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The Risk Premium on Italian Government Debt, 1976-88

Prepared by Carlo Cottarelli and Mauro Mecagni*

Authorized for Distribution by Erich Spitäller

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Abstract

This paper considers the behavior of the yield differential between government and nongovernment bonds in Italy between 1976 and 1988. It is shown that the trend increase of the differential observed in this period was significantly influenced by the deterioration of public finances, as reflected both by an increase in the relative supply of government with respect to nongovernment paper and by a worsening of selected default risk indicators. In addition, the effect of relative supply factors was found to be statistically more robust and quantitatively more important than the effect of risk indicators in explaining the movements of the yield differential.

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Summary

This paper analyzes the behavior of the real yield differential between fixed coupon bonds issued during 1976-88 by the Italian Government and by Special Credit Institutions, the largest nongovernment issuer of bonds in Italy. The yield differential, computed on bonds of equal maturity, turned from negative values in the 1970s to positive values in the 1980s. This behavior is interpreted in light of a simple model relating changes in relative yields between imperfectly substitutable assets to changes in relative supply and in relative default risk. The model is extended to allow for the investment requirement that forced commercial banks to hold a large share of their assets invested in Special Credit Institutions during 1973-87 and for possible differences in the propensity of banks and the nonbank public to purchase government paper.

The model is tested on a panel of 457 observations on yield differentials, as well as on quarterly averages of the same observations by generalized least squares. The trend increase in the differential observed in the sample period was significantly influenced by the deterioration of public finances, as reflected in an increase in the relative supply of government paper and in a worsening of selected indicators of default risk on public debt. As to the relative importance of supply factors vis-a-vis risk indicators, the paper finds that the effect of the former appears to have been quantitatively larger and statistically more robust with respect to changes in the specification of the estimated equation and in the estimation method.

These findings have implications for fiscal policy and for debt management. First, the evidence that the yield differential between government and nongovernment paper rose as a result of increasing fiscal imbalances implies that a policy of fiscal adjustment should benefit from a reduction both in the general level of interest rates and in the additional yield now paid by the Government on its debt. The predominance of supply factors, and, in general, of stock-over-flows indicators of public finance imbalances, implies that, in order to reap this reward, the fiscal effort must be sustained over time, so as to allow for the proper adjustment in the stock of financial assets. Second, the finding that the rise in the cost of public debt was influenced by increased pressure of government paper on Italian financial markets indicates that the interest rate burden could be reduced by increasing the share of debt sold to nonresidents, whose "appetite" for Italian government paper may not be entirely satisfied. Third, measures to increase the efficiency of the secondary market for Treasury paper in Italy may also be instrumental in bringing about a decline in the cost of debt, as they would reduce the component of the yield differential unrelated to default risk.

I. Introduction

Government liabilities are considered in many countries the risk free assets *par excellence*. Financial market studies take therefore the yield on Government bonds or Treasury Bills as a benchmark against the yield of assets issued by private agents and explain the frequently observed yield differential with the existence of default risks on the latter (see, for example, Barrett, Heuson and Kolb (1986), Van Horne (1970), Ferson (1989), Ferson, Kandel and Stambaugh (1987), Fisher (1959), Yawitz (1978)). This paradigm needs however to be revised in conditions of high and persistent imbalances in public sector finances, leading to rapid and substantial accumulation of public debt. As public debt increases, the market may start wondering how the government intertemporal budget constraint will be respected and the possibility of "default" on government debt may be explicitly considered.

A case in point may be provided by the Italian experience of the last 15 years. Between the middle of the 1970s and the end of the 1980s public debt rose in relation to GDP from 50 to 100 percent, with an increasing share of debt financed at rising market rates. With the burden of interest payments becoming progressively heavier (reaching 80 percent of the overall deficit in 1989) and with the average interest rate on the debt approaching the GDP growth rate, increasing attention is given to the existence of a "debt problem" and to the possible use of extraordinary measures to solve it, including forms of partial repudiation. Most economists have warned that those measures would not be at present feasible in Italy, at least not before primary deficits are eliminated, due to the cost in terms of reputation (see, for example, Alesina (1988) and Spaventa (1988)). All the same, discussions of debt consolidation, recourse to administrative measures (e.g. portfolio constraints on the banking system) and capital levies on public debt recurred in the policy arena in the last few years. 1/

Has the risk of default, a term by which henceforth we include also milder forms of repudiation or consolidation, already been discounted by the market? Is that risk, at least partially, responsible for the current high level of real rates on Italian government paper? Is it perceived as related to the size of the deficit and/or the debt? Answers to these questions are clearly important to debt management and, in general, to the formulation of fiscal adjustment programs. Indeed, if the high real rates paid by the Italian government are due, at least

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1/ Interest in previous examples of partial debt repudiation in Italy has also revived; see Alesina (1988) and De Cecco (1989). As recalled by Spinelli and Vismara (1989), the Economic Commission created in 1945 by the Constituent Assembly discussed the possibility of introducing in the new Republican Constitution the prohibition of debt repudiation; the proposal was, however, soon rejected.

in part, to an "issuer specific" risk premium, credible and sustained commitment to fiscal adjustment should lower this risk and lead to a relatively fast reduction in interest rates.

Despite the importance of the issue, evidence of the existence of a default risk premium on the Italian government debt is almost exclusively anecdotal. There is some evidence of a positive effect of increasing deficits on the level of real rates (see CER (1988), Modigliani and Jappelli (1988)); but this evidence is no proof of an increasing issuer specific risk premium as rising deficits may affect real rates without affecting the risk premium. 1/

The existence of an increasing risk premium on government paper should be signaled by a rise in the yield differential between government and nongovernment assets comparable in terms of currency denomination, maturity, and liquidity. Is there evidence of such a rise? Given the incompleteness of the Italian financial market and its institutional constraints, answering this question is not an easy task. 2/ In this paper we consider the yield differential between medium-term government paper (BTPs) and the fixed coupon bonds issued by Special Credit Institutions (henceforth SCI), the major nongovernment issuer of bonds in

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1/ Indeed, if Ricardian equivalence does not hold, an increase in the deficit (and/or the debt) will tend to increase aggregate demand; for given resources, and/or supply of money, general equilibrium requires an increase in the real interest rate. In addition, increasing deficits and debt could affect real rates by increasing the "currency specific" risk premium: as an accumulation of debt may increase the likelihood of future monetization and inflation, investors free to shift their investment across different countries will move away from the risky currency, pushing up interest rates on all assets (public and private) denominated in the risky currency. While the existence of a currency component in the risk premium on Italian government debt is an obvious possibility, in this paper we focus exclusively on the issuer specific component.

2/ Alesina, Prati and Tabellini (1989) consider in this respect the differential yield between treasury bills and bank CDs; this series is, however, available only for a few years, as bank CDs were introduced only in the course of the 1980s and had initially a limited development. Moreover, bank CDs have been, especially in the past, mainly a substitute for other forms of deposits, more than a direct competitor of treasury bills. It must also be recalled that the market for private commercial paper in Italy is extremely thin and no information is available on the corresponding interest rates. Another possibility would have been to compare the yield of bonds issued by the Italian and other governments on the Euromarkets. However, the default risk on the Italian Eurobonds, a rather small portion of the Italian government debt (3.3 percent at the end of 1989), is likely to be substantially different from the risk on domestic debt, which may be more easily subject to administrative measures. Indeed, the high rating given to the Italian government by main rating agencies refers exclusively to the bonds issued on Euromarkets; see, for example, Moody's (1989), p. 1.

Italy. 1/ We find that this differential, computed on bonds of equal maturity, turned from negative values in the 1970s to significantly positive values in the 1980s. Unfortunately, this finding does not necessarily imply an increasing risk premium, as changes in relative supply can explain changes in the yield differential between assets which are imperfect substitutes. 2/ The hypothesis that BTP and SCI bonds behaved as imperfect substitutes during the period under consideration is sustained here by three arguments: first, during most of the period, a portfolio investment requirement forced banks to purchase SCI bonds, thus reducing their yield and, possibly, their yield variance. Second, while the relative default risk may not have changed in the period, over a constant difference in the level of risk is per se to induce imperfect substitutability. Third, and most important, the imperfections of Italian bond markets are likely to have deeply influenced the relative liquidity of the two assets and the variance-covariance matrix of their returns, particularly their yield correlation. In conclusion, the interpretation of the observed change in the yield differential requires a comprehensive econometric analysis allowing for separate effects of default risk indicators, of relative supply effects and of institutional constraints.

The plan of the paper is the following. In Section II a simple portfolio choice model is used to derive testable propositions on the determinants of yield differentials. The theoretical framework allows for different propensities in the purchase of government paper by banks and by the nonbank private sector and for the investment requirements imposed for many years on bank portfolios. Section III presents the results of the econometric analysis of the behavior of the yield differential between BTPs and SCI fixed coupon bonds; GLS estimates of the relation between yield differential adjusted for maturity, relative asset supply and risk indicators are presented both for a panel of 457 quarterly observations between 1976-IV and 1988-IV and for aggregate data. The main conclusion of the analysis (Section IV) is that the trend increase in the differential was significantly influenced by the deterioration of public finances in the last 15 years, reflected both by the increase in the relative supply of BTPs and by the deterioration of selected risk indicators. While the breakdown of the respective impact

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1/ SCI are financial intermediaries specialized in long-term credit for industrial and real estate investment. Although a large share of these institutions are public bodies, they are largely independent from the government, they have their own capital endowment and legal status, their assets are represented mainly by loans to the private sector, and their bonds are rated independently from those of the government.

2/ As discussed in Section II, following the mean-variance approach to portfolio choice, the increase in the yield differential required to accommodate a change in relative supply may also be interpreted as an increasing risk premium because it offsets the utility loss occurring when investors move away from the minimum variance portfolio. Clearly, however, this risk premium has a different nature from a default risk premium.

of relative supply and default risk indicators may be affected by collinearity problems, the contribution of relative supply appears to have been quantitatively more important.

II. A Model for the Interest Rate Differential

1. Basic features

A model for the yield differential between BTPs and SCI bonds can be derived from the mean-variance approach to portfolio choice. In this context, demand curves for different assets can be derived as follows; let:

$$U = U(W^e, S_W^2) \qquad U_1>0 \quad U_2<0 \tag{1}$$

be the utility function of a representative investor, assumed to depend positively on the expected value of total real wealth at the end of the period (W^e) and, negatively, on the variance of wealth (S_W^2). 1/ Given N assets in which wealth can be invested, call Q the Nx1 decision vector including the shares of wealth invested in each asset and R the Nx1 vector including one plus the real yield of each asset; R is a vector of random variables distributed as:

$$E(R) = R^e \qquad Var(R) = \Omega \tag{2}$$

The investor maximizes (1) subject to the budget constraint:

$$Q'J = 1 \tag{3}$$

where J is a column vector of ones. The solution to this problem (see Appendix 1) is given by:

$$Q = (\gamma/\phi)\Omega^{-1}[I-J(J'\Omega^{-1}J)^{-1}J'\Omega^{-1}]R^e + \Omega^{-1}J(J'\Omega^{-1}J)^{-1} \tag{4}$$

where $\gamma = -W_{-1}U_1$, I is a NxN identity matrix, W_{-1} is the beginning of period wealth and $\phi = 2\,W^2_{-1}U_2$. If we consider the two asset case, system (4) reduces to the following demand equations for the two assets shares q_g and q_p:

$$q_g = \frac{(r_g^e - r_p^e) + \theta(s_{pp} - s_{gp})}{\theta(s_{gg} + s_{pp} - 2s_{gp})} \tag{5}$$

$$q_p = \frac{(r_p^e - r_g^e) + \theta(s_{gg} - s_{gp})}{\theta(s_{gg} + s_{pp} - 2s_{gp})} \tag{6}$$

1/ We also assume that the second order derivatives (U_{11} and U_{22}) are negative and that $U_{11}U_{22}-U_{12}^2>0$; under these conditions the investor is said to be "risk averse in a mean-variance sense" (Merton, 1982).

where $\theta = (-2W_{-1}U_2)/U_1>0$, r_g^e and r_p^e are the expected real yields on asset g (BTPs) and asset p (fixed coupon SCI bonds) and s_{gg}, s_{pp} and s_{gp} are, respectively, the variances of asset returns on assets g and p and their covariance. 1/ More compactly, we can also write:

$$q_g = \alpha_0 + \alpha_1(r_g^e - r_p^e) \tag{7}$$

$$q_p = \beta_0 + \beta_1(r_g^e - r_p^e) \tag{8}$$

In (7) - (8) the portfolio shares of government and nongovernment paper depend on the expected real differential yield and on parameters which are functions of the variance-covariance structure of the returns on the two assets and on θ, a parameter that, in a mean-variance context, reflects the risk aversion of investors (Dornbusch, 1983). 2/ Provided the asset yields do have not equal variance and covariance (i.e., unless $s_{gg} = s_{pp} = s_{gp}$), the elasticity of portfolio shares with respect to the yield differential is finite, i.e. the assets are imperfect substitutes. The imperfect substitutability of financial assets, often assumed as working hypothesis in empirical research on financial markets, 3/ appears appropriate also in the case of BTPs and SCI bonds. Apart from the effect of the investment requirement (see below) and of possible differences in default risk, the characteristics of both primary and secondary market for the two types of bonds are very different. BTPs issues are largely advertised and purchases can be made not only through the banking system but also through the Bank of Italy; on the other hand, their secondary market has been, until recently, highly imperfect. 4/ SCI bonds are mainly traded through commercial banks to which SCI are usually connected by ownership links. Thus, while a formal secondary market for SCI bonds is equally inefficient, their

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1/ Moving from an N asset to a two asset portfolio requires the separability of portfolio decisions. Our model assumes that agents apply mean-variance maximization to the portfolio represented by BTPs and SCI fixed coupon bonds; the relative size of this portfolio is here considered as exogenous, and possibly derived from previous application of mean-variance analysis to the choice between money, short term paper, indexed bonds and other financial (and real) assets. This stepwise decision process is also assumed to describe the behavior of both bank and nonbank sectors in the quarterly econometric model of Bank of Italy (see Banca d'Italia (1986)).

2/ The usual constraints among the coefficients of (7) and (8) hold. It can be shown, for example, that $\alpha_1>0$ and $\alpha_1=-\beta_1$. Moreover, an increase in the variance of each asset reduces, ceteris paribus, its demand.

3/ See, for example, Jaffee (1975), Friedman (1977), Masson (1978), Backus et al. (1980), Roley (1983) and, in the case of Italy, Banca d'Italia (1986 and 1988 b).

4/ See Banca d'Italia (1988 a). In order to improve its efficiency, the secondary market for government paper was reformed in May 1988 with the institution of a screen-based system of negotiations and the introduction of a group of primary dealers.

liquidity may have been enhanced by the intervention of commercial banks. All these differences, and more generally the imperfection of the Italian secondary bond market, explain why in the past price changes of BTPs and SCI bonds were not perfectly correlated. 1/

The imperfect substitutability hypothesis has an obvious implication for the analysis of market yield differentials. Deferring aggregation problems to Section II(2), equation (7) can be considered as the market demand for BTPs (expressed in terms of portfolio shares). By inverting equation (7) (as, for example, in Fair and Malkiel (1971), Modigliani and Sutch (1967), and Frankel (1983)), and assuming equilibrium in the bond market, we derive an equation for the expected yield differential as a function of the relative supply of different assets: 2/

$$\delta^e = r_g^e - r_p^e = -(\alpha_0/\alpha_1) + (1/\alpha_1)q_g \tag{9}$$

As $\alpha_1 > 0$, the expected yield differential δ^e required by the market increases with q_g. This increase is frequently interpreted as a "relative supply effect," although, in a mean-variance context, it should be interpreted as a risk premium as it corresponds to the increase in the expected yield differential required to move investors away from the minimum variance portfolio, i.e., to accept a higher risk; indeed, δ^e is zero only when the relative market supply of the two assets corresponds to the minimum variance portfolio. 3/

2. Aggregation and institutional constraints

Application of the previous approach to the analysis of the yield differential between BTPs and SCI bonds requires the consideration of

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1/ While direct information on the price changes of BTPs and SCI bonds of the same maturity is not available, the correlation between the average price changes of the outstanding stock of the two types of bonds appears relatively low; for example, the correlation coefficient of monthly price changes was 0.74 between 1977 and 1982 and 0.81 between 1983 and 1988; in some years (1978 and 1980) the coefficient was close to 0.5. Indeed, as observed by Penati and Formentini (1989), with reference to different government bonds the low yield correlation observed in the short run on Italian markets leaves room to profitable portfolio diversification.

2/ Given the relation between the parameters of the model, equation (9) can be equivalently derived from equation (8) instead of equation (7).

3/ The minimum variance portfolio is obtained by minimizing S_W^2 subject to (3). In the two asset case it is given by:

$$q_g = (s_{pp} - s_{gp})/ (s_{gg} + s_{pp} - 2s_{gp}).$$

By substitution of this expression in (5), $\delta = 0$.

two additional factors. The first relates to an aggregation problem. In Italy, both bank and nonbank sectors have in the past purchased large amounts of bonds, but the composition of demand substantially changed over time: at the end of 1976, 70 percent of the outstanding stock of medium and long-term bonds (excluding those in the portfolio of the Central Bank) was held by the banking system. At the end of 1988, this share had declined to 30 percent. As the propensity to buy government bonds may be different between banks and nonbank public, ignoring the change in the composition of demand may bias the econometric results.

To illustrate this point, call G^h and G^b the total demand of government bonds of nonbank public (henceforth "households", for simplicity) and of banks. Under the assumption that both banks and households are characterized by asset demand functions such as (7), 1/ the total demand for BTPs can be expressed as:

$$G = G^h + G^b = (\alpha_0^h + \alpha_1^h \delta^e)B^h + (\alpha_0^b + \alpha_1^b \delta^e)B^b \tag{10}$$

where B^h and B^b are the total holdings of bonds by households and banks. Dividing (10) by $B(= B^h + B^b)$ we obtain:

$$q_g = \frac{G}{B} = (\alpha_0^h + \alpha_1^h \delta^e)h + (\alpha_0^b + \alpha_1^b \delta^e)(1 - h) \tag{11}$$

where $h = B^h/B$ is the share of total bonds on the market held by the households sector. Solving (11) for δ^e yields:

$$\delta^e = -\frac{[\alpha_0^h h + \alpha_0^b(1-h)]}{[\alpha_1^h h + \alpha_1^b (1-h)]} + \frac{1}{[\alpha_1^h h + \alpha_1^b (1-h)]} q_g \tag{12}$$

which shows that, unless $\alpha_0^h = \alpha_0^b$ and $\alpha_1^h = \alpha_1^b$, or unless h is constant, the parameters of (12) will change over time, according to the distribution of bonds between households and banks. If we assume that $\alpha_1^b = \alpha_1^h = \alpha_1$, but we allow for different intercepts in the demand functions $(\alpha_0^h \neq \alpha_0^b)$, equation (12) becomes:

$$\delta^e = -\frac{\alpha_0^b}{\alpha_1} - \frac{(\alpha_0^h - \alpha_0^b)}{\alpha_1} h + \frac{1}{\alpha_1} q_g \tag{13}$$

1/ We are also assuming homogeneity of behavior and wealth endowments within each of the two groups with the exception of the distinction between "constrained" and unconstrained" banks introduced below. Alternatively, if we assume that individuals within each group have different wealth endowments and risk aversion, the parameters of equation (10) for both households and banks will reflect the averages of risk aversion coefficients weighted by the individual wealth endowments (Dornbusch (1983)).

Under these hypotheses, the level of the differential is also a function of the distribution of the stock of bonds between banks and nonbank public.

The second factor that has to be incorporated in the model is the 1973-86 portfolio investment requirement forcing commercial banks to hold a large share of their assets invested in SCI bonds as a measure to enhance investment growth. 1/ The existence of a portfolio constraint, which at a given time may not be binding for all banks, can be incorporated in the model as follows. Let N_c be the number of banks for which the constraint is binding and N_u the number of unconstrained banks, with $N = N_c + N_u$. The stock of SCI bonds (i.e. of "nongovernment bonds") held by constrained banks (P_c) is given by:

$$P_c = p_c N_c = kd_c N_c \tag{14}$$

where p_c, the average stock of SCI bonds held by constrained banks, is a proportion (k) of their deposits (d). The stock of SCI bonds held by unconstrained banks (P_u) is instead determined by the portfolio model previously discussed:

$$P_u = p_u N_u = (\beta_0^b + \beta_1^b \delta^e) b_u N_u \tag{15}$$

where p_u and b_u are, respectively, the average holdings of SCI bonds and the average holdings of total bonds by unconstrained banks. As P_c and N_c are not observed, the number of constrained banks is assumed to be a fixed proportion (γ) of the ratio between the minimum required amount of SCI bonds (P^*) and the total amount of SCI bonds held by the banking system (P^b):

$$N_c = \gamma(P^*/P^b)N \qquad 0 < \gamma < 1 \tag{16}$$

The total demand of SCI bonds by banks is then given by:

$$P^b = P_c + P_u = \gamma p_c \frac{P^*}{P^b} N + (\beta_0^b + \beta_1^b \delta^e) b_u (1 - \gamma \frac{P^*}{P^b}) N \tag{17}$$

1/ The regulation on the investment requirement changed over time. From June 1973 to September 1974, banks had to invest in SCI bonds a fixed proportion of the stock of bank deposits; between September 1974 and June 1978, the outstanding required stock of bonds had to be incremented by a fixed proportion of the increase in deposits. In July 1978 the minimum requirement on the bonds issued by SCI granting credit for industrial investment (on which this paper focuses) was frozen: banks were requested only to replace the maturing bonds. Between January and December 1983 only 50 percent of maturing bonds had to be replaced. As of January 1984 the replacement constraint was removed and, finally, in January 1987 the requirement was lifted. Initially the constraint also forced banks to buy a limited amount of government bonds. The bulk of the investment requirement, however, always referred to SCI bonds.

Under the additional simplifying hypothesis that total bond holdings are on average equal for constrained and unconstrained banks (i.e., that $b_u N = b_c N = B^b$), 1/ and by substitution of (14) into (17) we get:

$$P^b = \gamma k d_c \frac{P^*}{P^b} N + (\beta_0^b + \beta_1^b \delta^e)\,(1-\gamma\frac{P^*}{P^b})B^b \qquad (18)$$

If we further assume that the average deposit size of constrained and unconstrained banks is the same, then $d_c N = D$, where D is total bank deposits. As for the whole system $P^* = kD$, (18) can then be written as:

$$P^b = \gamma\frac{(P^*)^2}{P^b} + (\beta_0^b + \beta_1^b \delta^e)(1-\gamma\frac{P^*}{P^b})B^b \qquad (19)$$

where all variables are now observable.

If we now consider market equilibrium, on account of equation (19), the market yield differential can be expressed (Appendix II) as:

$$\delta^e = -\frac{\alpha_0^b}{\tilde{\alpha}_1} - \frac{(\alpha_0^h - \alpha_0^b)}{\tilde{\alpha}_1} h + \frac{1}{\tilde{\alpha}_1} q_g + \frac{\gamma}{\tilde{\alpha}_1} \frac{P^*}{B} \frac{P^*}{P^b} - \frac{\gamma}{\tilde{\alpha}_1} (1-\alpha_0^b) \frac{P^*}{P} (1-h) \qquad (20)$$

where $\tilde{\alpha}_1 = \alpha_1 [1 - \gamma \frac{P^*}{P^b} (1-h)] > 0$, or, more simply:

$$\delta^e = \phi_0 + \phi_1 h + \phi_2 q_g + \phi_3 \frac{P^*}{B} \frac{P^*}{P^b} + \phi_4 \frac{P^*}{P} (1-h) \qquad (21)$$

With respect to equation (13), two additional terms enter equation (21) with expected coefficients $\phi_3 > 0$ and $\phi_4 < 0$. The fourth term of (21) shows that an increase in the minimum investment requirement of SCI bonds raises the yield differential between government and SCI bonds. This effect is larger, the larger is the investment requirement in relation to the total demand of SCI bonds by banks (P^*/P^b) and in relation to the size of the bond market (P^*/B). The last term of (21) shows that the previous effect is partially moderated by the decline in the unconstrained demand for SCI bonds. 2/ A second relevant point is that the coefficients of equation (21) are a function of P*. In particular, as $P^* \neq 0$, $\tilde{\alpha}_1 < \alpha_1$, the coefficients exceed (in absolute terms) the corresponding coefficients of equation (13). 3/ Since in our empirical estimates we do not allow for the dependence of the coefficients on P*, the estimated parameters will reflect the average effect of the

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1/ This is indeed a restrictive assumption, since the constraint may alter not only the composition but also the size of the constrained bank portfolios, which may be higher than otherwise would be.

2/ In other words, an increase in the minimum investment requirement does not imply an equal increase in the demand for SCI bonds as some of the required increase is simply satisfied by the outstanding stock of SCI bonds.

3/ This is a known effect of constraints (see, for example, Angeloni (1985) with reference to the effect of credit ceilings in Italy).

regressors on the differential in presence of the investment requirement. Their values may therefore exceed those prevailing in the absence of the requirement.

3. Default risk

In the presence of a positive default probability, the expected yields on the two assets can be expressed as the sum of expected yields in the absence of default (r_g^{Ne} and r_p^{Ne}) minus the expected cost of default:

$$E(r_g) = r_g^{Ne} - p_g t_g \tag{22}$$

$$E(r_p) = r_p^{Ne} - p_p t_p \tag{23}$$

where p_g and p_p are the probabilities of default and t_g and t_p are the costs of default, respectively, for BTPs and SCI bonds. 1/ Recalling that $\delta^e = E(r_g) - E(r_p)$, and by substitution of (22) and (23) into (21), we obtain:

$$\delta = \phi_0 + \phi_1 h + \phi_2 q_g + \phi_3 \frac{P*}{B} \frac{P*}{p^b} + \phi_4 \frac{P*}{P}(1-h) + p_g t_g - p_p t_p \tag{24}$$

Equation (24) shows that δ (= $r_g^{Ne} - r_p^{Ne}$), the yield differential computed under the hypothesis of no default, is a function of the expected differential cost of default ($p_g t_g - p_p t_p$). The default probabilities p_g and p_p are not observed; while we assume that p_p (the probability of default on SCI bonds) was constant in the period under consideration, 2/ we correlate the probability of default on BTPs to a set of default risk indicators that could trigger confidence crises. Recent research suggests two variables that in the past may have significantly affected investors' confidence in Italian government paper: the maturity of the debt 3/ and the amount of debt that comes to maturity in each period. 4/ It is plausible to add to these variables two fiscal policy

1/ Note that r_g^{Ne} and r_p^{Ne} are expected values of stochastic yields because default risk is not the only source of uncertainty of bond yields.

2/ This seems to be a reasonable assumption as the performance of SCI remained satisfactory in terms of profitability and capital adequacy throughout the period.

3/ Alesina, Prati and Tabellini (1989) present a theoretical model in which "the maturity structure of public debt determines the likelihood of a confidence crisis on the debt: the shorter and the more concentrated is the maturity, the more likely is the crisis."

4/ In Giavazzi and Pagano (1989) a model is presented in which "the probability that the authorities will withstand a confidence crisis is critically affected by the extent to which they have to appeal to the market at each given date to roll over public debt. This depends on three factors: the amount of debt outstanding, its average maturity and the time pattern of maturing debt."

indicators on which public opinion usually focuses: the deficit to GDP ratio and the debt to GDP ratio. 1/ We therefore assume that:

$$p_g = \lambda_0 + \lambda_1 m + \lambda_2 \frac{MA}{D} + \lambda_3 \frac{DF}{Y} + \lambda_4 \frac{D}{Y} \qquad \lambda_1 < 0,\ \lambda_2 > 0,\ \lambda_3 > 0,\ \lambda_4 > 0 \quad (25)$$

where m is the average maturity of government debt, MA is the amount of debt coming to maturity in the period, DF is the deficit, Y is GDP and D is the stock of debt. By substitution of (25) into (24) we obtain an equation that relates the yield differential to the distribution of bonds between households and banks, to the relative supply of government paper with respect to nongovernment paper, to the investment requirement, and the effect of "risk factors." 2/

$$\delta = \phi_0' + \phi_1 h + \phi_2 q_g + \phi_3 \frac{P^*}{B} \frac{P^*}{P^b} + \phi_4 \frac{P^*}{P}(1-h) + \phi_5 m + \phi_6 \frac{MA}{D} + \phi_7 \frac{DF}{Y} + \phi_8 \frac{D}{Y} \quad (26)$$

where $\phi_0' = \phi_0 + t_g \lambda_0 - p_p t_p$, $\phi_5 = t_g \lambda_1$, $\phi_6 = t_g \lambda_2$, $\phi_7 = t_g \lambda_3$ and $\phi_8 = t_g \lambda_4$.

The sign of the coefficients is expected to be positive for ϕ_2, ϕ_3, ϕ_6, ϕ_7, and ϕ_8, and negative for ϕ_4 and ϕ_5, while the sign of ϕ_1 is not determined a priori depending on the relative propensity of households and banks to purchase government paper.

So far, the discussion has been cast for simplicity in a static framework; yet several reasons make it plausible to assume that the relation between yield differential and its determinants is dynamic. First, lags in (26) may be justified in case expectation formation makes use of past information; for example, not only the current, but also past values of the debt to GDP ratio could be considered as "risk indicators". Second, and most important, lags in the demand for government bonds, due for example to adjustment costs, should be reflected also in the equation of the yield differential. The simplest way to include adjustment lags in the demand for government bonds equation would be through a partial adjustment mechanism; in this case equation (7) should include the lagged value of q_g and, consequently, equation (26) should also include q_{g-1}, with expected negative sign and a coefficient smaller than ϕ_2 in absolute value.

1/ See, with respect to the effect of debt accumulation on confidence, Spaventa (1988, p. 16) and Treasury Committee on Financial Assets, Public Debt and Monetary Policy (1987, p. 304).

2/ We have already observed that, strictly speaking, what we call "supply effect" could be seen as a component of risk premium, and what we call "risk factors" should be seen as factors affecting the expected return of government bonds, not its variance (i.e., the portfolio risk). In what follows, however, we prefer to maintain the more immediate, albeit less precise, terminology used in the text.

III. Empirical Analysis

1. Stylized facts

According to the model discussed in Section II, the deterioration of Italian public finances in the course of the 1980s should have been accompanied by a rise in the yield differential, partly as a result of a "relative supply effect" and partly as a result of "increased default risk," as a consequence of the higher debt and deficit to GDP ratios.

The evidence depicted in Chart 1, prima facie, supports the theoretical model. The net of tax differential between the average yield of the outstanding stock of BTPs and of fixed coupon SCI bonds 1/ turned from negative values in the 1970s to positive values in the 1980s (top panel, solid line). The trend increase was accompanied both by a rise in the relative supply of BTPs (central panel) and by a deterioration of some of the above mentioned "risk indicators" (Chart 2). The differential dropped in 1987-88, following the removal of the portfolio

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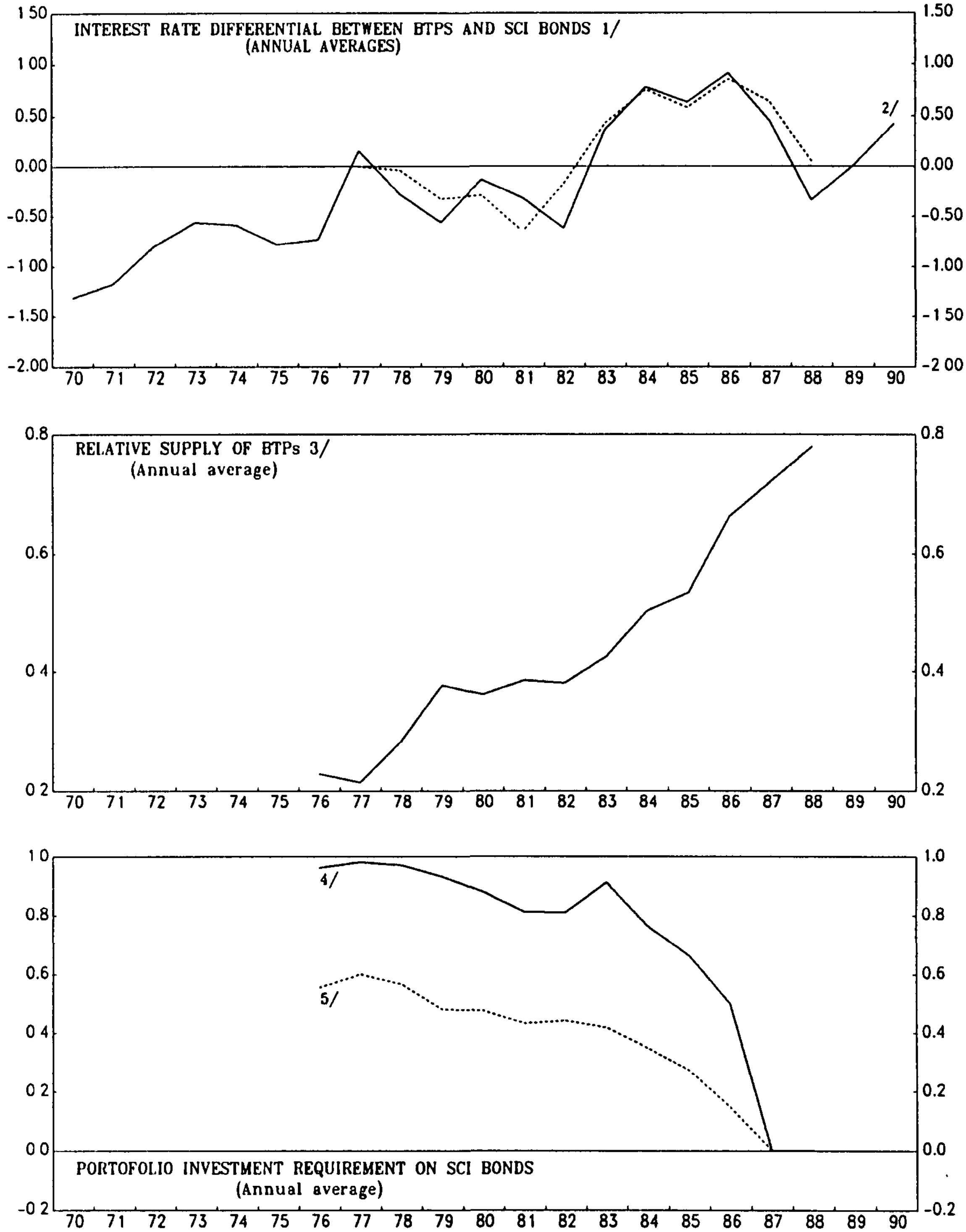
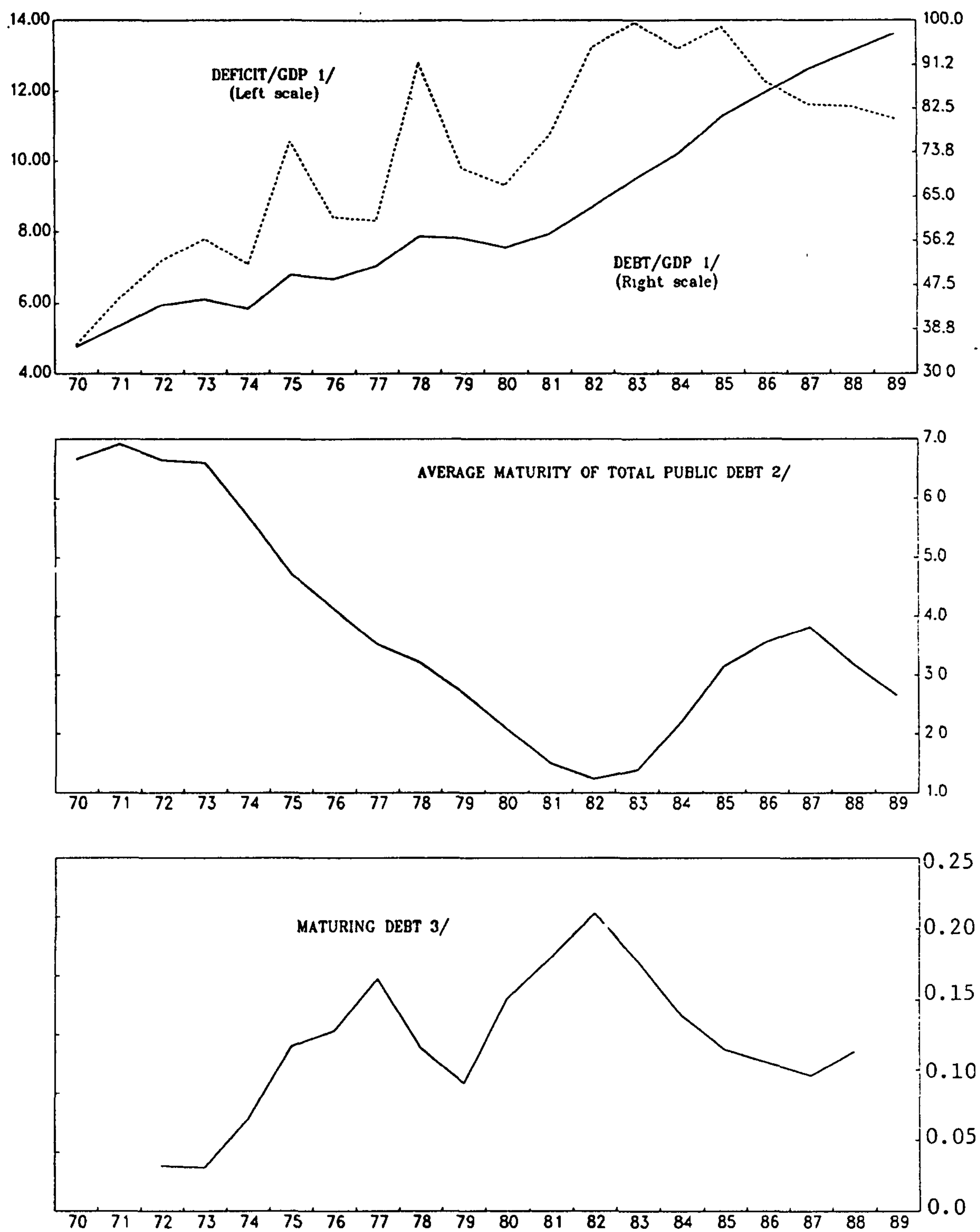
1/ The data used in Chart 1 and in the econometric estimates reported below refer to the bonds issued by SCI granting credit to finance industrial investment. The SCI financing real estate investment (whose bond issues at the end of 1988 represented around 18 percent of the outstanding stock of SCI and government fixed coupon bonds) are excluded because the market for these bonds appears to be highly segmented and in order to contain the size of the data base used for the panel data estimates presented below. The data refer to the Bank of Italy sample of net of tax yields to maturity published in the Bollettino Statistico. Alternatively, one period holding yields could have been used. While this alternative would have been more consistent with the mean variance analysis approach followed in Section II, it would have required taking into account explicitly the change in bond prices expected (in the absence of default) during the holding period (e.g., one year). To avoid the usual intricacies of measuring the unobservable, it was decided to use the yields to maturity, as, for example, in Barrett, Heuson, and Kolb (1986), Roley (1983), and Fisher (1959). As we are considering yield differentials computed under the hypothesis of no default (see equation (24)), the omission of expected capital gains matters only if the expected price change, under this hypothesis, differs for the two categories of bonds.

CHART 1
ITALY
INTEREST RATE DIFFERENTIALS, RELATIVE SUPPLY OF BTPs, AND PORTOFOLIO INVESTMENT REQUIREMENT ON SCI BONDS, 1970-90



Sources Bank of Italy, BOLLETINO STATISTICO.
1/ Yield to maturity differential between BTPs and Special Credit Institutions (SCI) bonds (industrial credit); the dashed line refers to the average differential computed on bonds of equal maturity
2/ The figure for 1990 refers to the first quarter.
3/ Ratio of the stock of BTPs to the stock of total non indexed bonds.
4/ Mininum investment requirement in SCI bonds over total SCI bonds.
5/ Mininum investment requirement in SCI bonds over total SCI bonds and BTPs held by market.

CHART 2
ITALY
RISK INDICATORS, 1970-89.



Sources: Bank of Italy, RELAZIONE ANNUALE, BOLLETTINO STATISTICO.
1/ The data are in percentage points and refer to the state sector.
2/ In years; excluding postal savings and overdraft facility at the central bank.
3/ Ratio beteen the amount of debt coming to maturity in each quarter and the stock of debt at the beginning of the quarter (annual averages).

investment requirement described in Section II(2) (Chart 1, bottom panel). After this temporary decline, 1/ the differential continued to rise, reaching 50 basis points in the first quarter of 1990.

2. Econometric estimates: data and methodological issues

The interpretation of movements in the yield differential depicted in Chart 1 is complicated by changes in the average maturity of BTPs and of SCI bonds during the period under consideration; as changes in the relative maturity of the two types of bonds may have heavily influenced the movements of the average yield differentials, the econometric analysis of Section III(3) will be based on yield differentials measured on bonds of the same residual maturity. 2/ The following procedure was used: the yield of individual SCI bonds and of BTPs was first collected on a quarterly basis from 1976 to 1988. 3/ A linear interpolation of the yields of SCI bonds was then computed for each quarter. This interpolation served two purposes: first, it provided an estimate of SCI bond yields also for maturities for which no SCI issue was outstanding; second, it helped to remove the high "noise" in individual SCI bond yields, most likely connected to market imperfections. 4/ Chart 3 plots, for a representative quarter, the yield of all SCI bonds against their maturity, together with their linear interpolation (solid line).

1/ As Chart 1 refers to annual averages the decline appears to occur mainly in 1988. On the contrary, the differential dropped in the last quarter of 1987 when the price of SCI bonds declined markedly pushing up their yields. This three-quarter lag between the removal of the constraint and the decline in the price of SCI bonds may be explained, not only by the usual adjustment lags in bank portfolios, but also by the conditions of Italian financial markets in 1987. When the investment requirement was removed interest rates were rapidly falling which made it convenient for banks to hold their fixed coupon bond portfolios, thus freezing their portfolio composition. The portfolio reshuffling was postponed to the last quarter of 1987 in the context of expectations of rising interest rates which followed the September credit crunch. In light of this behavior, the portfolio constraint indicator was introduced in the economic estimates of Section II(3) with a three-quarter lag; similar results were obtained with a two-quarter lag.

2/ It would have been more correct to compare bonds with the same financial duration rather than with the same maturity (see, for example, Carr, Halpern and McCallum(1974) and Barrett, Heuson and Kolb (1986)). This would have required detailed knowledge of the amortization plan (including interest payments) for each bond. Partial consideration of different amortization plans is obtained by using, instead of the residual maturity, the average residual life (which is published by the Bollettino Statistico of Banca d'Italia). For brevity, in the text, the term maturity will henceforth refer to the residual average life.

3/ The average yield in the last month of each quarter was used.

4/ Since the outstanding amount of each SCI bond issue is rather small, yields in specific months can be widely affected by random factors.

In the Chart, vertical dashed lines mark the yield differential for each maturity of the outstanding BTP issues. Thus, for each quarter, the number of available observations on the yield differential is equal to the number of outstanding BTP issues. Our sample covers 49 quarters (from 1976-IV to 1988-IV), with a total of 457 observations on the yield differential. The annual averages of these observations are plotted in Chart 1 (top panel, dashed line); also the data adjusted for maturity differences confirm the trend increase already observed on unadjusted differentials.

The individual observations on the differential, adjusted for expected inflation, 1/ were then regressed on the variables in the right of equation (26), and on the maturity of each yield differential. The first set of regressors is meant to explain the variability of the differential over time, while the individual maturity tries to explain the variability of the differentials for each given period. More formally, the estimated equation was the following:

$$\delta_{nt} = X_t'\Phi + \phi_9 f_{nt} + \eta_{nt} \tag{27}$$

where the subscript nt refers to the differential computed on the nth BTP (i.e. a certain BTP issue) observed at time t, X_t' is the matrix containing, together with a vector of ones for the constant, the eight time varying regressors included in equation (26), Φ is the vector of coefficients on these regressors, f_{nt} is the residual maturity (in months) of each BTP issue at time t and η_{it} is a stochastic error term. Note that ϕ_9 cannot be signed _a priori_ and will depend on the relative slope of the term structure of the two different types of bonds.

In light of the specific characterization of "individuals" (which in our panel correspond to certain BTP issues), it was decided not to introduce in the model individual effects, neither as nonstochastic components of equation (27) (as in a "dummy variable" model), nor as part of its error term (as in an "error components" model). Indeed, as BTPs present standardized features (in terms, for instance, of amortization plan) there is no plausible reason to assume _a priori_ that the differential yield computed on a specific BTP issue should systematically differ from the differential computed on other issues, except on account of different maturities, which are explicitly taken into consideration in (27).

1/ The adjustment is required because equation (26) refers to the differential between _real_ yields; real yields are here defined as $[(1+i)/(1+\pi)-1]$, where i and π are, respectively, the nominal yield to maturity and the expected inflation rate. Therefore the real yield differential differs from the nominal yield differential by a factor equal to $(1+\pi)$; as inflation in Italy reached 20 percent in the sample period the adjustment is not irrelevant for the result of the estimates. Expected inflation is here derived from the _Forum-Mondo Economico_ survey of expectation on inflation; see Visco (1987).

The estimation of equation (27) was based on a set of simplifying assumptions on the equation parameters and on the error term. All the parameters were assumed to be time invariant, although, according to the model expounded in Section II, they should depend on the tightness of the portfolio constraint, on the variance covariance matrix of yields and on the risk aversion parameter, all of which might change over time. 1/ As to the error term, we assumed that:

$$\mathrm{Cov}(\eta_{nt}, X_t) = \mathrm{Cov}(\eta_{nt}, f_{nt}) = 0 \tag{28}$$

$$E(\eta_{nt}) = 0 \qquad E(\eta_{nt}^2) = \sigma_{jt}^2 = \sigma_{it}^2 \qquad E(\eta_{it}, \eta_{jt}) = 0 \text{ for } i \neq j \tag{29}$$

$$E(\eta_{nt}^2) \neq E(\eta_{ns}^2) \text{ for } s \neq t \tag{30}$$

Equation (28) rules out nonzero correlations between the error term and the regressors. In this respect, the main reason for the inclusion of an error term in equation (27) is associated with the existence of a random disturbance in the demand for government bonds. Therefore, unless the relative supply of government bonds with respect to the total supply of bonds (i.e., q_g) is independent of demand conditions, q_g in equation (27) is likely to be correlated with η_{nt}. Indeed, q_g = G/(P+G) and, even if we assume that G (the supply of government bonds) is exogenous, P (the supply of SCI bonds) is likely to be affected by the level of interest rates. Moreover, when the portfolio model of Section II is included in a macroeconomic model of the economy, it is clear that the interest rate level, the interest rate differential, and P (and hence q_g) are determined simultaneously and that therefore q_g is likely to be correlated with η_{nt}. In what follows, however, we assume that at the quarterly level here considered the composition of supply is not affected by the level of interest rates and that, therefore, a random shock in the demand for government bonds is entirely reflected in changes in the interest rate differential. This assumption appears to be sustained by the long lags characterizing the supply response of SCI bonds to changes in the level of interest rates, due to lagged response of investments and of lengthy administrative procedure in the issue of SCI bonds.

Equations (29) and (30) summarize additional simplifying assumption on the error term. In particular, the variance of the error term is allowed to vary over time, but is assumed to be the same for all

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1/ The time invariance of the coefficients ϕ_0, ϕ_5, ϕ_6 and ϕ_8, also requires that the expected cost of default t_g is constant. The assumption of time invariance of parameters of the asset demand functions, from which (27) is derived, is fairly common in empirical research in the context of mean variance optimization; see, for example, Friedman (1985) and Roley (1983). An exception is Frankel (1983) in which demand parameters change according to the estimated changes of the variance covariance matrix.

observations in the same quarter; in addition, the covariance between disturbances related to different observations is assumed to be equal to zero. 1/

The heteroscedastic nature of the stochastic term in equation (27) required the adoption of a GLS estimation method, which under appropriate conditions provides consistent and asymptotically efficient estimates. With a view to improve efficiency in finite samples, an iterative estimation procedure was implemented. An initial estimate of the variance-covariance matrix was obtained from residuals of the OLS estimator applied to the pooled vectors of observations. This estimate was then used to obtain initial GLS estimates, producing residuals for a second estimate of the variance-covariance matrix. The procedure was iterated until convergence.

3. Econometric estimates: empirical results

Table 1 presents the GLS estimates of equation (27) obtained from panel data. Equation A refers to the most general specification here considered. With respect to equation (27), and its dynamic specification described at the end of Section II, there are, however, two differences. First, seasonal dummies have been included on account of the seasonality of some regressors and, possibly, of the yield differential. 2/ Second, two variables have been omitted: the lagged value of q_g and the second portfolio constraint variable $(P^*/P)(1-h)$. The exclusion of these variables is justified by the severe collinearity between these variables and the other regressors. The correlation coefficient between q_g and its lagged value is 0.997; the correlation coefficients between $(P^*/P)(1-h)$ and, respectively, h and $(P^*/B)(P^*/P^b)$ are 0.983 and

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1/ The assumption of zero correlation between disturbances within each quarter rules out the existence of a time specific effect; preliminary analysis including a time specific component did not yield satisfactory results (as expected given that almost all regressors vary over time but are constant over individuals). The hypothesis of a time correlation of residuals for the same individual (i.e., BTP issue) can also be ruled out for the same reasons for which individual effects were not included. It would be more attractive to test the hypothesis that the errors referring to specific maturities are correlated in time; however, in our panel the interest rate differentials observed in each period rarely refer to the same maturities and therefore it is not possible to test this hypothesis.

2/ A dummy on the differential computed on one BTP issue between the first quarter of 1981 and the second quarter of 1982 was also included. The coefficient on this dummy resulted to be very high (between 200 and 300 basis points in all specification) and could hardly be explained by market mechanisms; most likely, it was due to a measurement error, removed in the third quarter of 1982.

Table 1. Panel Data Estimates of the Equation of the Real Yield Differential
(GLS; 1976 IV-1988 IV; 457 Observations)

	Eq. A	Eq. B	Eq. C	Eq. D	Eq. E	Eq. F	Eq. G	Eq. H	Eq. I
	Equations								
Constant	-7.70	-7.68	-7.47	-7.05	-7.41	-7.81	-7.37	-4.69	-7.04
	(-23.61)	(-24.69)	(-26.50)	(-21.50)	(-22.12)	(-24.23)	(-17.91)	(-18.77)	(-22.27)
Bond distribution (h)	0.48	--	--	--	--	3.64	--	--	--
	(0.73)					(6.98)			
Relative supply (q_g)	7.21	7.24	7.79	4.91	5.16	5.35	--	6.42	4.92
	(13.33)	(15.42)	(25.83)	(10.84)	(11.25)	(11.57)		(20.81)	(11.08)
Portfolio constraint $(\frac{P^*}{B}\frac{P^*}{P^b})_{-3}$	0.068	0.067	0.066	0.063	0.069	0.076	0.053	0.054	0.063
	(21.26)	(26.91)	(26.76)	(23.49)	(23.56)	(26.06)	(20.77)	(20.52)	(24.17)
Debt maturity (m)	0.32	0.33	0.36	--	--	--	--	--	--
	(7.14)	(9.99)	(12.50)						
Maturing debt $(\frac{MA}{D})$	13.63	13.48	13.47	8.40	8.86	11.41	9.14	--	8.35
	(7.94)	(7.90)	(7.94)	(5.07)	(5.42)	(6.91)	(4.39)		(5.27)
Deficit ratio $(\frac{DF}{Y})$	0.20	--	--	--	-2.64	--	--	--	--
	(0.24)				(-3.44)				
Debt ratio $(\frac{D}{Y})$	0.13	0.21	--	0.89	1.03	0.28	2.01	--	0.89
	(0.84)	(1.60)		(8.51)	(8.52)	(2.20)	(19.89)		(8.85)
BTP maturity (f_n)	-0.0075	-0.0075	-0.0079	-0.0069	-0.0074	-0.0071	-0.0060	-0.0123	-0.0069
	(-7.69)	(-7.77)	(-8.56)	(-7.14)	(-7.57)	(-7.30)	(-4.80)	(-13.54)	(-7.17)
R^2	0.743	0.744	0.747	0.718	0.719	0.718	0.620	0.671	0.719
Standard error	0.28	0.28	0.28	0.27	0.28	0.28	0.32	0.25	0.27
Convergence indicator 1/	16.6	5.6	0.1	0.3	0.5	11.7	0.5	0.02	0.02

1/ Largest percentage change of coefficients between the nineth and tenth iteration (in absolute terms). All equations (except equation I) include three seasonal dummies and a dummy on one BTP issue between the first quarter of 1981 and the second quarter of 1982; equation I includes only a seasonal dummy (in the second quarter) and the 1981-82 dummy.

0.976. 1/ Despite these exclusions, the collinearity among regressors remains quite high and is probably the main reason behind the difficulty in the convergence of equation A. As shown in the last row of the table, after 10 iterations, at least one coefficient still changes by more than 16 percent. 2/ Despite this drawback, the results of specification A are broadly consistent with the theoretical model of Section II, with only one coefficient (that on the average maturity of the debt) having the "wrong" sign. 3/ In specification B, the two variables with lowest t statistics (i.e. the bond distribution and the deficit ratio) are omitted; the convergence indicator improves, but still convergence is not obtained after 10 iterations. Convergence is achieved in specifications C and D. In specification C the debt to GDP ratio is omitted without loss in terms of goodness of fit; however, the coefficient of the debt maturity still has positive sign. In specification D, instead, the debt to GDP ratio is reintroduced and the debt maturity is excluded, with a small decline in the adjusted R^2 and a slight improvement of the standard error. Clearly, there are no statistical grounds to select specification D over specification C; however, the signs of the coefficients of the former are consistent with the theoretical model and also the order of magnitude of the coefficients appears more plausible (see below). Specifications E and F differ from D in the reintroduction of the deficit ratio (E) and the bond distribution (F). The results are again unsatisfactory; in specification E the deficit ratio is now significant but has wrong sign, while specification F does not achieve

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1/ Preliminary estimate including the lagged value of q_g did not yield results very different from those of equation A. The signs of the coefficients on q_g and q_{g-1} was, however, opposite than what was expected (the first was negative and the second was positive); the sum of the two coefficients was very close to the coefficient shown in equation A.

2/ All the results in Table 1 refer to the estimate of the 10th GLS iteration. It was decided initially to accept convergence when all coefficients changed by less than one percent. Whenever this condition was not satisfied the number of iterations was increased to 30; it was, however, found that the equations not converging after 10 iteration would not converge either after 30 iterations.

3/ As mentioned above, the coefficient on f_{nt} (the maturity of the BTP on which the differential is computed) cannot be signed a priori; the fact that this coefficient is always negative in the estimates implies that the term of structure of interest rates, in the sample average, rises more steeply (or declines more gradually) for SCI bonds than for BTPs, a feature easily observable by simple inspection of the data. This feature may be connected to differences in the relative supply of BTPs and SCI bonds along the maturity axis. Indeed, the supply of BTPs has always been relatively larger on shorter maturities.

convergence. 1/ Starting again from D, specifications G and H were estimated to gather information on the relative importance of "relative supply" vis-a-vis "risk indicators", by removing them in turn. Both factors seem to be relevant, as both equations G and F are worse in terms of goodness of fit with respect to equation D. Note, however, that the deterioration is much stronger when supply factors are removed; indeed, the specification without risk factors, while having a lower R^2 than equation D, has the lowest standard error of all specifications and converges very rapidly. Finally, in equation I the statistically non-significant seasonal dummies included in equation D are removed without relevant changes in the results.

Given the characteristics of the panel used for the regressions, the usual diagnostic tests, particularly those on residual autocorrelation, cannot be applied to the regressions presented in Table 1. 2/ To circumvent this obstacle, and also as a check on the results obtained with disaggregated data, equation I was re-estimated on aggregate data obtained by averaging the cross-sectional observations for each time period. The first two columns of Table 2 show the OLS estimates and t-statistics, respectively, noncorrected and corrected for possible heteroscedasticity of unknown form. 3/ The comparison of the standard errors obtained in this way with those computed by the usual variance-covariance matrix estimator provides informal support to the hypothesis of heteroscedastic residuals. The equation was then re-estimated by GLS, weighting the observations with an estimate of the (time varying) variance of the disturbances computed from the residuals of the corresponding panel data estimates. The GLS estimates (Table 2, third column) are remarkably similar to those obtained from panel data, the main difference being the loss of significance on the coefficient on the BTP maturity; the high level of the adjusted R^2 and the inspection of actual and fitted values (Chart 4, top panel) confirm that the model

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1/ The lack of convergence of the equations in which the bond distribution is included, together with the low t-statistics of the corresponding coefficient in equation A, suggest that changes in the bond distribution were not a major determinant of the movements in the yield differential or, at least, that the available sample data does not allow to identify a specific effect of this variable.

2/ As already mentioned, there are n_t residuals for each period, but it is not clear what should be considered the lagged value of each residual: the residual on an interest rate differential of the same maturity in the previous period would be economically meaningful but is almost never observed, while the use of the residual on the same BTP issue observed in the previous period (i.e., on the residual on the BTP characterized by a specific serial number) could hardly be explained in economic terms.

3/ We used the variance-covariance estimator proposed by White (1980), which is consistent when the heteroscedasticity is of unknown form.

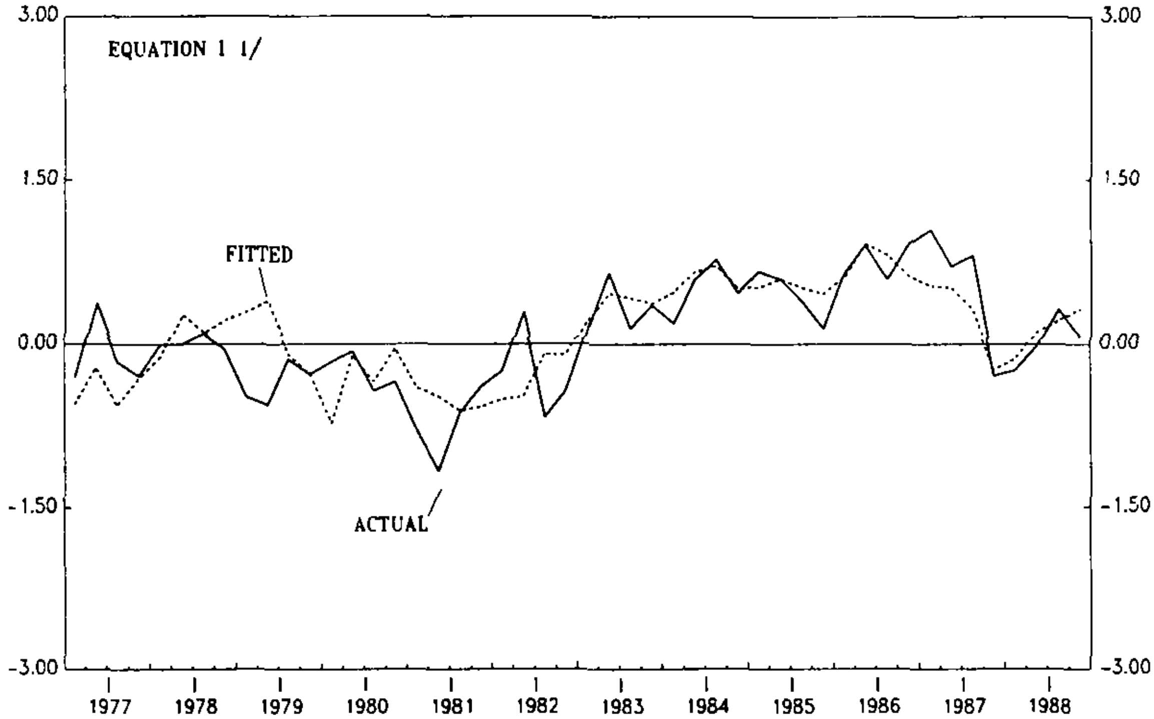
Table 2. Aggregate Data Estimates of the Equation of the Real Yield Differential (1976 IV-1988 IV)

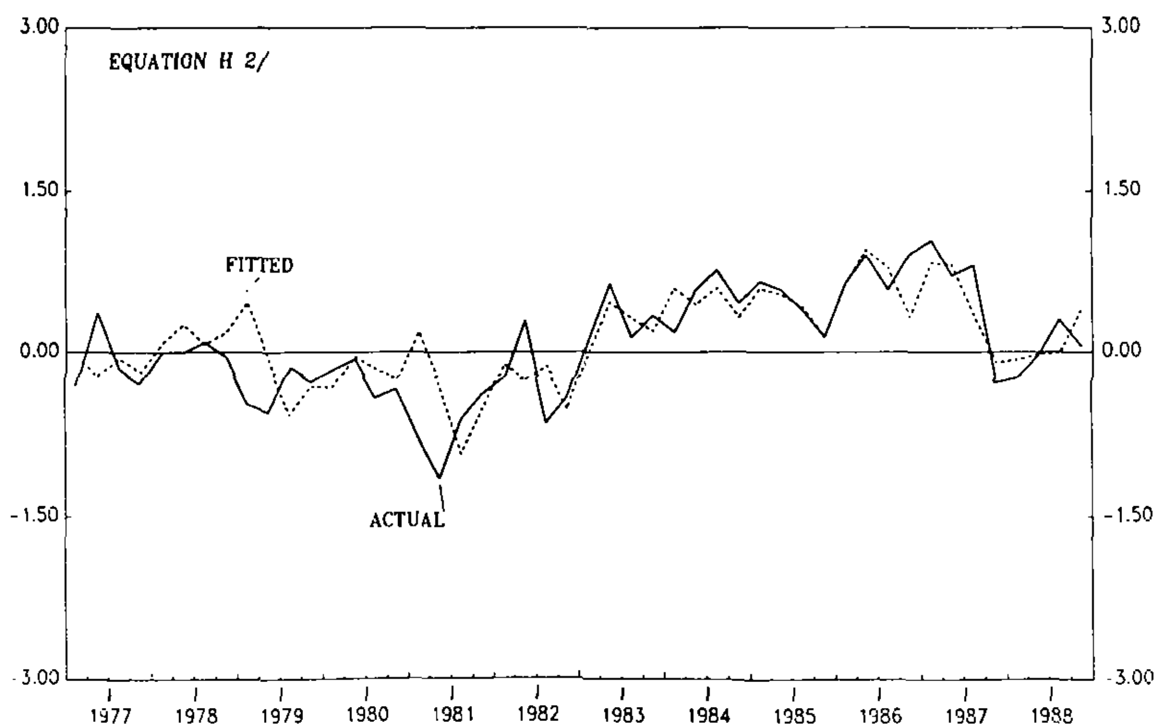
	Equation I				Equation H			
	OLS	OLS 1/	GLS	GLS 2/	OLS	OLS 1/	GLS	GLS 2/
Constant	-2.32	-2.32	-6.39	-6.34	-1.05	-1.05	-4.02	-5.72
	(-1.21)	(-1.15)	(-4.87)	(-3.66)	(-1.27)	(-1.16)	(-6.99)	(-6.31)
Relative supply (q_g)	2.50	2.50	4.28	8.59	2.76	2.76	5.70	7.27
	(1.26)	(1.60)	(3.57)	(6.87)	(2.90)	(2.74)	(8.70)	(7.31)
Portfolio constraint $(\frac{P^*}{B}\frac{P^*}{P^b})_{-3}$	0.025	0.025	0.056	0.070	0.019	0.019	0.047	0.063
	(2.38)	(2.33)	(8.59)	(6.43)	(2.13)	(2.08)	(7.89)	(6.73)
Maturing debt $(\frac{MA}{D})$	4.77	4.77	6.97	-0.46	--	--	--	--
	(0.96)	(1.12)	(2.54)	(-0.13)				
Debt ratio $(\frac{D}{Y})$	0.33	0.33	0.86	-0.02	--	--	--	--
	(0.46)	(0.51)	(1.85)	(-0.004)				
BTP maturity $(\frac{1}{n}\Sigma_n f_n)$	-0.026	-0.026	-0.0061	-0.0063	-0.033	-0.033	-0.016	-0.0038
	(-1.83)	(-1.79)	(0.61)	(-0.47)	(-4.23)	(-4.94)	(-4.61)	(-0.45)
$\bar{R}^2$	0.560	0.560	0.879	0.642	0.555	0.555	0.835	0.761
Standard error	0.33	0.33	0.14	0.16	0.33	0.33	0.13	0.13
DW	1.40	1.40	1.29	1.96	1.30	1.30	1.17	2.04
Residual autocorrelation coefficient	--	--	--	0.68	--	--	--	0.70
				(6.45)				(6.75)

1/ The t- statistics are computed on the basis of the White heteroscedasticity consistent covariance matrix estimator.

2/ Adjusted for first order residual correlation with the Cochrane-Orcutt technique.

CHART 4
ITALY
ESTIMATED EQUATIONS FOR REAL YIED DIFFERENTIAL
BETWEEN BTPs AND SCI BONDS, 1977(I)-1988(IV)





1/ GLS aggregate data estimates not adjusted for serial correlation.
2/ GLS aggregate data estimates adjusted for serial correlation.

is able to reproduce the main movements of the yield differential. 1/ The aggregate estimates reveal, however, the presence of some residuals autocorrelation: the DW statistics is on the lower margin of the inconclusive range of the test (which is 1.29-1.82 at the 5 percent level). The presence of serial autocorrelation implies a loss of efficiency in the estimates and may be revealing of some misspecification. 2/ As a first check, the equation was estimated again with the Cochrane-Orcutt technique (fourth column); the high value of the residual correlation coefficient and of its t-statistics confirm the presence of autocorrelation. After this correction, the coefficients on the risk factors collapse in both size and significance, while the opposite occurs to the coefficient of relative supply.

On account of these results, also equation H (which excludes the risk factors) was estimated on aggregate data. The new estimates (last four columns of Table 2) show that, as the DW remains low, the presence of the risk factors was not the reason for the serial correlation. The adjustment for serial correlation (last column of Table 2) reduces the value and the significance of the coefficient on the BTP maturity, but does not alter substantially the other coefficients, which remain close to those of the corresponding panel data estimates; actual and fitted values for this equation are plotted in Chart 4 (lower panel).

The presence of autocorrelation in the OLS and GLS aggregate estimates, and presumably also in the corresponding panel data estimates, 3/ can be due to several reasons. A first, quite obvious, reason is the static nature of the estimated regressions. The relevance of this factor was confirmed by computing the COMFAC test for equation H of Table 2. The value of the test statistics was 3.20 against a critical

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1/ In the aggregate estimates the maturity of BTP on which the interest rate differentials are computed is the average maturity computed in each quarter. In the averaging, most of the variability of this regressor is lost, which probably explain the loss of significance.

2/ In order to check for the possibility that the residuals autocorrelation could be a symptom of spurious regression among nonstationary variables, Phillips-Perron unit root tests were applied to the variables used in the GLS estimation procedures, following the testing strategy indicated in Perron (1988). For all the weighted time-series of observations on the relevant dependent and independent variables, the presence of a unit root was always decisively rejected. With all the caution required when applying to regression residuals tests procedures designed for univariate analysis, as expected the Phillips-Perron tests strongly rejected also the nonstationarity of residuals, confirming the indications provided by the very fast decay of their autocorrelation function.

3/ As recalled, a formal test of autocorrelation on the residuals of the panel data estimates is not possible; an informal indicator of the presence of autocorrelation in the panel data regressions comes from the computation of the DW statistics from the period averages of the residuals. This statistics was close to 1 for both equations I and H.

value of 9.49 of the χ^2 distribution at the 5 percent probability level. Thus the hypothesis that the correction for autocorrelation is "a convenient simplification" for a more complex dynamic process cannot be rejected.

An additional reason for autocorrelated residuals may be the imposition of time invariant parameters on a data-generating process characterized by coefficients varying over time. On account of the theoretical discussion of Section II, this may indeed be a concrete possibility in our case. The stability of the coefficients in the disaggregated equations I and H of Table 1 was, therefore, checked in an exercise of recursive GLS and by Chow tests. This was implemented by adding in sequence subsamples of observations corresponding to different time periods. As expected, the larger sample size increased the precision of the estimates, but mixed indications were obtained on the constancy of the parameters. While both point and interval estimates for the entire sample and their profiles during the recursions remained approximately within the initial confidence intervals in most cases, for some parameters the assumption of invariance over time appeared questionable. For example, in specification I (including risk indicators) the response parameter of the real interest rate differential to the debt to GDP ratio resulted statistically indistinguishable from zero in samples until approximately the end of 1983, while, with the addition of the most recent information the parameter increased in value and precision. Similarly, in specification H, the increasing sample size coincided with a gradual increase in value of the relative supply parameter. Formal Chow tests for the equality of parameters in the two subsamples 1976(IV)-1983(I) and 1983(II)-1988(IV), rejected the null of parameter invariance for both specifications. Similar indications on the parameter instability come from recursive application of the Chow test comparing the estimates for the whole sample with those obtained by the gradual addition of new information to an initial subsample. While indicative, these results should be considered with caution, since the distribution of the Chow test is known to be sensitive to the restrictive assumptions of nonstochastic regressors, and of normality and independence of the disturbances.

In order to gain additional insights, the recursive estimation procedure was also applied to equation H on aggregate data (Table 2), for which the corrections for heteroscedasticity and first-order serial correlation and the observed behavior of the residuals' autocorrelation function make more reasonable the application of Chow tests. In this case, the hypothesis of parameter constancy was always accepted at the 5 percent level both in the recursive applications of the Chow test and for the two separate subsamples. Even in this instance, however, the addition of the most recent information was accompanied by an increasing value and significance of the relative supply parameter.

Overall, the recursive estimates indicate that the estimates for the whole sample are remarkably close in value to those obtained on the observations for the most recent years, and that the explanatory power of the model is higher in the second part of the sample (approximately, after 1983) than in the first, possibly because of a greater sample variability in the regressors in the second subsample.

To allow, at least partially, for time dependence of the parameters, equation I was re-estimated on aggregate data by nonlinear least squares entering the debt ratio in a nonlinear fashion. Indeed, the perception of risk may be connected nonlinearly to public imbalance indicators: increases in the debt to GDP ratio may be considered irrelevant when the ratio is low, but may attract much attention when the ratio is already high and/or is rising rapidly. In order to explore this possibility, the response parameter of the debt to GDP ratio was allowed to vary according to a logistic function of the level of the ratio itself. The logistic form seemed attractive a priori since it allows a parsimonious parametrization (only 2 additional parameters are needed), and since it broadly replicates the behavior of the coefficient on the debt ratio observed in the recursive estimates. However, the results did not improve upon those presented in Table 2. The significance of relative supply and of the portfolio constraint was confirmed, but the estimates for the debt ratio parameter and for the parameters of the logistic curve were statistically insignificant and nonrobust to selected starting values in sensitivity analysis. Convergent estimates could not be achieved in some cases, depending again on starting values, and overall goodness of fit measures deteriorated. While informative, this attempt to model nonlinearities is by no means conclusive: more attention will have to be dedicated in future research to a better selection of the functional form and to alternative estimation methods involving switching regimes.

In conclusion, the available empirical evidence seems to confirm the relevance of supply effects, risk indicators and institutional constraints in explaining the movements in the yield differential between government and SCI bonds, while relevance of the bond distribution between banks and nonbank public is not confirmed. The evidence also suggests that supply factors were more important than risk indicators; indeed, the simple specification H of Tables 1 and 2, which excludes risk indicators, seems to describe adequately the behavior of the yield differential and passes the statistical diagnostic tests.

As to the multipliers implicit in the point estimates so far discussed, Table 3 summarizes the effect of changes in supply and risk factors on the yield differential. More specifically, the Table shows the effect of a change in the public debt of Lit 10 trillion at the end of 1988 (around 1 percent of total debt and also of GDP), financed to one third by issuing BTPs. 1/ While the different specifications differ in the split of the total effect between risk and supply factors the overall effect appears to be close to 20 basis points in all

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1/ For simplicity, it is assumed that the average maturity of the debt does not change.

Table 3. Effect on the Yield Differential of a One Percent Increase in the Public Debt 1/

	Equation I		Equation H	
	Panel Data	Aggregate data 2/	Panel Data	Aggregate Data 3/
Total	19	16	20	22
Of which:				
Supply factors	15	13	20	22
Risk factors	4	3	--	--

1/ In basis points; the effect is computed relatively to the value of the variables at the end of 1988. At the same date, a change in total debt by 1 percent corresponded to around Lit 10 trillion (1 percent of GDP). It is assumed that one third of this increase is financed by BTP issues.

2/ GLS estimates not adjusted for serial correlation.

3/ GLS estimates adjusted for serial correlation.

specifications. Note also that the specifications in which risk factors are present indicate that 80 percent of the overall effect is due to a change in the relative supply of assets. In order to allow a better appreciation of the relative importance of these effects in explaining the movements in the yield differential, Table 4 presents the decomposition of the change in the yield differential between the beginning and the end of the sample period. Again, all specifications agree that the overall effect of supply and risk factors was close to 400 basis points; when risk factors are present, they are estimated to account for one third of the overall effect, mainly as a consequence in the increase in the debt ratio. The effect on the yield differential of supply and risk factors was to a large extent offset by the progressive removal of the portfolio constraint; this removal allowed a decline in the yield differential of over 300 basis points.

One important caveat has to be recalled at this point. The estimate of the effect of changes in the relative bond supply may appear quite large and it would imply low substitutability between BTPs and SCI bonds of the same maturity. However, as mentioned in Section II, the estimates presented in this paper reflect the dominance in the sample period of the portfolio constraint which largely reduced the substitutability between the two types of bonds; as a consequence the estimates here presented tend to overestimate the effect on the yield differential of changes in relative supply (and indeed also of risk factors) in the absence of a portfolio constraint.

IV. Conclusions

This paper presented econometric evidence on determinants of the movements of the real yield differential between government and non-government paper in Italy between the middle of the 1970s and the end of the 1980s. We showed that the increase in the differential that occurred through 1986 was heavily influenced by the deterioration of public finances. This deterioration affected the differential in two ways: first, through an increase in the relative supply of government bonds with respect to SCI bonds, in the context of imperfect substitutability between the two assets; second, through an increase in the default risk premium, reflected by changes in selected default risk indicators (specifically, the debt to GDP ratio and the share of maturing debt over total debt). According to our analysis, the decline in the differential at the end of 1987 was mainly due to the removal of the investment requirement forcing banks to hold a large share of their bond portfolio in SCI bonds. The factors explaining the increase in the differential through 1986 remained, however, at work, setting the differential again on a rising trend in 1989 and 1990. Thus, the question asked in the introduction (i.e., whether the available data on financial market yields reveal the existence of an increasing default risk premium on Government debt) can be answered in the affirmative, but with qualifications. Indeed, we found that the factors here interpreted as risk indicators accounted to some extent for the observed movements of the

Table 4. Decomposition of the Change in the Yield Differential (1976 IV-1988 IV) 1/

	Equation I		Equation H	
	Panel data 2/	Aggregate data	Panel data 2/	Aggregate data
Relative supply	297	258	386	438
Portfolio constraint	-343	-305	-294	-343
Maturing debt	-25	-21	--	--
Debt ratio	164	161	--	--
BTP maturity	-7	-6	-12	-4
Residual	-71	-72	-61	-76
Total change	15	15	15	15
Of which:				
Supply and risk factors	436	398	386	438

1/ In basis points.
2/ The total change, the effect of the BTP maturity and the residual refer to the average of the observations.

yield differential. However, we also found that relative supply factors were statistically more robust and quantitatively more important than risk indicators in explaining the trend increase in the differential.

These conclusions are of relevance for fiscal policy and for debt management. The evidence that the yield differential between government and nongovernment paper rose as a result of increasing fiscal imbalances, a result which contrasts with the Ricardian equivalence hypothesis, implies that a policy of fiscal adjustment should potentially benefit from a reduction not only in the general level of interest rates, due to the standard macroeconomic effect of restrictive fiscal policies, but also from a reduction in the additional yield paid at present by the government on its debt. The predominance of relative supply factors, and in general of stock over flow indicators of public finance imbalances, implies also that, in order to reap this reward, the fiscal effort must be sustained over time, as to allow for the proper adjustment in the stocks of financial assets. Moreover, the evidence that part of the increase in the cost of public debt is due to the imperfect substitutability between government and nongovernment assets in the portfolio of Italian investors supports the suggestion recently advanced (see Ministero del Tesoro (1989)) that the interest burden could be reduced by increasing the share of debt sold to nonresidents, whose "appetite" for Government paper may not yet be entirely satisfied. Finally, measures to increase the efficiency of the secondary market for treasury paper in Italy, hence raising its marketability and liquidity, may also be instrumental in bringing about a decline in the cost of debt, as they would tend to reduce the component of the risk differential unrelated to default risk.

While the conclusions presented in this paper appear fairly robust with respect to changes in the specification of the estimated equation, use of aggregate data versus panel data and different estimation techniques, some caveats are nonetheless required. The analysis allowed only partially for time dependence of parameters, which is instead possible in light of the theoretical analysis of Section II. Moreover, during the time interval considered in this paper, market behavior was distorted by the investment requirement on bank bond portfolios; as mentioned, we partially considered in our estimates the effect of this requirement; yet, as the constraint reduced the elasticity of portfolio shares to changes in the interest rate differentials in the sample period, the parameters reflecting the effect of supply and risk factors on the differential are probably overestimated relative to their values in the absence of constraints.

Asset Demand Curves in a Mean-Variance Framework

The maximization problem of section 2.1 in the text is solved as follows: the Lagrangean is given by:

$$\text{(A.1)} \quad L = U(W_{-1}Q'R^e,\ W_{-1}^2 Q'\Omega Q) + \lambda(Q'J - 1)$$

and the first order conditions are:

$$W_{-1}U_1R^e + 2\ W_{-1}^2 U_2\Omega\ Q + \lambda J = 0$$

and $Q'J = 1$

or:

$$\text{(A.2)} \quad \begin{bmatrix} 2W_{-1}^2U_2\Omega & J \\ J' & 0 \end{bmatrix} \begin{bmatrix} Q \\ \lambda \end{bmatrix} = \begin{bmatrix} -W_{-1}U_1R^e \\ 1 \end{bmatrix}$$

The inversion rule for partitioned matrices provides the solution for Q given by equation (4) in the text.

Market Equilibrium in the Presence of a Portfolio Constraint

The demand for total SCI bonds is given by:

$$P^h + P^b = P = (\beta_0^h + \beta_1^h \delta^e) B^h + \gamma \frac{(P*)^2}{P^b} + (\beta_0^b + \beta_1^b \delta^e)\ (1 - \gamma \frac{P*}{P^b}) B^b \qquad (A.3)$$

and, after dividing for B and rearranging the terms:

$$1 - q_g = \beta_0^h h + \beta_0^b (1-h) + \beta_1^h h \delta^e + \beta_1^b (1-h) \delta^e + \gamma \frac{P*}{B} \frac{P*}{P^b} - \gamma\ \beta_0^b \frac{P*}{P^b}(1-h) - \gamma \beta_1^b \delta^e \frac{P*}{P^b}(1-h)$$

$$(A.4)$$

If, as done for equation (13) in the text, we assume $\beta_1^h = \beta_1^b = \beta_1$, equation (21) becomes:

$$1 - q_g = \beta_0^b + (\beta_0^h - \beta_0^b) h + \gamma \frac{P*}{B} \frac{P*}{P^b} - \gamma \beta_0^b \frac{P*}{P}(1-h) + \beta_1 [1 - \gamma \frac{P*}{P^b}(1-h)] \delta^e \qquad (A.5)$$

Recalling that $\beta_0^b = 1 - \alpha_0^b$, $\beta_0^h = 1 - \alpha_0^h$ and $\beta_1 = - \alpha_1$ (where all the α parameters refer to the coefficients of the demand for government bonds; see Section 2.1) equation (A.5) becomes:

$$q_g = \alpha_0^b + (\alpha_0^h - \alpha_0^b) h - \gamma \frac{P*}{B} \frac{P*}{P^b} + \gamma (1 - \alpha_0^b) \frac{P*}{P}(1-h) + \alpha_1 [1 - \gamma \frac{P*}{P^b}\ (1-h)] \delta^e \qquad (A.6)$$

Equation (20) in the text can be obtained by solving (A.6) for δ^e.

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