.5419	101.9001	103.5677
ADF(2)	-1.7361	108.4757	104.4757	100.9533	103.1767
ADF(3)	-1.5934	108.4937	103.4937	99.0907	101.8700
ADF(4)	-1.3640	108.6723	102.6723	97.3887	100.7239

95% critical value for the augmented Dickey-Fuller statistic = -2.9303

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table A1e

UNIT ROOT TESTS FOR VARIABLE WSSY

DF regressions include intercept and linear trend

43 observations used in the estimation of all ADF regressions (1956-98)

	Test Statistic	LL	AIC	SBC	HQC
DF	.21230	-23.7477	-26.7477	-29.3895	-27.7220
ADF(1)	-.24155	-23.0156	-27.0156	-30.5380	-28.3146
ADF(2)	.23394	-22.3527	-27.3527	-31.7557	-28.9764
ADF(3)	.077986	-22.3060	-28.3060	-33.5896	-30.2545
ADF(4)	-.17617	-22.1347	-29.1347	-35.2989	-31.4079

95% critical value for the augmented Dickey-Fuller statistic = -3.5162

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table A1f

UNIT ROOT TESTS FOR VARIABLE WSSY

OLS estimation based on Newey-West
adjusted S.E.'s

Dependent variable is DWSSY; 47 observations from 1956 to 1998

Regressor	Coefficient	Standard Error	T-Ratio
INTP	.41963	.17364	2.4167
WSSY (-1)	-.066453	.039350	-1.6888

Table A1g

UNIT ROOT ANALYSIS FOR WSSY

Structural break as in model C in Perron (1989)

Dependent variable is WSSY; OLS Estimation, with 43 observations from 1956 to 1998;

DU=1 if t=1990+1, 0 otherwise – DTB=1 if t>1990, 0 otherwise – DT=0 if t≤1989, t otherwise

T_B=1990, λ=0.8; 5% critical value for $\tau_{\tilde{}}$ equal to –4.04

Regressor	Coefficient	S.E.	Regressor	Coefficient	S.E.
INTP	0.704	0.206	DWSSY(-3)	0.029	0.147
WSSY(-1)	0.511	0.196	DWSSY(-4)	0.089	0.137
TREND	0.094	0.039	DT	-0.206	0.137
DWSSY(-1)	0.191	0.207	DTB	6.742	5.615
DWSSY(-2)	-0.059	0.173	DU	-0.168	0.501

$$\tau_{\hat{\alpha}} = -2.5503$$

Table A2

UNIVARIATE COINTEGRATION ANALYSIS

Stock-Watson Approach

OLS estimation based on Newey-West adjusted S.E's

Dependent variable is LC. 43 observations used for estimation from 1954 to 1996

Regressor	Coefficient	S.E.	Regressor	Coefficient	S.E.
INTP	1.828	0.537	DWYP	-0.069	0.022
LYP	0.820	0.044	DWYP(-1)	-0.051	0.020
WSSY	0.033	0.006	DWYP(-2)	-0.032	0.024
WYP	0.088	0.020	DWYP(+1)	0.017	0.023
DZ	-0.077	0.117	DWYP(+2)	0.028	0.018
DPO	0.068	0.035	DWSSY	-0.031	0.009
DLYP	-0.701	0.250	DWSSY(-1)	-0.018	0.008
DLYP(-1)	-0.509	0.233	DWSSY(-2)	-0.013	0.007
DLYP(-2)	-0.395	0.201	DWSSY(+1)	0.001	0.008
DLYP(+1)	0.000	0.194	DWSSY(+2)	-0.002	0.008
DLYP(+2)	0.084	0.155			

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	Pesaran-Shin Approach ARDL (1, 1, 0,0)		Engle-Granger Approach OLS Estimation	

Dependent variable is LC. 46 observations used for estimation from 1952 to 1998

Regressor	Coefficient	S. E.	Coefficient	S.E.
LYP	.902	.072	.865	.037
WYP	.067	.028	.115	.012
WSSY	.022	.010	.020	.005
INTP	.964	.812	1.148	.419
DZ	-.259	.265	-.270	.135

Table A3

ERROR CORRECTION MODEL AND WU-HAUSMANN TEST FOR EXOGENEITY OF DLYP, DR, DWSSY, UNC

Dependent variable is DLC - List of variables added to the regression: RDLYP, RDR, RDWSSY, RUNC
43 observations used for estimation from 1956 to 1998

Regressor	Coefficient	S. E.	T-Ratio	Regressor	Coefficient	S. E.	T-Ratio
DLC(-1)	0.44825	0.1163	3.5435	RUNC	-0.0043	0.0027	-1.6291
DLC(-2)	-0.2091	0.10918	-1.9313	UNC	0.0014	0.0024	0.5870
DLYP	0.2415	0.21347	1.1313	DPO	0.0294	0.0118	2.4871
DZ	-0.1310	0.0558	-2.5940	DPO(-2)	0.0179	0.0121	1.4804
DZ(-1)	0.1347	0.0476	2.3294	DASP	-0.0478	0.0125	-3.8344
DR	-0.0372	0.1518	-0.2451	DASP(-2)	0.0281	0.0159	1.7659
DR(-1)	-0.0737	0.0862	-0.8554	DWSSY	0.0098	0.0038	2.5810
RDLYP	-0.0796	0.1971	-0.4040	DWSSY(-1)	0.0078	0.0041	1.9122
RDR	0.1979	0.1692	1.1694	ECM(-1)	0.0848	0.0539	1.5726
RDWSSY	-0.0171	0.0105	-1.6307				

Variables RDLYP, RDR, RDWSSY, RUNC are residuals of regressions of DLYP, DR, DWSSY, UNC, respectively, on: const, DLC(-1), DLC(-2), DLYP(-1), DLYP(-2), DR(-1), DR(-2), DWSSY(-1), DWSSY(-2), UNC(-1), UNC(-2)

Joint test of zero restrictions on the coefficients of additional variables:
LM Statistic.: CHSQ(4)= 8.3460; LR Statistic: CHSQ(4)= 9.268; F Statistic F(4,24)=1.445

Appendix II: Data set

For most of the variables for which specific computations have been required, sources and methods are broadly the same as in Rossi-Visco (1994). With respect to the data set adopted in that paper, however, important revisions have occasionally resulted from updated evidence about some basic parameters and statistics. Here we define the main variables considered in the empirical analysis and sketch their computation procedure.

LYP stands for the logarithm of households' disposable income at constant prices, net of pension payments[32] and adjusted for the loss in purchasing power of net financial assets due to inflation. As to the former, which substitute for social transfers considered in Rossi-Visco, they are computed after taxes on the basis of the average tax rate given by the ratio of total direct taxes paid by households and their income. The Hicksian correction for monetary erosion of net financial assets is computed with respect to expected inflation, whose index has been retrieved from the results of the business survey run by *Forum-Mondo Economico*.

WYP is the ratio of net real and financial assets to our measure of disposable income. Real assets are the sum of end-of-period stock of durables, dwellings and land. Each component is estimated by projecting backward and forward the relevant value for a benchmark year, on the basis of the flow of investments and depreciation. With reference to dwellings, which accounts for more than 70 per cent of total real wealth of Italian households, major revisions have occurred with respect to the estimate adopted in Rossi-Visco. They mainly reflect a lower value for the benchmark year 1980 and a new price index adopted to turn stock into market value, which now comes from the Bank of Italy based on data provided by *Il Consulente Immobiliare*. As a consequence, on average for the first five years of the nineties the estimated stock of dwelling wealth adopted in the paper is around 10 per cent lower than in Rossi-Visco.

R stands for the real interest rate, which is measured by $\log[(1+i)/(1+\pi)]$, with i the nominal interest rate on Treasury bonds with a residual life longer than one year and π the expected inflation rate computed as described.

[32] Their source is ISTAT and, for years before 1973, Morcaldo (1977).

CAPS is households' total capital gains on dwellings, bonds and equities. Each of the three components has been estimated separately on a yearly basis, by applying annual changes in the respective capitalization index as to the value of the stock at the end of the previous year. In the case of dwellings, revaluation has been proxied by changes in the price index of new and recently restructured buildings devised by the Bank of Italy, as noted. In the case of equities and bonds, estimates of capital gains include the capitalization of dividends or interests, that has not been transferred to the assets' holders.

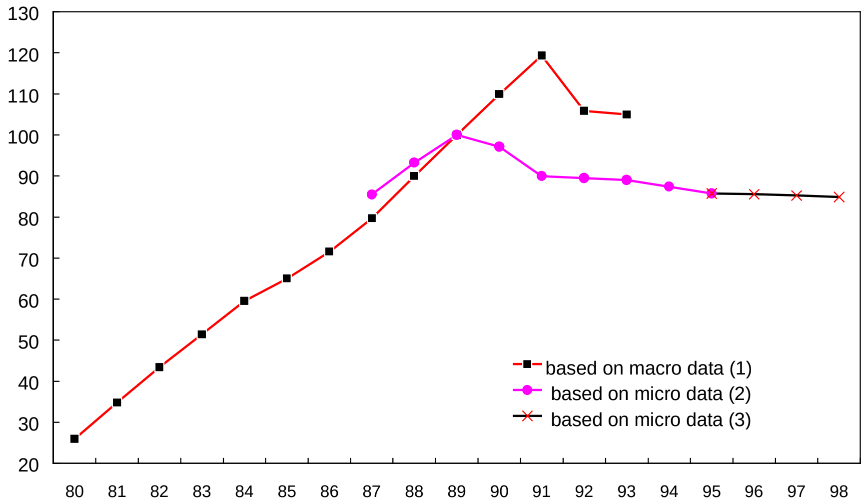
WSSY is the ratio of gross social security wealth to disposable income. In the definition we adopted, it is given by the present discounted value of the sum of benefits that current contributors expect to receive after retirement and those that current retirees will keep receiving (Feldstein, 1974). The computation of this variable is controversial, since it involves a variety of unobservable factors, such as expected age at retirement and death, future labor incomes, the discount factor. In the absence of a single time series covering the full period, the measure we used comes from a combination of different sources. They include: a) estimates made by Beltrametti for the period 1951-93 (Beltrametti, 1995); b) estimates made by Nicoletti-Altimari (1999) for any two years between 1987 and 1995; c) provisional Bank of Italy simulations for 1994-98 (Ando et al., 1999). Whereas the last two are based on the same micro-model of the Italian economy, which has been estimated with data from the Bank of Italy Survey of Household Income and Wealth, the first is obtained from the traditional, more aggregated approach. The measure we adopted is equal to the latter until 1989; for later years it is given by a projection of the same series on the basis of the growth rates of the micro estimates.

The identification of the cut point as 1989, which proves crucial for the recent development of variable WSSY, is predicated on the following argument. The "micro" measure sub b), shows a correction in social security wealth earlier than the "macro" measure sub a), suggesting adjustment in WSSY prior to the actual reform in 1992. This is presumably due to the fact that the "micro" measure allows one to estimate a key parameter like retirement time, which otherwise is restricted to be equal to the legal requirement, and takes better account of composition effects coming from heterogeneous workers. Both factors are likely to prove especially important under recurrent institutional changes, which may induce stratification of different regimes[33]. There is in fact some indication of this in the much larger discrepancies between the two measures that we observe towards the end of the seven-year period in which they overlap (see chart).

Fig. A1

ESTIMATES OF GROSS SOCIAL SECURITY WEALTH

(Index 1989=100)



(1) Based on data from Beltrametti (1995). (2) Based on data from Nicoletti (1999). (3) Based on data from Ando et al. (1999).

The computation of the measure we adopted proceeded in several steps. First, we linearly interpolated the original biannual estimates by Nicoletti-Altimari and merged the resulting yearly series with that referred to sub c), after converting both into indexes (1995=100) to correct for discrepancies. Second, the micro estimates, now running from 1987 to 1998, have been merged with the "macro" one, again after converting both into indexes (1989=100). Finally, levels have been retrieved.

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[33] In Italy, pension benefits currently conform to three different regimes, identified according to the length of the contribution period: a) the "old" one, as heavily reformed in 1992, still holds for people with at least 18 years of contributions in 1995; b) the "new" one, introduced in 1995, applies to people whose contribution period started in 1993; c) the "intermediate" one, concerning people with less than 18 years of contributions in 1995, combines the old regime for contributions until 1995 and the new regime for the subsequent years. The specific formulas and details on evolution of Italian pension system are discussed in Bank of Italy (1995) and Peracchi et al. (1995).

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