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National saving and social security in Italy

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Summary

We have recently suggested that net government transfers have substantially contributed to the decline of the aggregate Italian saving rate. Changes in social security laws and regulations which took place in the late 1960s and early 1970s apparently weakened the link between contributions and benefits permitting a time path of aggregate consumption in excess of what would have occurred in the absence of such changes. In this paper, these results are revised and extended and, if anything, strengthened: slightly less than half the reduction in the private "equilibrium" saving rate observed over the last 30 years appears to be due to the increase of social security wealth. The paper also assesses the likely impact of the 1992 social security reform and suggests that in the long run the impact might be sizeable accounting for up to a 3 points higher saving ratio.

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1. Introduction

According to official sources, aggregate Italian saving rates have been recently experiencing a substantial decline. The timing and the extent of such decline largely depend on the specific definition of saving as well as on the researcher's attitude toward the host of measurement issues involved.† However, for gross as well as

† The standard national income accounts (NIA) definition is inadequate on various respects. First of all, consumption is usually defined as total expenditures on consumer goods and services. The former include expenditures on consumer durables which should be more properly considered as a component of household investment. A proper (economic) definition of consumption should instead only include the services provided to households by the stock of durables. A second issue arises when we consider that in the NIA definition of income interest payments and receipts are included at their nominal value. As it has been recognized for quite a time, however, income should be adjusted for the loss of

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for net rates, for national and for private saving ratios, for both "adjusted" and "unadjusted" magnitudes, major trends appear to be beyond dispute and provide a somewhat unexpected description of the behaviour of Italian consumers. As it turns out, from a peak of about 24% at the beginning of the 1960s, the net national saving rate has declined to about 9% in 1993.

In particular, in the early 1990s private saving rates (defined as net of depreciation, adjusted for durables and inflation and computed as ratios to net national disposable income) have been fluctuating around 10%, nine percentage points down from the average saving rate prevailing in the 1960s. In the same years, government saving (adjusted for inflation) has plunged to an unprecedented −1%, after more than 30 years of current account surpluses averaging to 2%. As a result, today's overall Italian saving rate is about 15 percentage points below the one observed during the so-called "Italian economic miracle".

As can be easily gauged from Figure A where national, private and government saving rates (corrected for durables and inflation) are presented, the beginning of the decline in the Italian saving rate can be placed just short of the middle of the 1970s, in conjunction with the first oil shock, when the growth rate of the Italian economy first started to fall. It is only in the 1980s, however, that the reduction of the national saving rate became quite substantial, fully reflecting the progressive decline of government saving along with the negative trend of private saving.

Several alternative explanations for the decline of the saving rate have been put forward by the most recent research.† In a recent article (Rossi & Visco,1994), we have scrutinized the impact of fiscal policy on private sector behaviour to suggest that a thorough assessment of the interplay of private and public decisions can shed some light on the evolution of the Italian saving rate in the most recent decades. In particular, focusing on the 1951–1990

purchasing power that inflation induces on the stock of nominally denominated debt. This is particularly relevant when national saving is split between its private and government components. Ignoring the loss of principal due to inflation implies an overestimation of private savings: this simply amounts to neglecting a possibly important source of revenue for the government, the so-called "inflation tax". Finally, even if it is controversial, it is likely that saving should be considered net of the depreciation of the capital stock. In this section, the movements of national, private and government saving rates (all computed in relation to the net national disposable income) are examined allowing for a proper treatment of consumer durables and accounting for the capital stock depreciation and the "inflation tax". For detailed statistics on post war Italian saving trends see the Statistical Appendix to Ando, Guiso and Visco (1994). The updated time series used in this paper are contained in its working paper version, available on request from the authors.

† Besides the papers by Fornero and Castellino (1990) and Jappelli (1991), the interested reader is referred to Modigliani and Jappelli (1993) and to Ando, Guiso and Visco (1994).

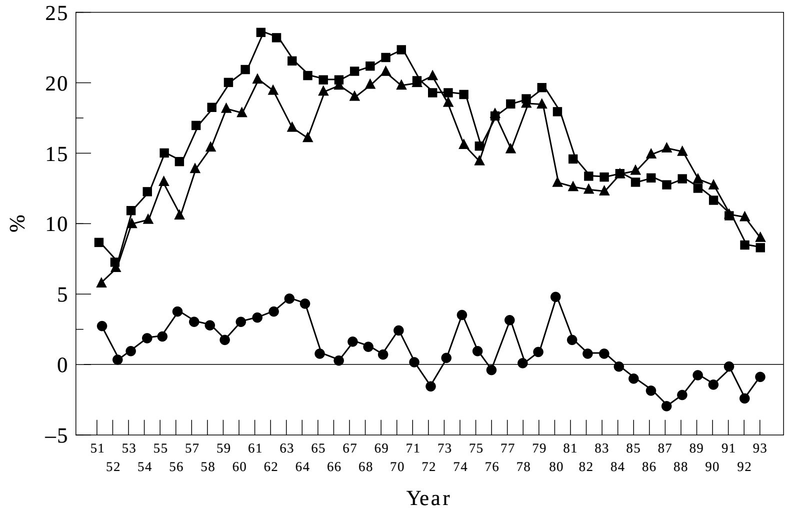

FIGURE A. Adjusted net saving rates. –■–: National; –●–: government; –▲–: private.

Source: Ando, Guiso and Visco (1994), Statistical Appendix, updated and revised.

period, we submitted that the changes in social security laws and regulations which took place in the late 1960s and early 1970s severely weakened the link between contributions and benefits permitting a time path of aggregate consumption in excess of what would have occurred in the absence of such changes. It is worth recalling that those changes entailed both a replacement effect due to the introduction of an unfunded social security system for specific segments of the population as well as, for a given institutional structure, a lifetime wealth increment. The evidence presented in our paper apparently provides a clear-cut outcome, indicating that the increase in social security benefits has substantially contributed to the decline in private saving that has been observed in the course of the 1970s and the 1980s. We estimated that one-third (or about three percentage points) of the decline of private saving (in terms of net private disposable income) observed since the 1960s may be attributed to the rise in the net social security wealth to income ratio.

That changes in laws and regulations determining retirement benefits might be at the root of important economic phenomena of the last 20 to 30 years comes as no surprise. Italy is the OECD country with the highest pension expenditures to GDP ratio and, in fact, Italy's pension system, since the early 1980s, has been the subject of lively discussions aiming, on the one hand, at controlling the future pattern of expenditures and at suggesting, on the other hand, appropriate legislative changes. Many attempts at structurally reforming the system took place in the 1980s in an effort to halt the rapid growth in the pension expenditures. At last, in 1992, a social security reform was enacted: while certainly contributing to reverse the present negative trends, it was however far from sufficient. An overall social security reform has been recently approved by Parliament: its effects still need to be thoroughly assessed.

The time is ripe, therefore, to revise and update Rossi and Visco's (1994) exercise. A reassessment of the likely impact of social security on the aggregate saving rate is the (admittedly limited) purpose of this paper. In Section 2 a brief review of the main patterns of the Italian social security system is presented, providing quantitative evidence on its more recent evolution. Section 3 considers the time-series evidence on the substitutability of social security and fungible wealth. Section 4 provides some microeconomic evidence on the same topic. Section 5 makes use of the time-series and cross-section evidence to provide a preliminary assessment of the impact of social security reforms on the future evolution of the saving rate and concludes the paper.

2. The Italian social security system

Social security transfers have been rising in Italy dramatically over the last 40 years from a level of 7 to one around 22% of net national income. The entire rise (about 15 points) is accounted for by the movements of pensions, public health services having grown in line with income over the period.

An analysis of the determinants of the rapid growth in pension expenditures highlights the role of demographic trends and of substantial increases in the number of pensions (throughout the period) as well as in average pension benefits (from 1976 onward).† The growth of pension expenditures which took place from the turn of the 1960s onwards reflects "the impact of the introduction of special plans for self-employed workers and increased use of disability pensions for welfare purposes. This rapid growth carried over into the 1970s in connection with the switch in 1968 from a contributions-based system to an earnings-based system and the far-reaching reform of INPS pensions in 1969, in particular the introduction, and subsequent improvement, of methods of indexing pension benefits to prices and real earnings. By contrast, in the 1980s the growth of pension expenditure slowed down. This reflected the decrease in the number of disability pensions . . . , the

† Franco (1992) reviews the post-war evolution of government transfers in Italy and provides a detailed account of their determinants.

gradual winding down of the hyperindexation of pensions to prices . . . , the slowdown in growth due to the adjustment of pensions to real earnings, the attenuation of the effects of self-employed schemes when they reached steady state, and the introduction of means testing for welfare benefits." (Franco & Morcaldo, 1990: p. 107).

From the point of view of government finances, it is crucial to underline that the spectacular increase of social transfers (and particularly of pensions) has not been matched by that of social contributions from whatever source (from employers as well as from employees). Between 1951 and 1993 total transfers increased in relation to net national income about three times (with three "jumps" in 1965, 1975 and 1981, when major institutional changes took place, with further consequences for the following years). By no means the corresponding increase of social contributions was a modest one, having amounted to two and a half times that of national income. This was not sufficient, however, to avoid a rise of net transfers (from about 1% of income until the middle of the 1970s to about 5% in most of the 1980s) entirely due to the generosity of the national social security system. As a result, in 1992, the equilibrium contribution rate (that is, in a pay-as-you-go system such as the Italian one, the ratio of pension transfers to wages) was projected to reach by the year 2010 levels well above any future foreseeable level of the contribution rate. For example, the 2010 equilibrium rate for the (private) dependent workers scheme was projected to reach 54%.†

As Franco *et al.* (1994) have shown, there was a serious imbalance in Italian generational policy: as of 1990, future generations of Italians would have to face lifetime net tax burdens much higher (almost double) than those faced by those just born. A deep reform of the system was therefore unavoidable and took place at the end of 1992. As a result, theoretical equilibrium rates were substantially lowered and, for the dependent workers' scheme, were projected to reach between 40 and 47% in the year 2010 depending on the indexation scheme (INPS, 1993). Equilibrium in the long run was however far from achieved, projections for some of the major pension funds still implying a long-run equilibrium rate which largely exceeded any economically meaningful level of the contribution rate. Moreover, the existence of separate schemes still tended to generate undesirable distributional effects.

The post-war evolution of the Italian social security system is clearly reflected in the time path of social security wealth (Figure B). The net pension wealth to national disposable income ratio

† See INPS (1993: p. 132).

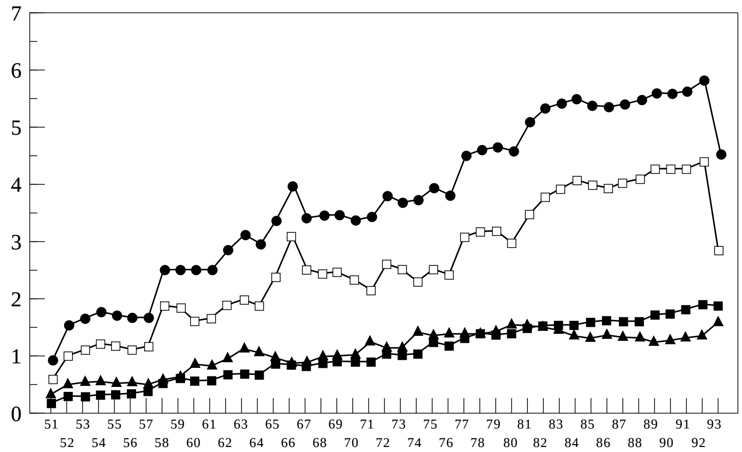

FIGURE B. Social security wealth to net disposable income ratios. –■–: Social security wealth (retirees); –●–: social security wealth (gross); –▲–: social security wealth (contributions); –■–: social security wealth (net).

Source: Beltrametti (1995).

rises dramatically from 0·6 in 1951 to about 4·5 in 1992 to drop to 2·9 after the 1992 reform.† The increase is due to demographic trends as well as to the deepening and widening of the system which took place since the late 1950s. At each point in time, it concerns both the workforce and (to a lesser extent) the retired population.

The information content of social security wealth should not be over emphasized, though. Net social security wealth should measure the actuarial value of the pension benefits which individuals expect to receive net of the actuarial value of the social security taxes which they expect to pay. As such, this quantity should reflect changes over time in such factors as the population age structure, social security coverage and benefit and tax rules. It is, therefore, a rather elusive quantity whose measurement involves formidable assumptions at the individual as well as at the aggregate level. It should be stressed, though, that, whatever the measurement issues involved, social security wealth is positive by definition in institutional settings such as the Italian one: it simply measures the cost of switching from a public pay-as-you-go system to a private fully-funded system. It tells almost nothing about the

† Social security wealth computations for 1993 assume changes in workers' retirement choices corresponding to the changes in the legal retirement age. See Beltrametti (1995), where the major assumptions behind pension wealth computations are spelled out in detail.

long-run sustainability of the social security system which is, instead, correctly measured by the equilibrium contribution rate. Consequently, international comparison of pension liabilities (such as in van den Noord & Herd, 1993) can to a large extent be questioned (see also Franco, 1995).

However, social security wealth might have a lot to say about household saving behaviour and labour supply choices. An increase in pension benefits in the form of, say, changes in the rules determining social security payments (such as the changes repeatedly observed in the 1970s and 1980s) could affect the saving behaviour of all individuals through their impact on expectations, since, in a life-cycle framework and for a given retirement age, such changes would, *ceteris paribus*, create an incentive for the current generation to increase its consumption.† Therefore, in such circumstances, an increase in transfer payments signalling a permanent increase in pension benefits‡ (the bill being eventually footed by later generations) could be expected to have an appreciable negative impact on private savings and hence, for a given level of current income, on the private saving rate. The shrinking private saving rate would add up to a larger government deficit, thereby leading to an even larger cut in the national saving rate.

3. Private saving and social security wealth: time-series evidence

Needless to say, the impact of social security on saving is far from being a novel issue, having been debated widely at both the theoretical and empirical levels.§ The controversy over the validity of the life-cycle model extended *à la* Feldstein (1976) is still very much an open issue and, indeed, Auerbach and Kotlikoff (1983), thoroughly examining the available cross-section and time-series tests of the role of social security, have argued that, given current data, neither type of information has much potential for settling the controversy.

Rossi and Visco's (1994) results suggest, however, that the Italian

† It should be recalled that, if the retirement age is not fixed, such effect may be partly offset by induced phenomena of early retirement.

‡ It is implicitly assumed here that the increase in transfer payments is perceived as sustainable by the agents. This warning forcefully applies to the Italian case where the inherent risk of unsustainability of the social security system had been recognized since the late 1970s.

§ See the pioneering works of Feldstein (1974) and Barro (1974). Apart from Rossi and Visco (1994), aggregate (but inconclusive) time-series evidence for Italy is reported in Modigliani and Sterling (1983).

experience may represent an ideal case study providing for substantial variability in social security regulations and hence in the social security wealth variable. As Figure B shows, the net social security wealth to income ratio accurately reflects the discretional adjustment of pensions to prices which took place in 1952, 1958, 1962 and 1965 as well as the automatic indexation first introduced in 1969, extended in 1976 and subsequently amended in 1983. It also incorporates the effects of the introduction of special schemes for self-employed farmers in 1957, for artisans in 1959 and for shopkeepers in 1966. It points out the major switch from a contributions-based system to an earning-based system (in two steps, the first one in 1969 and the second in 1976). It finally highlights the far-reaching consequences of the 1992 reform (as possibly perceived by the agents). As in Rossi and Visco (1994), we shall exploit the information contained in the behaviour of the economy around the periods of change, hopefully avoiding what Auerbach and Kotlikoff correctly regarded as a major source of misspecification.

Following Modigliani (1986), the aggregate behaviour of Italian consumers may be described by the simple tale of rational, utility maximizing, agents allocating optimally resources to consumption over their entire life, as depicted by the stripped down version of the life-cycle hypothesis of saving. As Feldstein (1974) first suggested, though, the social security system should influence family's consumption. If labour-leisure choices are unaffected, this influence should only pass through the effect of social security benefits and contributions on its lifetime budget constraint. If there are no constraints on net worth prior to the time of death, the social security impact is summarized by the family's lifetime social security wealth which is simply the difference between the discounted value of lifetime net social security benefits and contributions.

In such a case, consumer decisions would be conditional on human wealth (which may be assumed roughly proportional to current labour income) and on a variable given by the sum of net non-human (real and financial) wealth and net pension wealth. The latter could be weighted by a constant parameter, θ, allowing for a difference between private assets and net social security wealth due, among other things, (i) to the presence of a bequest motive, which would imply $\theta>1$ (Williamson & Jones, 1983), (ii) to income uncertainty, which would entail $\theta<1$ (Samwick, 1994), and (iii) to the perception of a risk of unsustainability of the social security system that would also imply $\theta<1$.

In the specification of an aggregate saving function within such an extended life-cycle framework, the issue of double counting between disposable income and social security wealth should however be addressed. In particular, according to the standard NIA definition, disposable income (y^d) is inclusive of social security benefits, essentially to current retirees (y^s), and is net of social security contributions by workers and firms (y^c).† In a pay-as-you-go system such as the Italian one, whereby the largest part of contributions are directly paid by firms, the latter choice is justified. This would amount to define, as usual, y^d as the net real disposable income of the private sector, and to revert to a gross definition of social security wealth ($w^{sg} = w^s + w^{sc}$, where w^s is net pension wealth and w^{sc} is the discounted stream of current and future contributions).‡

Furthermore, if current social security transfers to households are already accounted for in the pension wealth variable, they should either be subtracted from disposable income to avoid double counting or should enter the aggregate saving function in a restricted form. Under a set of somewhat restrictive assumptions, the aggregate equilibrium behaviour of consumers would then translate into a simple saving rate function such as

$$\frac{s}{y^d} = \varphi_0 + \varphi_1 g + \varphi_2 r + \varphi_3 \left(\frac{w^r}{y^d} + \theta \frac{w^{sg}}{y^d} \right) + \varphi_4 \frac{y^s}{y^d}, \tag{1}$$

where the φ's are (approximately) constant aggregate coefficients, and the restriction $\varphi_4 = 1 - \varphi_0$ should be satisfied.

Equation (1)—that is an equilibrium linear relationship between the saving rate (s/y^d), the rate of growth of population and productivity (g), the real interest rate (r), the net non-human (real and financial) wealth to income ratio (w^r/y^d), the gross social security wealth to income ratio (w^{sg}/y^d) and the social transfers to income ratio (y^s/y^d)—may be arrived at expanding around constant values of r, g, (w^r/y^d) and (w^{sg}/y^d) the extended aggregate life-cycle consumption function,§ with y^d standing then for the sum of aggregate net labour income

† In what follows, all consumption, income and wealth variables are in real terms.

‡ The alternative specification, with disposable income defined as gross of current social security contributions and pension wealth being net of their present value counterpart, could also be tested. In the empirical analysis summarized in Appendix 1, this has been done with such a specification always turning out to be dominated by the one discussed in the text.

§ See Appendix A in the working paper version of this paper.

(y^l), net income from capital (rw^r) and net social transfers to households (y^s).†

A thorough statistical analysis, summarized in Appendix 1, led to the error-correction representation of the data generating process reported in Table 1,‡ where coefficient estimates, associated standard errors and diagnostic statistics, when available, for both least squares and instrumental variables estimates are also reported. Comparison of LS and IV estimates suggests, however, simultaneity bias to be negligible. In Table 1,§ $\ln(c/y^d)$ is the logarithm of the consumption to income ratio,¶ where c stands for "economic" consumption and y^d is net private disposable income, adjusted for inflation;†† z stands for real government consumption; and r is defined as $\ln[(1+R)/(1+\pi)]$, where R is the nominal interest rate on long-term Government bonds and π is the expected annual rate of change of consumer prices.‡‡

The error-correction estimates have a plausible economic interpretation, and appear to possess acceptable statistical properties. Needless to say, to assess the relevance of Auerbach and Kotlikoff's (1983) assertion about the likely instability of aggregate

† An equilibrium condition equivalent to equation (1) (with $\varphi_4 = 1 - \varphi_0$) might also be written excluding the y^s/y^d term from the right-hand side of the equation and replacing y^d by $y = y^d - y^s$. Alternatively, one might still focus on the standard definition of the private saving rate, leaving the definition of disposable income unchanged, replacing w^{sg} with w^{sw} (the discounted stream of pension benefits for current workers) and leaving φ_4 unconstrained. In this case y^s would also catch the effect on private saving due to w^{sr} (equal to $w^{sg} - w^{sw}$, i.e. the discounted stream of pensions for current retirees). The relative merits of the three alternative (and, in principle, entirely equivalent) formulations are scrutinized in Appendix 1, where the reasons supporting the use of equation (1) are indicated.

‡ Apart from adequately representing the data process, the error-correction representation may implicitly account for other phenomena likely to affect the Italian saving rate, such as financial market imperfections. Liquidity constraints may be caught by the short-term components of the error-correction equation while borrowing constraints may affect the target wealth to income ratio and thereby the long-run relationship (see the "Introduction" to Ando, Guiso & Visco, 1994).

§ All variables are reported, with sources and definitions, in a data appendix to the working paper version of this article. As far as the social security wealth variables are concerned, it should be noticed that for the year 1992 they were constructed as weighted averages of pre-reform and post-reform figures, with weights equal to 0·75 and 0·25, respectively, to account for the fact that, although enacted by the end of the year, the 1992 reform of social security was publicly announced by the end of the summer of 1992.

¶ Clearly, $-\ln(c/y^d) \approx s/y^d$. Notice that while the saving ratio is a bounded variable (at least from above), $\ln(c/y^d)$ is not.

†† The empirical exercise undertaken in this section, therefore, assumes that while households pierce the corporate veil, they do not pierce the government veil. Evidence on this assumption is presented in Rossi and Visco (1994: Appendix).

‡‡ Unweighted average of semi-annual survey estimates of expected inflation.

TABLE 1 *Non-linear error-correction representation (sample: 1954–1993, least squares and instrumental variables estimates)*

$$\underset{}{\Delta \ln c_t =} \underset{(0{\cdot}042)}{0{\cdot}459\Delta \ln y^d_t} + \underset{(0{\cdot}043)}{0{\cdot}098\Delta \ln(y^s/y^d)_{t-1}} - \underset{(0{\cdot}130)}{0{\cdot}476\Delta \ln z_t} - \underset{(0{\cdot}168)}{0{\cdot}340\Delta \ln z_{t-1}} + \underset{(0{\cdot}081)}{0{\cdot}275\Delta \ln c_{t-1}}$$

$$- \underset{(0{\cdot}091)}{0{\cdot}545}[\ln(c/y^d)_{t-1} + (\underset{(0{\cdot}041)}{0{\cdot}288} - \underset{(0{\cdot}003)}{0{\cdot}033}((w^r/y^d)_{t-1} + \underset{(0{\cdot}091)}{0{\cdot}658}(w^{sg}/y^d)_{t-1}) + \underset{(0{\cdot}106)}{0{\cdot}321}r_{t-1} + \underset{(0{\cdot}041)}{(1–0{\cdot}288)}(y^s/y^d)_{t-1})]$$

Note: LS estimates; in parentheses, White's standard errors.

$\bar{R}^2 = 0{\cdot}866$, $dw = 2{\cdot}010$, $\hat{\sigma} = 0{\cdot}0084$, serial correlation: $F(1, 29) = 0{\cdot}095$, functional form: $F(1, 29) = 1{\cdot}237$, normality: $\chi^2(2) = 0{\cdot}952$, heteroskedasticity: $F(1, 38) = 0{\cdot}537$, arch: $F(1, 29) = 0{\cdot}607$.

$$\Delta \ln c_t = \underset{(0{\cdot}085)}{0{\cdot}538\Delta \ln y^d_t} + \underset{(0{\cdot}047)}{0{\cdot}080\Delta \ln(y^s/y^d)_{t-1}} - \underset{(0{\cdot}150)}{0{\cdot}533\Delta \ln z_t} - \underset{(0{\cdot}187)}{0{\cdot}426\Delta \ln z_{t-1}} + \underset{(0{\cdot}096)}{0{\cdot}315\Delta \ln c_{t-1}}$$

$$- \underset{(0{\cdot}104)}{0{\cdot}567}[\ln(c/y^d)_{t-1} + (\underset{(0{\cdot}041)}{0{\cdot}298} - \underset{(0{\cdot}004)}{0{\cdot}033}((w^r/y^d)_{t-1} + \underset{(0{\cdot}090)}{0{\cdot}650}(w^{sg}/y^d)_{t-1}) + \underset{(0{\cdot}102)}{0{\cdot}336}r_{t-1} + \underset{(0{\cdot}041)}{(1–0{\cdot}298)}(y^s/y^d)_{t-1})]$$

Note: IV estimates; in parentheses, White's standard errors.

$\bar{R}^2 = 0{\cdot}850$, $dw = 2{\cdot}066$, $\hat{\sigma} = 0{\cdot}0089$, validity of instruments: $\chi^2(5) = 5{\cdot}081$, serial correlation: $\chi^2(1) = 0{\cdot}159$, functional form: $\chi^2(1) = 2{\cdot}169$, normality: $\chi^2(1) = 0{\cdot}522$, heteroskedasticity: $\chi^2(1) = 1{\cdot}683$.

Instruments: besides predetermined variables and $\Delta \ln z_t$, the list of instruments includes $\ln y^d_{t-1}$, $\ln(c/y^d)_{t-2}$, $(w^r/y^d)_{t-2}$, $(w^{sg}/y^d)_{t-2}$ and r_{t-2}.

time-series relationships, we have carefully investigated the stability of the estimated model and inspected the sequence of one-step ahead residuals. The equation standard errors corresponding to the latter show little variation, in spite of the substantial changes in the series under observation. Inspection of similar evidence for the estimated coefficients based on the recursive estimator, has provided further evidence of parameter constancy.

All long-run coefficients appear to be rather precisely determined and the restriction $\varphi_4 = 1 - \varphi_0$ turns out not to be rejected by the data. The estimates indicate θ to be about two-thirds, thereby implying in equilibrium a non-negligible replacement of private savings by social security.†

The present estimate of θ is about twice as large as the corresponding IV estimate in Rossi and Visco's (1994). The large and significant difference originates entirely from the proper treatment of the double counting issue and not from the information contained in the 1991–1993 observations. The estimate of θ presented in Table 1 is, in fact, very robust to changes in specification (in both the short-run and the long-run components of the non-linear error-correction specification) and in the methodology of estimation (if anything, when compared with the other estimates presented in Appendix 1, those of Table 1 seem to be on the lower side).

Turning to the interpretation of recent trends in the Italian private saving rates on the basis of the equilibrium parameters derived from the IV estimates presented in Table 1 (the upper panel of Table 1 makes it clear that looking at LS estimates would not change the results appreciably), Table 2 presents in the first row changes in the net private saving rate (that is s/y^d) over the 1954–1992 period, leaving 1993 for later discussion. The second row reports the fitted equilibrium values of the latter and the following rows present its decomposition.‡

It should be underlined that the fitted equilibrium values do not incorporate the short-run dynamic effects of the error-correction specification. Therefore they are not directly comparable to the actual changes of the net private saving rate. Moreover, it is crucial to understand that, in the light of equation (1), the saving rate can be affected by the pension system in two conceptually different ways. On the one hand, pension income adds up to other income components to determine private disposable income and hence the

† It should be noticed that θ measures the extent by which private non-human wealth is replaced by social security wealth conditional on given interest rate, productivity growth and net saving rate. In the light of the "error-correcting" nature of equation (1) (see the "Introduction" to Ando, Guiso & Visco, 1994), it can be shown that in a steady state equilibrium the replacement rate between pension and non-human wealth may be much lower than θ.

‡ The stability of the short- and long-run terms of the error-correction representation suggests that such in-sample simulation exercise is a legitimate one.

TABLE 2 *Changes in the aggregate private saving rate and their determinants (changes between period averages: 1954–1992)*

	1962–72 / *1954–61*	*1973–79* / *1962–72*	*1980–92* / *1973–79*	*1980–92* / *1962–72*
Changes in actual saving rate:				
$\Delta(s/y^d)$	0·049	−0·040	−0·051	−0·094
Changes in equilibrium saving				
rate: $\Delta[(s/\hat{y}^d)]$	0·031	−0·083	−0·048	−0·131
Contributions of:				
Δg	−0·015	−0·047	−0·015	−0·062
Δr	−0·008	−0·010	0·020	0·010
$\Delta(w^r/y^d)$	0·066	−0·028	−0·034	−0·062
$\Delta(w^{sg}/y^d)$	−0·034	−0·022	−0·035	−0·057
$\Delta(y^s/y^d)$	0·022	0·024	0·016	0·040

Note: Changes in the aggregate private saving rate are computed from period averages of ratios of actual saving figures to the net disposable income of the private sector adjusted for inflation. The contribution of g is arrived at setting the short-run components of the error-correction representation at their equilibrium values and solving for the growth equilibrium rate.

observed saving rate. On the other hand, present and future developments of the pension system influence present consumption behaviour thereby affecting the saving rate.

As it can be seen from Table 2, the large increase of the private saving rate observed between the 1950s and the 1960s turns out to be mostly attributable to the declining non-human wealth to income ratio which reflects the unprecedented growth experienced by the postwar Italian economy until 1962. At the same time, the negative contribution of social security wealth arising from the appearance of new pay-as-you-go pension schemes is far from negligible and averages to about three percentage points. In the following time interval (between the 1960s and the years following the first oil shock) rising housing prices coupled with the productivity slowdown to induce an upward trend in the non-human wealth to income ratio. The result, strengthened by a declining real interest rate, is a three percentage points fall in the equilibrium saving rate. In the same period, though, further structural changes in the social security system began to feed into consumption decisions thereby adding to the fall in the saving rate. In the following period (ending in 1992), notwithstanding the rapid growth of private financial assets (mainly public debt), the joint contribution of the non-human wealth to income ratio and of productivity growth somewhat tapers off and is partly counterbalanced by the positive effect of the rising real interest rates, with the "pension effect" accounting for much of the further fall in the private saving rate. Finally, comparing the 1980s with the 1960s, slightly less than half of the entire fall of the private equilibrium saving rate appears to be due to the increased gross social security wealth to income ratio. The negative pension wealth effect on saving in the period (some six percentage points) is partially offset by the direct effect of pension expenditures on disposable income (with a positive impact on the saving rate of about four percentage points). To put it differently, it might be said that, had the Italian households adhered to the standard life-cycle framework (with $\theta = 0$), the net private saving rate currently observed would not be much different from the one prevailing in the mid-1970s.

4. Household saving and social security wealth: cross-section evidence

The foregoing time-series analysis has proposed that the rise in the pension wealth to income ratio could have been responsible for a substantial part of the decrease in the private saving rate observed over the last quarter of a century. Recognizing that aggregation and other sources of bias may affect the standard aggregate consumption functions, in this section we supplement the time-series evidence with the insights provided by the estimation of a saving function on the microeconomic information derived from the Bank of Italy 1991 Survey on Household Income and Wealth (SHIW).†

The measurement of the net present value of social security benefits and taxes is discussed in Appendix 2, where summary information on the households included in the sample is also provided. Measures of expected social security benefits were computed under two sets of alternative assumptions. First, an attempt was made to assign each worker to a specific social security fund and to define consequently, on the basis of the relevant legislation, age eligibility requirements, reference wages for computing pensionable earnings, accrual rates and contribution requirements. Second, reference was made to the subjective assessment of expected pension benefits and expected retirement age of active population. Differently than Jappelli (1995), the latter alternative was not confined to the subsample of respondents. To avoid sample selection bias and account for endogeneity of households' expectations about future benefits, separate reduced form equations were estimated for the expected replacement rate and retirement age and the results used to impute expectations to non-respondents.

† See Brandolini and Cannari (1994) for a full length description of the Bank of Italy SHIW.

The empirical specification of the saving function replicates, at the household level, equation (1):†

$$s_i/y_i^n = \psi_0 + \psi_1(w_i^r/y_i^n) + \psi_2(w_i^{sg}/y_i^n), \tag{2}$$

where real interest and productivity growth rates are assumed to be constant across households, $s_i/y_i^n \approx -\ln(c_i/y_i^n)$, y_i^n stands for normal age-adjusted net earnings of household *i* defined as the average present discounted value of the stream of net future earnings, and coefficients ψ_0, ψ_1 and ψ_2 depend on a number of socio-demographic variables including age of the household's head, number of children, education, location, and the like. In addition, regressors include (the log of) normal earnings and dummies indicating the purchase of life insurance or pension funds.‡ As in the case of the time-series estimates, a gross definition of social security wealth is used.

Appendix 2 reports and comments on the cross-section data and the parameter estimates based, alternatively, on the two measures of social security wealth. Since both ψ_1 and ψ_2 are allowed to vary with household-specific socio-demographic characteristics, the cross-section results permit to derive household-specific estimates of the degree of substitutability between fungible and pension wealth, as measured by $(\theta = \psi_2/\psi_1)$. The latter parameter, however, only captures the (static) displacement effect of social security wealth under the assumption of constant saving rates. The distributions of the θ's over the sample of about 5 600 households are summarized in Table 3 in terms of their medians, means and standard deviations.

Both measures of social security wealth convey the same message: as of 1991, the static substitution effect of social security on private accumulation tended to be large, and, on average, above unity. However, both distributions tend to be skewed to the left, with the overall median as low as 0·9, not far from the time-series estimates.§

† It is worth noting that, contrary to the previous microeconometric literature, we estimate a saving function and not a fungible assets equation.

‡ Normal earnings should not affect the saving ratio if preferences are homothetic. It should be kept in mind, though, that other variables included in the regression (such as the region of residence or dummies signalling contributions to life insurance and pension funds) may also proxy for lifetime earnings.

§ The extent to which social security wealth substitutes for private wealth might seem much larger than that obtained by Brugiavini (1987) and Jappelli (1995), where estimation of a fungible assets equation yields an estimated offset between one tenth and one fifth. It is worth stressing, however, that the difference between those estimates and the present ones is entirely due to the reference theoretical framework, in that estimation of a fungible asset equation (as opposed to the estimation of a saving function) would yield results comparable to those reached by the previous authors. This is not surprising since, as observed, the two estimation strategies refer to two different notions of equilibrium.

TABLE 3 *Distribution of θ: summary information (1991 SHIW)*

	Social security wealth based on					
	legislation			*expectations*		
	median	*mean*	*S.D.*	*median*	*mean*	*S.D.*
Overall	0·933	1·052	0·521	1·172	1·275	0·547
By age of head:						
less than 34	0·737	0·777	0·237	0·918	0·946	0·258
between 34 and 44	0·720	0·757	0·295	1·150	1·142	0·396
between 44 and 54	1·169	1·220	0·488	1·379	1·401	0·536
between 54 and 64	1·249	1·289	0·524	1·417	1·404	0·555
By region:						
North and Central Italy	1·133	1·245	0·540	1·381	1·438	0·588
Southern Italy	0·688	0·734	0·276	0·989	1·008	0·327
By residence:						
Rural areas and towns	1·009	1·131	0·521	1·236	1·357	0·545
Cities	0·617	0·643	0·263	0·828	0·858	0·324
By household size:						
3 or less	1·004	1·109	0·497	1·174	1·290	0·569
more than 3	0·885	0·998	0·537	1·172	1·262	0·525
By number of income recipients:						
1	0·849	0·927	0·437	1·215	1·330	0·504
more than one	1·009	1·134	0·554	1·113	1·240	0·571
By number of children:						
2 or less	0·935	1·062	0·520	1·172	1·281	0·556
more than 2	0·913	0·990	0·522	1·212	1·242	0·491
No. of observations		5587			5587	

Notes: (i) estimates of θ are based on the cross-section estimates reported in Table 7, (ii) the estimation sample excludes households with pension income only; moreover, in computing the present table, households showing negative values of θ (for a total of 41 observations) have been excluded.

As Table 3 indicates, as of 1991, the effect of pension wealth on saving tended to exceed that of non-human wealth in the case of older households, while for younger households the opposite result appears to hold. This outcome possibly indicates the perception among the young of the unsustainability of the Italian social security system. Interestingly, both the old and the young households may have proved to be right in that the 1992 reform left the former group basically unaffected, severely hitting the latter. The age effect is apparently the predominant one, the differences between estimated θ's in different regions of the country being largely determined by their different age compositions.

5. 1993 and beyond

The central tenet of this paper is that the changes in social security laws and regulations which took place in the late 1960s and early 1970s severely weakened the link between contributions and benefits permitting a time path of aggregate consumption in excess of what would have occurred in the absence of such changes. It is worth recalling that those changes entailed both a replacement effect due to the introduction of an unfunded social security system for specific segments of the population as well as, for a given-institutional structure, a lifetime wealth increment.

The evidence emerging from aggregate time series as well as from cross-section data provides a rather clear-cut outcome, indicating that a large part of the decline of private saving (in terms of net private disposable income) observed since the 1960s may be attributed to the rise in the social security wealth to income ratio. Given the specification adopted in this work, a major role is played by the productivity slowdown.

Past trends have been, however, substantially reversed in the last couple of years. A major reform was enacted in 1992 and a further attempt to halt social security expenditure has taken place in the spring of 1995. As Figure B shows, the 1992 reform may have determined a substantial fall (as large as 20%) in the gross social security wealth (a gigantic “capital levy” of about 2000 trillions lire), leading the gross pension wealth to net national disposable income ratio to fall to 4·6 in 1993 from 5·9 in 1992. This outcome appears to be confirmed by the information available at the individual level: preliminary computation of expected household social security wealth for 1991 and 1993 suggests an average fall of about 10% (the difference being given by postponed retirement).

To the extent that the overall reform of the Italian pension system just approved by Parliament will confirm the implications of the 1992 measures, on the basis of the parameter estimates of Table 1 it might be inferred that such dramatic changes in the social security arrangements could bring about a rise of the private saving ratio of up to 3 percentage points in the steady state, with figures lower than 3% accounting for the fact that the saving augmenting replacement effect may be counterbalanced by a saving reducing retirement effect.

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References

Ando, A., Guiso, L. & Terlizzese, D. (1994). Dissaving by the elderly, transfer motives and liquidity constraints. In A. Ando, L. Guiso & I. Visco, Eds. *Saving and the Accumulation of Wealth*, pp. 188–213. Cambridge: Cambridge University Press.

Ando, A., Guiso, L. & Visco, I. Eds. (1994). *Saving and the Accumulation of Wealth*. Cambridge: Cambridge University Press.

Ando, A. & Modigliani, F. (1963). The "life cycle" hypothesis of saving: aggregate implications and tests. *American Economic Review*, **53**, 55–84.

Auerbach, A.J. & Kotlikoff, L.J. (1983). An examination of empirical tests of social security and saving. In E. Helpman, A. Razin & E. Sadka, Eds. *Social Policy Evaluation: An Economic Perspective*, pp. 161–179. New York: Academic Press.

Barro, R.J. (1974). Are government bonds net wealth? *Journal of Political Economy*, **82**, 1095–1117.

Beltrametti, L. (1995). On pension liabilities in Italy. *Ricerche Economiche*, **49**, 405–428.

Brandolini, A. & Cannari, L. (1994). Methodological appendix: the Bank of Italy's survey of household income and wealth. In A. Ando, L. Guiso & I. Visco, Eds. *Saving and the Accumulation of Wealth*, pp. 369–386. Cambridge: Cambridge University Press.

Brugiavini, A. (1991). Effetti delle pensioni sul risparmio nella teoria del ciclo vitale. In T. Jappelli, Ed. *Bilancio pubblico e risparmio privato*, pp. 43–68. Milano: F. Angeli.

Cannari, L., D'Alessio, G., Raimondi, G. & Rinaldi, A.I. (1990). Le attività finanziarie delle famiglie italiane. Temi di Discussione No. 172. Roma: Banca d'Italia.

Doornick, J.A. & Hendry, D.F. (1994). *PcFiml 8.0*. London: International Thomson Publishing.

Engle, R.F. & Granger, C.W. (1987). Co-integration and error correction representation, estimation and testing. *Econometrica*, **55**, 255–276.

Feldstein, M. (1974). Social security, induced retirement, and aggregate accumulation. *Journal of Political Economy*, **82**, 905–926.

Feldstein, M. (1976). Social security and saving: the extended life cycle theory. *American Economic Review*, **66**, 77–87.

Fornero, E. & Castellino, O. Eds. (1990). *Formazione e impiego della ricchezza delle famiglie*. Torino: Fondazione Giorgio Rota.

Franco, D. (1992). Alcune note sull'espansione della spesa pubblica in Italia (1960–1990). In *Ente Einaudi, Il Disavanzo Pubblico in Italia: Natura Strutturale e Politiche di Rientro*, pp. 211–355. Bologna: Il Mulino.

Franco, D. (1995). Pension liabilities. Their use and misuse in the assessment of fiscal policies. Economic Papers, No. 110. Brussels: European Commission.

Franco, D. & Morcaldo, G. (1990). The Italian pension system: development and effects on income distribution and poverty. *Labour*, **4**, 105–124 and 133–159.

Franco, D., Gokhale, J., Guiso, L., Kotlikoff, L.J. & Sartor, N. (1994). Generational accounting. The case of Italy. In A. Ando, L. Guiso & I. Visco, Eds. *Saving and the Accumulation of Wealth*, pp. 128–160. Cambridge: Cambridge University Press.

Hansen, H. & Johansen, S. (1993). Recursive estimation of cointegrated VAR-models. Preprint No. 1. Copenhagen: Institute of Mathematics and Statistics.

INPS (1993). *Le pensioni domani*. Bologna: Il Mulino.

Jappelli, T., Ed. (1991). *Bilancio pubblico e risparmio privato*. Milano, F. Angeli.

Jappelli, T. (1995). Does social security reduce wealth accumulation? Evidence from Italian survey data. *Ricerche Economiche* **49**, 1–31.

Johansen, S. (1988). Statistical analysis of cointegrating vectors. *Journal of Economic Dynamics and Control*, **12**, 231–254.

Modigliani, F. (1986). Life cycle, individual thrift and the wealth of nations. *American Economic Review*, **76**, 297–313.

Modigliani, F. & Jappelli, T. (1993). Why has the Italian national saving rate declined? In S. Biasco, A. Roncaglia & M. Salvati, Eds. *Market and Institutions in Economic Development*, pp. 117–137. London: MacMillan.

Modigliani, F. & Sterling, A. (1983). Determinants of private saving with special reference to the role of social security. Cross-country tests. In F. Modigliani & R. Hemming, Eds. *The Determinants of National Saving and Wealth*, pp. 24–55. London: MacMillan.

Pesaran, M.H. & Pesaran, B. (1991). *MICROFIT 3.0. An Interactive Econometric Software Package*. Oxford: Oxford University Press.

Rossi, N. & Visco, I. (1994). Private saving and government deficits. In A. Ando, L. Guiso & I. Visco, Eds. *Saving and the Accumulation of Wealth*, pp. 70–105. Cambridge: Cambridge University Press.

Samwick, A. (1994). The limited offset between pension wealth and other private wealth: implications of buffer stock saving (unpublished).

van den Noord, P. & Herd, R. (1993). Pension liabilities in the seven major economies. Working Paper No. 142. Paris: OECD.

Williamson, S.H. & Jones, W.J. (1983). Computing the impact of social security using the life cycle consumption function. *American Economic Review*, **73**, 1036–1052.

Appendix 1: Time-series estimates

As in Rossi and Visco (1994) an interesting starting point for a time-series analysis of the determinants of the private saving rate might be the equilibrium specification

$$-\ln(c/y^d) \approx s/y^d = \varphi_0 + \varphi_1 g + \varphi_2 r + \varphi_3\left(\frac{w^r}{y^d} + \theta\frac{w^s}{y^d}\right) \qquad (1.1)$$

where the variables are defined as in Section 3 of the text. In Rossi & Visco (1994) y^d and w^s, defined as usual, included (the former as the current flow, the latter as the current and future discounted flows) the pensions of current retirees (y^s) and (with a negative sign) the social security contributions (y^c) paid in front of the pension benefits of individuals currently at work.† In what follows, we benefit from the availability of the different components of w^s (from Beltrametti, 1995) and introduce an estimate of the net of tax social transfers component of disposable income, so that $w^s = w^{sw} + w^{sr} - w^{sc}$ and $y^d = y + y^s$ (disregarding net government transfers to firms which are assumed to have been already netted out from all income variables) where w^{sw} is the discounted stream of pension benefits for current workers, w^{sr} is the discounted stream of pensions for current retirees, w^{sc} is the discounted stream of social security contributions by current workers and we recall that y^s is the (net of tax) pension income flow. In equation (1.1), then, one should replace y^d with y and w^s with $w^{sg} = w^{sr} + w^{sw} = w^s + w^{sc}$. Alternatively, disposable income might be defined as $y + y^c$ and pension wealth as w^s. The former case, on which the following estimates are based, seems more appropriate for a pay-as-you-go system like the Italian one.‡

It should be noticed, however, that w^{sr} (as w^{sw}) is constructed under a set of specific assumptions which purport to mimic the formation of expectations by current retirees on their future net pensions. One could then include in the equilibrium relation only the component of social security wealth that relates to current workers (for which no substitute to w^{sw} is currently available), and use y^s as a proxy for the expected stream of current pensions and future benefits to the current retirees.§ We might then assume two alternative equilibrium specifications, that is

$$-\ln(c/y) = \varphi_0 + \varphi_1 g + \varphi_2 r + \varphi_3\left(\frac{w^r}{y} + \theta\frac{w^{sg}}{y}\right) \tag{1.2}$$

and

$$-\ln(c/y^d) = \varphi_0 + \varphi_1 g + \varphi_2 r + \varphi_3\left(\frac{w^r}{y^d} + \theta\frac{w^{sw}}{y^d}\right) + \varphi_4\left(\frac{y^s}{y^d}\right) \tag{1.3}$$

† In this paper, as in Rossi and Visco (1994), the income flows are net of taxes while the present values of current and future discounted flows are not.

‡ For a formal test of the alternative specification, see footnote † on p. 23.

§ Being aware of the double counting issue and to evaluate the robustness of our results, in our 1994 paper we also considered a specification such as (1.1) with w^s replaced by w^{sw} (but without allowing y^s to enter the regression with a different coefficient than y), to obtain results very close to those eventually reported in that paper, with, if anything, a slightly stronger effect of social security wealth on private consumption. Moreover, in conducting a simulation experiment on microeconomic data we fixed θ at 0·5 (as opposed to an estimate of 0·3 in the time series model) to account, again, for double counting.

where φ_4 is obviously a function of the w^{sr}/y^s ratio. In principle, proper representations of such long-run equilibrium relations might validly be estimated. However, it should be noticed that the ratio of w^{sr} to y^s increases from about 3 in the early 1950s to over 11 in the early 1990s. It then follows that relying on y^s as a proxy of w^{sr}, if may allow to consider a different expectations hypothesis, might induce substantial bias in the regression estimates to accommodate that rising trend. On these grounds we shall therefore disregard equation (1.3) in what follows.

However, it is easy to show that equation (1.2) may be rewritten as

$$-\ln(c/y^d) = \varphi_0 + \varphi_1 g + \varphi_2 r + \varphi_3\left(\frac{w^r}{y^d} + \theta\frac{w^{sg}}{y^d}\right) + \varphi_4\left(\frac{y^s}{y^d}\right) \qquad (1.4)$$

with $\varphi_4 = 1 - \varphi_0$.

If the above constraint is satisfied, equations (1.2) and (1.4) may be considered equivalent representations of consumer behaviour. In what follows we shall first focus on the somewhat simpler equation (1.2) and proceed, subsequently, to test the restriction $\varphi_4 = 1 - \varphi_0$.

Recent developments in the theory of cointegration have allowed to explicitly link economic equilibrium relationships between a set of variables such as those depicted by equation (1.2) with statistical models of those variables (Engle & Granger, 1987). In testing for the existence of cointegration, it is necessary to establish first that the individual series are $I(1)$ and then that there exists some non trivial function of them which is $I(0)$. The most common approach is to start testing for a unit root in the univariate representations of the individual series and then in the least squares linear combination. Dickey–Fuller (DF) and augmented Dickey–Fuller (ADF) statistics have been computed for the following variables:† (i) the log of the consumption to income ratio, $\ln(c/y)$, (ii) the private wealth to income ratio, w^r/y, (iii) the gross social security wealth to income ratio, w^{sg}/y, (iv) the real interest rate, r.

In the 1951–1993 sample, the hypothesis of a unit root at the 5% level could never be rejected for the variables listed above.‡ In a multivariate context, though, there may be a number of cointegrating regressions linking the saving rate, the real interest rate and the wealth to income ratios. We therefore revert to the

† All estimates reported in this Appendix have been obtained using Microfit 3.0 (Version 386, see Pesaran and Pesaran, 1991) or Pcfiml (Version 8.0, see Doornick and Hendry, 1994), where specific references are reported.

‡ The hypothesis of a unit root could, instead, be rejected for productivity and population growth.

TABLE 4 *Maximum likelihood cointegration analysis (equation 1.2, 1954–1993)*

Lag length in VAR (in years): 2; trended variables, no trend in DGP; exogenous variable: $\Delta\ ln(z_t)$

Non-zero eigenvalues of the stochastic matrix:
0·655, 0·485, 0·298, 0·169, 0·102

W test (H_0: $v \leq 4$; H_1: $v = 5$)	4·289
W test (H_0: $v \leq 3$; H_1: $v = 4$)	7·399
W test (H_0: $v \leq 2$; H_1: $v = 3$)	14·61
W test (H_0: $v = 1$; H_1: $v = 2$)	26·519
W test (H_0: $v = 0$; H_1: $v = 1$)	42·585*

Cointegration vector associated with the largest eigenvalue:

$\ln c$	$\ln y$	r	w^r/y	w^{sg}/y
$-1{\cdot}000$	0·966*	$-0{\cdot}155$	0·022*	0·023*

Homogeneity restriction: $\chi^2(1) = 1{\cdot}631$

* significant at the 95% level.

recent work of Johansen (1988), testing for the maximum number of cointegration vectors (v) and providing a maximum likelihood estimate of such vector(s) (Table 4). In designing the vector autoregression underlying the cointegration analysis it should be carefully noticed that focusing on the $\ln(c/y)$ variable entails unduly restrictive constraints on the short-run response of consumption to income. It seems wiser to let $\ln(c)$ and $\ln(y)$ enter the VAR separately and then test for long-run homogeneity of consumption to income. As in Rossi and Visco (1994), government expenditure is also allowed to play a role and hence $\Delta \ln(z_t)$ is allowed to enter the VAR as an exogenous variable.† As it turns out, the hypothesis of a single cointegration vector cannot be rejected at customary confidence levels.‡ Moreover, the hypothesis of long-run homogeneity of consumption and income is not rejected by the data.

† Integration tests for this additional variable show it to be $I(0)$.

‡ Statistically, it would be possible to further restrict the VAR to lag one. We prefer though a slightly overparameterized system and we stress that the (single) cointegrating vector implied by a VAR(1) turns out to be almost identical to those above. It may be interesting to notice that in our previous paper we opted for a VAR(4), imposing however short-run homogeneity between consumption and income. In that case the overparameterized VAR apparently accounted for the misspecification induced by the short-run homogeneity restrictions, delivering a single cointegration vector rather similar to the one shown in Table 4. Overlooking this point and simultaneously restricting the VAR's length and the short-run consumption to income response ends up misspecifying the cointegrating vector.

Finally, the stability of the cointegration space was explored, carefully inspecting the constancy of the cointegration rank (see Hansen & Johansen, 1993). In the two cases of a variable or a constant short-term dynamics, the cointegration rank was found to be remarkably stable.

In a relatively small sample, like the present one, an attractive alternative to the multivariate time-series methods used above is the estimation of a conditional error-correction model and the formal test of the significance of the adjustment coefficient. Assuming, therefore, on the basis of the results reported in Table 4, that the variables entering the long-run equilibrium relations are cointegrated, Table 5 presents the estimated linear error-correction models which have the general form of the theoretical model but contain additional lags and variables designed to shed light on the short-run impact of income and government expenditure as well as to take into account the possibility of some agents being constrained from their optimal consumption plan in any period thereby showing excess sensitivity of consumption to current income. In particular, our specification search allowed for a rich short-run dynamics which included, as in Rossi & Visco (1994), the lagged value of the dependent variable, $\Delta \ln c$, and the first differences of $\ln z$ and of $\ln y^d$. To allow for the possibility of different short-run responses with respect to y and y^s, besides the (current and lagged) first differences of $\ln(y^d)$, those of $\ln(y^s/y^d)$ were introduced with separate coefficients. The possible presence of current and lagged first differences of the real interest rate and of the rates of change of the wealth variables (in real terms) was also investigated. Beside the error-correction representation of equation (1.2), Table 5 also reports the estimates obtained with the saving rate specification (in terms of the standard definition of disposable income) of equation (1.4). Both the unrestricted and restricted ($\varphi_4 = 1 - \varphi_0$) estimates are presented.

In Table 5† besides the short-term effects coming from changes in y, y^s and z, the saving rate adjusts to its equilibrium level at a rate of about 0·5 per year; in the long-run equilibrium the saving rate is significantly influenced by the real rate of interest, the substitution effect ($-\varphi_2$) being about 0·3 (that is the negative of the coefficient of r, as estimated in Table 5, divided by the adjustment coefficient). In the long-run equilibrium the

† In Table 5 White's standard errors are reported under the coefficient estimates. The mis-specification tests statistics are distributed as follows: serial correlation $F(1, n-k-1)$, functional form $F(1, n-k-1)$, normality $\chi^2(2)$, heteroskedasticity $F(1, n-2)$, predictive failure $F(n_1, n-k)$, where n_1 is the number of observations in the shorter time period (1991–1993), additional regressors $F(8, n-k-8)$.

TABLE 5 *Linear error-correction representation (dependent variable: $\Delta \ln c_t$; sample: 1954–1993)*

	Equation (1.2)	*Equation (1.4) (unrestricted)*	*Equation (1.4) (restricted)*
Short-run terms			
$\Delta \ln(y^d)_t$	0·456 (0·044)	0·457 (0·045)	0·459 (0·042)
$\Delta \ln(y^s/y^d)_{t-1}$	0·093 (0·041)	0·093 (0·044)	0·098 (0·043)
$\Delta \ln z_t$	−0·482 (0·126)	−0·507 (0·142)	−0·476 (0·130)
$\Delta \ln z_{t-1}$	−0·332 (0·168)	−0·358 (0·174)	−0·340 (0·168)
$\Delta \ln c_{t-1}$	0·271 (0·078)	0·253 (0·083)	0·275 (0·081)
Error-correction term			
$\ln(c/y)_{t-1}$	−0·538 (0·088)		
$\ln(c/y^d)_{t-1}$		−0·516 (0·102)	−0·545 (0·091)
Long-run terms			
intercept	−0·126 (0·031)	−0·128 (0·057)	−0·157 (0·039)
r_{t-1}	−0·198 (0·066)	−0·184 (0·064)	−0·175 (0·063)
$(w^r/y)_{t-1}$	0·015 (0·003)		
$(w^r/y^d)_{t-1}$		0·016 (0·005)	0·018 (0·004)
$(w^{sg}/y)_{t-1}$	0·010 (0·002)		
$(w^{sg}/y^d)_{t-1}$		0·014 (0·004)	0·012 (0·002)
$(y^s/y^d)_{t-1}$		−0·546 (0·209)	−0·388 (0·063)
Misspecification tests			
$\bar{R}^2$	0·866	0·864	0·866
dw	2·030	2·079	2·010
$\hat{\sigma}100$	0·836	0·845	0·838
Serial correlation	0·147	0·273	0·095
Functional form	1·387	1·135	1·237
Normality	1·017	1·117	0·952
Heteroskedasticity	0·843	1·449	0·537
Predictive failure	0·911	0·729	0·894
Additional regressors	0·485	0·489	0·337
No. of observations (n)	40	40	40
No. of regressors (k)	10	11	10

Note: Additional regressors: $\Delta \ln(y^d)_{t-1}$, $\Delta \ln(y^s/y^d)_t$, Δr_t, Δr_{t-1}, $\Delta \ln w^r_t$, $\Delta \ln w^r_{t-1}$, $\Delta \ln w^{sg}_t$, $\Delta \ln w^{sg}_{t-1}$.

propensity to consume out of non-human wealth (φ_3) is about 0·03 and the coefficient of w^{sg} is between 65 and 90% that of w^r. Finally, the estimates appear to be remarkably stable, even when the very turbulent 1991–1993 period is included in the sample.†

† It should be observed that if we test for the equality of the coefficients of w^{sw} and w^{sr} by adding $(w^{sr}/y)_{t-1}$ to the regressors of equation (1.2) we cannot reject the null hypothesis (with a t-statistic equal to 0·938). Similarly, when we add $(w^{sc}/y)_{t-1}$ and $(y^c/y)_{t-1}$ to the regressors, to check for a possible bias in considering disposable income net of current contributions (rather than social security wealth net of the stream of current and future contributions) we cannot reject the assumption of no bias (with an $F(2, 28)$ statistic of 1·343).

TABLE 6 *Estimates of θ*

	Equation (1.2)		*Equation (1.4) (unrestricted)*		*Equation (1.4) (restricted)*
(1)	0·705 (0·095)	(2)	0·871 (0·337)	(3)	0·658 (0·087)
(4)	0·599 (0·099)	(5)	1·149 (0·538)	(6)	0·576 (0·104)
(7)	0·922 (0·125)	(8)	0·952 (0·345)	(9)	0·867 (0·114)

Note: in parentheses, White's standard errors. Equations (1), (2) and (3) as in Table 5. Equations (4), (5) and (6) as in Table 5 with additional regressors as indicated. Equations (7), (8) and (9) as (1), (2) and (3), respectively, with w^{sg}_{t-1} replaced by w^{s}_{t-1}.

The evidence from the ECM estimates then confirms that stemming from the analysis of cointegration. Table 6 presents a summary of the estimates of θ in equations (1.2) and (1.4) as can be obtained by the estimated coefficients presented in Table 5. Estimates of the respective standard errors are also provided. It is clear that this parameter appears to be sizeable and, as already observed, not far from one. To further check the robustness of this result a few attempts have been made, as shown in Table 6, to check whether restricting to zero the coefficient of w^{sc} would dramatically change the picture. Again we obtain very similar results, with, if anything, estimates of θ even closer to one.†

Since model (1.4) can be derived from model (1.2) if the restriction $\varphi_4 = 1 - \varphi_0$ is satisfied and since such restriction turns out not to be rejected by the data ($F(1, 29) = 0{\cdot}555$) it seemed safe to use in the text the specification of column (3) of Table 5, which is written in terms of the standard saving to disposable income ratio. The corresponding (non-linear) restricted least-squares and instrumental variables estimates are presented in Table 1 of the text.

† These estimates should be compared with those obtained in Rossi and Visco (1994), where the long-run equilibrium relation (1.1) was considered, and estimates of θ between one fifth and one third were obtained. New estimates of the same specification (with the revised time series and the extended sample considered in this paper) come out very close to the original ones, with θ turning out to be about one fourth. Most strikingly, however, adding the pension to income ratio, the estimate of θ rises dramatically, again being in the neighbourhood of one. The pension to income ratio provides then a correction to the double counting (and subsequent constraint) implicit in the definitions of disposable income and social security wealth considered in Rossi and Visco (1994).

Appendix 2: Cross-section estimates

The 1991 and 1993 SHIW provide data on consumption, income, wealth and socio-demographic characteristics from two samples of households representative of the Italian population. The 1991 and 1993 SHIW also report specific information on retirement behaviour, providing data on *ex post* and *ex ante* retirement age of current workers and (for 1991 only) on *ex post* and *ex ante* replacement rates. In order to evaluate the empirical exercise reported in Section 4, some data definitions and manipulations are worth recalling.

Both 1991 and 1993 SHIW do not provide information on the imputed flow of the consumption of durables. Consumption data therefore refer to consumption of non-durables and services. The normal earnings variable is supposed to purge current earnings from transitory earnings and age effects. In estimating adjusted normal earnings, defined as the annualized value of the recipient's human wealth, the procedure suggested by Ando, Guiso and Terlizzese (1994:208–209) has been followed. The household net worth is given by the sum of its net financial and real assets. The former are adjusted for non- and under-reporting as in Cannari *et al.* (1990).

Gross expected social security wealth is defined as in Jappelli (1995), as the stream of social security benefits of a person of a given age, expecting to retire at a given point in time and to die, with a given probability, anytime in the future. However, contrary to Jappelli we account for survivors' benefits. Since the future stream of social security benefits may be conventionally defined as a percentage of the agent's expected earnings at retirement, expected pension wealth crucially depends on the expected replacement rate and on the expected retirement rate.

Information on both quantities are provided by the 1991 SHIW as answers to the following questions: (i) "Consider the moment when you will retire. Setting your final monthly income before retirement equal to 100, what do you expect your first monthly pension to be?", and (ii) "When do you expect to retire?". Answers to the above questions are available only for 4400 and 8500 labour income recipients, respectively. Contrary to Jappelli (1995), to avoid sample selection biases arising from a non random distribution of non-responses in the population, reduced form equations for expected replacement rate and expected retirement age were estimated on the respondents subsample and fitted values imputed to the overall sample of labour income recipients, thereby implicitly accounting for potential endogeneity of the social security wealth variable arising from endogeneity of expected retirement age and replacement rate.

In particular, expected replacement rates at the individual level turned out to be affected significantly by variables defining the social security rules applying to the specific individual such as sex, number of years worked (proxying for the number of years of contributions), occupation and sector of occupation, as well as their interaction with other variables at the individual level such as education. Expected retirement age, instead, proved to be strongly dependent, along with the individual history of the agent, on his social environment (location of the household, number of children and income recipients in the household, and the like). Interestingly, the expected retirement age turned out to depend, rather clearly, on the expected replacement rate with a five percentage point decrease of the pension to earnings ratio leading to a one-year shift forward of the expected retirement age.

Alternatively, to check the previous computations and cross-check the time-series aggregate estimates discussed in the main text, gross social security wealth was computed assigning each worker to a specific social security fund and defining consequently age eligibility requirements, reference wage and period for computing pensionable earnings, accrual rates and contribution requirements. The task was made rather difficult by the complexity of the Italian social security system, by the presence of a large number of social security funds for various categories of employees and self-employed and by the uncontrolled evolution of the legislation in the last decades. It proved necessary to revert to a number of simplifying assumptions of which the most important ones related to the timing of retirement, assumed equal to the earliest possible one.

For the 1991 sample household social security wealth averaged at about 310 and 340 million lire (for the measures based on expectations and legislation, respectively), a level largely comparable with Beltrametti's (1995) aggregate time-series figures used in Section 3. The difference between the two measures of social security wealth was mostly concentrated in those occupations and sectors where early retirement is allowed but is apparently not accounted for in agents' expectations. In particular, the simple correlation coefficient between the two quantities is about 0·95 in the 1991 sample.

Table 7 presents a summary of the estimates based on equation (2) in the text. Since both intercept and slopes of equation (2) shift with socio-demographic characteristics, estimated coefficients refer to a specific household as described in the notes to Table 7.

TABLE 7 *Cross-section estimates (1991 SHIW, least squares estimates, dependent variable: negative of log of non-durables consumption to normal net earnings ratio)*

	Social security wealth based on	
	Legislation	*Expectations*
w^r/y^n	−0·0190 (0·0030)	−0·0174 (0·0030)
w^{sg}/y^n	−0·0163 (0·0011)	−0·0174 (0·0012)
$\ln(y^n)$	0·3366 (0·0140)	0·3532 (0·0158)
Constant	−2·9847 (0·1474)	−3·1602 (0·1717)
Purchase of life insurance ($\times 100$)	−0·0024 (0·0005)	−0·0023 (0·0005)
Contributions to pension funds ($\times 100$)	−0·0030 (0·0006)	−0·0029 (0·0006)
Plus socio-demographic dummies and interaction terms		
No. of observations	5628	5628
R^2	0·5360	0·5440

Notes: (i) in parentheses, White's standard errors, (ii) the sample excludes households with pension income only, (iii) the reported estimates and standard errors refer to a household living in a small town in Northern Italy, formed by two adults, both income recipients, and two children, and with the head of the household in his late twenties or early thirties.