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STOCK RETURNS, THE INTEREST RATE AND INFLATION IN THE ITALIAN STOCK MARKET: A LONG-RUN PERSPECTIVE

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This paper investigates the behaviour of stock prices in Italy over the 1963-1995 period. By means of a time-series analysis of both the long- and the short-run properties of stock prices and other macroeconomic variables, we find strong evidence of a long-run equilibrium negative relation between the inflation rate and a real stock price index. The dynamic adjustment of stock prices towards the equilibrium relation is also analysed.

JEL Classification: *C32, E44, G10*
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INTRODUCTION

This paper explores the relations among stock market returns, the nominal short-term interest rate and inflation in Italy over the last thirty-five years. The emphasis on the rate of inflation as an important factor in explaining bond and stock market behaviour is largely due to the key role of inflation in the determination of economic activity. Indeed, according to some theoretical models, the connection between financial markets and inflation simply reflects deeper links between inflation and macroeconomic variables affecting equilibrium in the financial and goods markets.

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Informal evidence confirms the relevance of inflation as an indicator of the state of macroeconomy. Rising inflation is usually a bad piece of news for financial markets. Recently, financial markets have reacted very strongly to any direct or indirect news about inflation; for example, in the US, downward revisions of the inflation rate boost the stock market. Even historically, financial markets seem to be influenced by the prevailing rate of inflation, to the extent that the depressed US stock market of the 1970s is suspected to be related to the upsurge in inflation in the same decade. Several applied studies provide empirical evidence extending this negative relation also to the bond market. Starting with Fisher, most researchers have found evidence of an insufficient reaction of nominal interest rates to inflation, yielding volatile real interest rates. To the extent that inflation affects both the interest rate and stock returns, there may be also a connection between inflation and the equity premium, defined as the difference between stock returns and the riskless interest rate (Blanchard, 1993, has recently considered this possibility in an applied study for the US).

The relation between the stock market and inflation can be justified on the basis of a number of theoretical models, briefly reviewed below. In general, these models suggest the existence of a negative relation both between inflation and stock returns and between expected inflation and stock returns. As discussed below, there are various reasons to explain this negative comovement: tax effects, risk, macroeconomic links.

However, the focus of this paper is empirical. In order to explain our methodology, it may be useful to start with a preliminary look at some relevant data. Fig. 1 shows the real value of an investment in stocks and in short-term, risk-free assets (three-month Treasury bills) over the 1962-1995 period. Fig. 2 compares the rate of inflation with the level of the real stock market index over the same timespan. Several considerations are suggested by the data.

First, in the period 1962-1995 investing in short-run assets provided a return which is superior to that yielded by stocks. This may be viewed as an Italian anomaly, contrary to the international evidence on the equity premium puzzle (Campbell, 1996; Goetzmann - Jorion, 1997, analysing a longer period, 1928-1995, confirm the anomaly of the Italian market). Second, there seems to be a strong negative correlation between inflation and the *level* of the index as opposed to its rate of change. This peculiarity needs to be considered with care in the empirical analysis in the light of the time-series properties of the variables.

In order to characterise empirically the long- and short-run comovements between stock prices and inflation, we consider an atheoretical representation of the variables in a multivariate time-series framework. Using a

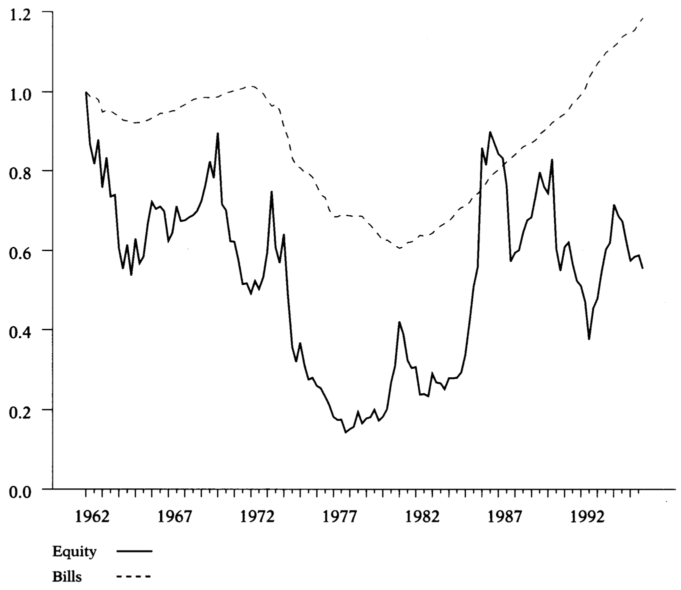

Fig. 1 - The Real Stock Index and the Real Short-Term Rate Index (1962(1) = 1)

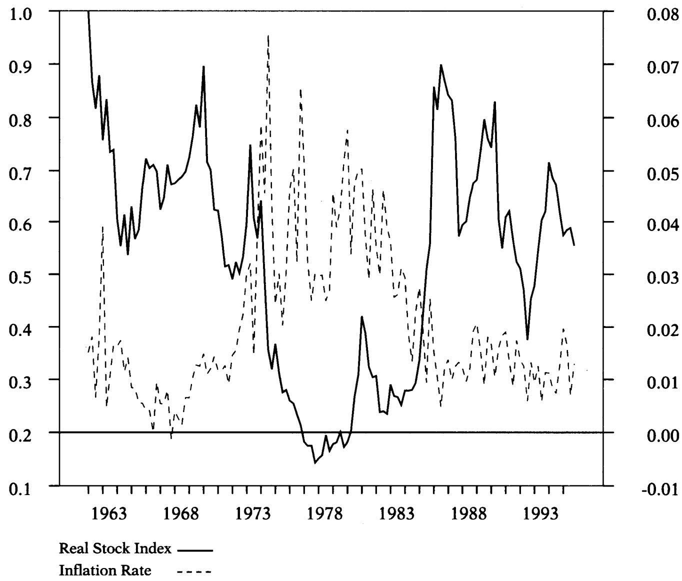

Fig. 2 - The Real Stock Index (Left Scale, 1962(1) = 1) and the Inflation Rate (Right Scale, Quarterly Rate): 1962-1995

cointegration methodology we assess the existence of long-run relations among the relevant variables and discuss a short-run model consistent with the detected long-run relations.

In our opinion, application of econometric methodologies to Italian financial data is useful given the paucity of applied research on the Italian market. It is true that such a market is rather small if compared with that of other developed countries. However, it is also true that in general Italian financial intermediaries do not ignore it. Some studies show that many Italian mutual funds invest a significant amount of resources in the Italian stock exchange, and this is in line with the available empirical evidence on insufficient international diversification by institutional investors (the so-called diversification puzzle, see French - Poterba, 1991). Also, the introduction of Italian pension funds will create institutions with a particular interest in hedging long-run inflation risks that seem particularly relevant to households saving for retirement. In this light it is essential to understand the hedging properties of various financial assets.

The paper is organized as follows. Section 1 briefly reviews theories and empirical results. Section 2 presents the empirical analysis. Section 3 discusses the consequences for the application of asset pricing theories to the Italian case. There follow conclusions.

1. THEORIES AND FACTS

A large literature has studied both from an empirical (mainly for the US case) and a theoretical point of view the relations between stock returns and inflation and between inflation and the nominal interest rate. Various explanations have been advanced in the literature, based on complete macroeconomic models, on tax effects and on particular assumptions on investors' behaviour.

A well-known macroeconomic model is due to Fama (1981), who explains such a correlation on the basis of a structural model involving a positive relation between current real stock returns and expected output growth and a negative relation between expected output growth and current inflation. A macroeconomic model also seems to be implicit in the 1996-1997 dynamics of the US stock market, where any news of high demand signals a possible tightening of monetary policy as a preventive move against inflation. According to this interpretation, inflation is important also as a signal of contractionary monetary policy.

Feldstein (1980) proposes an explanation based on tax considerations. He distinguishes between a constant expected rate of inflation, determining a steady increase in stock prices, and an unexpected jump in the rate of inflation causing a discrete fall in stock prices. Profits and dividends may be negatively affected by inflation because of the use of historic-cost methods of depreciation. This causes an increase in taxable profits so that real after-tax profits fall with an increase in inflation.

A behavioural explanation is proposed by Modigliani - Cohn (1979), according to whom agents may mistakenly discount real dividends paid by firms with the nominal, rather than the real, rate of interest. It follows that an increase in inflation, usually resulting in a surge in nominal interest rates, will decrease stock prices also in an economy where real activity is independent of inflation.

The Lucas (1978) model, as used for example by Mehra - Prescott (1985), represents the standard general equilibrium model for the analysis of asset prices. The monetary version of the model developed by Labadie (1989) brings out the important role of inflation. The first-order conditions of the representative agent maximization problem show that the expected real stock return depends on the covariance between the return and the marginal rate of substitution between current and future consumption. Labadie also shows that «stochastic inflation affects the risk characteristics of the equity's return through the assessment of the entire path of the inflation tax and the correlation of the marginal rate of substitution with the rate of appreciation in the purchasing power of money» (p. 282)[1].

The relations between stock returns, the interest rate and inflation have been carefully investigated in the applied literature too. One striking empirical regularity is that over the last century the average real rate of return on equity in the US market has been about 7% while the average rate of return on short-term assets has been about 1%, yielding a 6% equity premium. Since the work of Mehra - Prescott (1985) the magnitude of the historical equity premium has been considered a puzzle because of the impossibility to generate it from a theoretical model. After calibration to "reasonable" parameters for the description of technologies and preferences, standard models could generate an equity premium of about 1% only, far from the observed level. Similarly, Grossman - Shiller (1981) estimate the coefficient of relative risk aversion of the representative agent from the Euler condition for the choice of stocks. The equity premium puzzle manifests itself in the finding that only a very large coefficient of relative risk aversion is able to reconcile

[1] Other theoretical models have been developed to study in detail the effects of inflation on equity returns, see for example Stultz (1986) and Day (1984).

the large variability of stock returns with the low variability of the rate of growth of consumption[2].

Before concluding this survey section it is useful to mention the atheoretical work of Blanchard (1993) who notes that a large differential between US stock and bond returns in the 1970s is associated with a sharp increase in inflation while the small differential in the 1980s is associated with a sharp decline in inflation. We shall come back to these empirical results while discussing our own results for the Italian case.

2. Empirical analysis

In this section an empirical analysis of the Italian data is carried out, with particular emphasis on the determination of the long-run and short-run relations among stock prices, the interest rate and inflation.

We take a different empirical approach with respect to other recent investigations of the stock prices behaviour in the "equity premium puzzle" tradition (see in particular the study by Panetta - Violi, 1997). Here we are not directly interested in assessing whether the magnitude of the historical equity premium on the Italian stock market may be explained by an equilibrium model of stock price formation. Therefore, we are not going to either estimate the first-order conditions derived from the representative agent problem or test the time-series restrictions implied by an intertemporal model. Instead, we shall provide a data-respectful econometric analysis of the Italian evidence, applying the tools of time-series analysis to a cointegrated system composed of an index of the stock market and other macroeconomic variables to study both the long- and the short-run relations among them. Such relations are estimated in the context of a small-scale "structural" Vector Autoregressive (VAR) model[3].

[2] There has been much applied and theoretical work since the original paper by Mehra and Prescott, recently surveyed by Kocherlakota (1996). Attempts at solving the equity premium puzzle have been based on transaction costs (Heaton - Lucas, 1995), non-separable preferences (Campbell - Cochrane, 1995) and endogenous uncertainty (Kurz - Beltratti, 1996).

[3] This methodology, building on the traditional VAR techniques pioneered by Sims (1980), has been widely applied in macroeconomics by, among others, Blanchard - Watson (1986), Bernanke (1986), Blanchard (1989) and Gali (1992). Thorough treatments of this and related techniques are provided by Giannini (1992) and Hamilton (1994).

2.1. *The Data*

The data we use in the empirical analysis are the following:

s : the logarithm of the real stock price index, obtained by deflating the *BCI* index adjusted for dividends (as in BCI, 1996) with the consumer price index;

i : the nominal short-term risk-free rate (the higher between the bank deposit rate and the end-of-quarter interest rate on three-month Treasury bills);

π : the log-difference of the consumer price index as a measure of the quarterly rate of inflation;

y : the logarithm of the index of industrial production (seasonally adjusted);

m : the logarithm of the real stock of money, obtained by deflating nominal end-of-quarter M2 balances (seasonally adjusted) with the consumer price index.

The data are quarterly, from March 1963 to December 1995. There are several reasons for our choice of the quarterly frequency. First, our analysis is aimed at studying the mechanism of dynamic adjustment of the variables in the face of business cycle fluctuations, for which quarterly series seem appropriate. Second, we avoid monthly data because of the large noise component which may obscure the relationship among macroeconomic and financial variables.

Several additional data issues are worth mentioning before analysing the time-series properties of our series. Firstly, we use an index which certainly suffers from the "survival" problem (see Brown - Goetzmann - Ross, 1995). The stock index collects the value of those companies which did not disappear over the years and can therefore give a false impression about the rate of return which would have been obtained by investing in the stock market. However, this does not seem to be a major problem *per se* as our main focus is not the equity premium but the relation between stock prices and inflation, even though in principle it is possible that in inflationary periods the companies which go bankrupt are those that are particularly sensitive to inflation.

A more relevant problem might be the use of a general stock index together with a measure of industrial production. In principle one would like to compare the behaviour of industrial production with a stock index of the industrial sector. Our choice here was dictated by data availability. There is no currently available stock index of the industrial sector which includes dividend payments. Moreover, revised data for *GDP* are not available for the 1960s (nevertheless, in order to evaluate the robustness of our findings, we also used the available *GDP* series instead of the industrial production index, obtaining qualitatively similar results).

The use of a general index may affect the conclusions regarding the hedging ability of the stock market in the face of inflationary shocks. An index focused on industrial shares may well show superior hedging abilities due to the intrinsic nature of firms as real assets, while the Italian general index contains a large proportion of firms operating in the financial and banking sectors which may be sensitive in various ways to inflation. All the results that follow should therefore be interpreted in the light of the above observations.

Finally, we use a time series which goes back to the 1960s only. It is possible that extending the analysis to a longer timespan (for example, using the series in Panetta - Violi, 1997, going back to the 1880s) might produce different results. However, the use of a time series including two world wars and major structural transformations of financial markets (besides the difficulties due to the statistical treatment of structural breaks in the data) may be less informative than more recent data for the characterization of current economic conditions.

2.2. *The Time-Series Properties of the Data*

As a preliminary step towards the correct formulation of a cointegrated VAR system, we investigate the time-series properties of the data used in the analysis, starting from Phillips - Perron (1988) unit-root tests on the levels and first differences of s, i, π, y and m. Table 1A reports the values of the $z(t_{\alpha^*})$ Phillips-Perron statistics for the variables listed above, showing clear evidence of a unit root in all series. In addition, the hypothesis of a deterministic (linear) trend in the industrial production index and the real money stock has been explicitly tested and rejected. When the tests are applied to the real interest rate $(i - \pi)$ the results are not overwhelmingly in favour of the presence of a unit root (the value of the $z(t_{\alpha^*})$ test ranges from -2.60 to -2.95, depending on the number of lags used); we will perform a more informative test on the long-run relation between inflation and the nominal interest rate below in the context of the multivariate system.

The *I(1)* nature of our basic series justifies their inclusion as endogenous variables in a VAR system in order to investigate the existence of long-run comovements in stock prices, the interest rate, inflation, an index of economic activity and a monetary aggregate. The vector of the endogenous va-

TABLE 1

A. *Univariate Phillips-Perron Unit-Root Tests*

VARIABLE	$z(t_{\alpha^*})$ STAT.	VARIABLE	$z(t_{\alpha^*})$ STAT.
s	–1.88	Δs	–10.69**
i	–1.63	Δi	–9.20**
π	–1.69	$\Delta \pi$	–7.43**
y	–2.08	Δy	–12.72**
m	–2.57	Δm	–9.45**

** denote significance at the 1 % level.

B. *Cointegration Analysis*

EIGENVALUE:	0.205	0.173	0.115	0.089	0.001
HYPOTHESIS:	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	$r \leq 4$
λ_{MAX}	29.2	24.1	15.6	11.8	0.2
90% CRIT. VALUE	20.9	17.2	13.4	10.6	2.7
λ_{TRACE}	80.8	51.7	27.5	11.9	0.2
90% CRIT. VALUE	64.7	43.8	26.7	13.3	2.7

r denotes the number of valid cointegrating vectors.

Test of weak-exogeneity of m for the long-run parameters:

$\chi^2(3) = 5.06$ p-value: 0.17

riables is then specified as $\mathbf{x}_t = (s_t, i_t, \pi_t, y_t, m_t)'$. The cointegration analysis is initially performed on this five-variable system by means of the Johansen (1988, 1991) maximum-likelihood procedure, with five lags of the endogenous variables and deterministic centered seasonals. In error-correction form, the estimated system is expressed as:

$$\Delta \mathbf{x}_t = \sum_{i=1}^{4} \mathbf{A}_i \Delta \mathbf{x}_{t-i} + \mathbf{\Pi} \mathbf{x}_{t-1} + \mathbf{K} \mathbf{d}_t + \mathbf{e}_t \tag{1}$$

where the matrix $\mathbf{\Pi}$ contains all relevant information on the long-run relations linking the variables in $\mathbf{x}$, and $\mathbf{d}$ is the vector of deterministic terms. As shown by Johansen (1988) and Johansen - Juselius (1990) the existence of long-run cointegrating relations among the elements of $\mathbf{x}$ imposes crossequation restrictions on the coefficients of the $\mathbf{\Pi}$ matrix, reducing its rank r. In particular, if $0 < r < n$ (with n being the dimension of the system in (1)), then there exist r linear combinations of the n $I(1)$ variables and $\mathbf{\Pi}$ may be expressed as the product of two $n \times r$ matrices: $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$. The columns of $\boldsymbol{\beta}$ contain the coefficients of the r cointegrating vectors, forming the stationary combinations $\boldsymbol{\beta}'\mathbf{x}_{t-1}$, whereas the elements of $\boldsymbol{\alpha}$ are the weights of each cointegrating relation in the system equations. The Johansen procedure provides a test for the number of valid cointegrating vectors in the system together with estimates of the coefficients of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ under the reduced rank assumption on $\mathbf{\Pi}$.

Correct implementation of the estimation procedure requires the system in (1) to be a well-specified representation of the underlying data generating process (below we report results from diagnostic and stability tests). As a first check, focusing on the more recent part of the sample period (1993-1995), Fig. 3 displays the forecasting performance of the VAR for the levels of each variable, along with a Chow forecast test for the system as a whole. No sign of overall predicting failure is shown, though some forecasting problems may be detected for individual variables (the interest rate and industrial production in 1993 and money in 1994).

From the results of the Johansen cointegration tests reported in Table 1B we argue that there is evidence of three valid long-run relations: the trace statistics (λ_{TRACE}) does not reject the null hypothesis of $r \leq 3$ at the 10% significance level, whereas the maximal eigenvalue statistics (λ_{MAX}) yields only marginal rejection. The estimated vectors, plotted in Fig. 4, display the required stationary behaviour, though with some persistent deviations from equilibrium. In order to assess the stability over time of the estimated long-run coefficients, the cointegration analysis has been performed recursively over the 1976-1995 period (observations up to 1995 being necessary for initialization). As the second column of Fig. 4 shows, the recursive estimates of the system three largest eigenvalues (associated with the valid cointegrating vectors) are sufficiently stable throughout the sample to suggest stability of the long-run coefficient estimates[4].

The estimates of system (1) with the imposition of a reduced rank of three on the long-run matrix $\mathbf{\Pi}$ yields very small coefficients for the elements of $\mathbf{\Pi}$ in the equation for the real money stock. Formal testing of the hypothesis that none of the three cointegrating vectors enters significantly

[4] In implementing the recursive procedure the estimated coefficients in the $\mathbf{A}_i$ and $\mathbf{K}$ matrices, capturing the system short-run dynamics, are kept fixed at their full-sample values (therefore adopting the R-representation of the recursion in the terminology of HANSEN - JOHANSEN, 1992).

the equation for Δm (*i.e.*, the last row of $\boldsymbol{\alpha}$ has all elements equal to zero) is carried out by means of a LR test following Johansen - Juselius (1992): the test result indicates that the null hypothesis cannot be rejected at standard levels of significance. Therefore the money stock can be considered weakly exogenous with respect to the parameters in the $\boldsymbol{\beta}$ matrix and there is no loss of information in conditioning the system on Δm_t, reducing its dimension to four equations. On the basis of this result, the final formulation of the system that we will use in the remainder of the empirical analysis is:

$$\Delta \mathbf{z}_t = \sum_{i=1}^{4} \mathbf{B}_i \Delta \mathbf{z}_{t-i} + \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{x}_{t-1} + \textstyle\sum_{i=0}^{4} \mathbf{C}_i \Delta m_{t-i} + \mathbf{G}\mathbf{d}_t + \mathbf{u}_t \qquad (2)$$

where $\mathbf{z}_t = (s_t, i_t, \pi_t, y_t)'$, whereas the vector $\mathbf{x}_t$ includes as before the level of the real money stock, and $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are 4×3 and 5×3 matrices respectively.

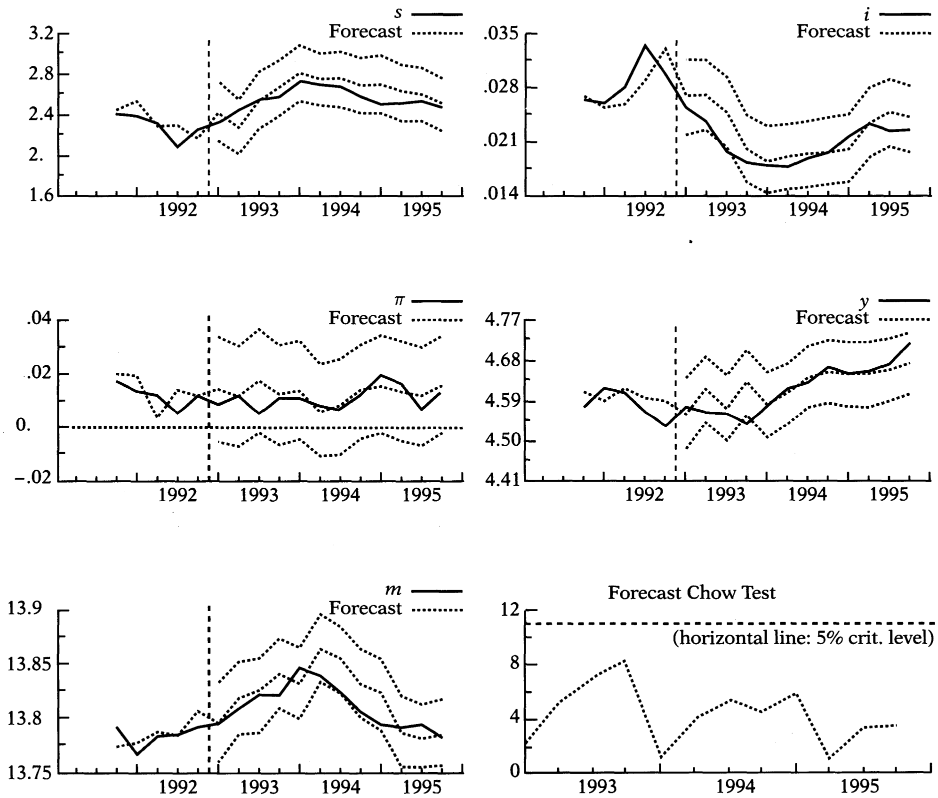

Fig. 3 - Forecasts from the Restricted VAR (1993-1995, with 95% Error Bands)

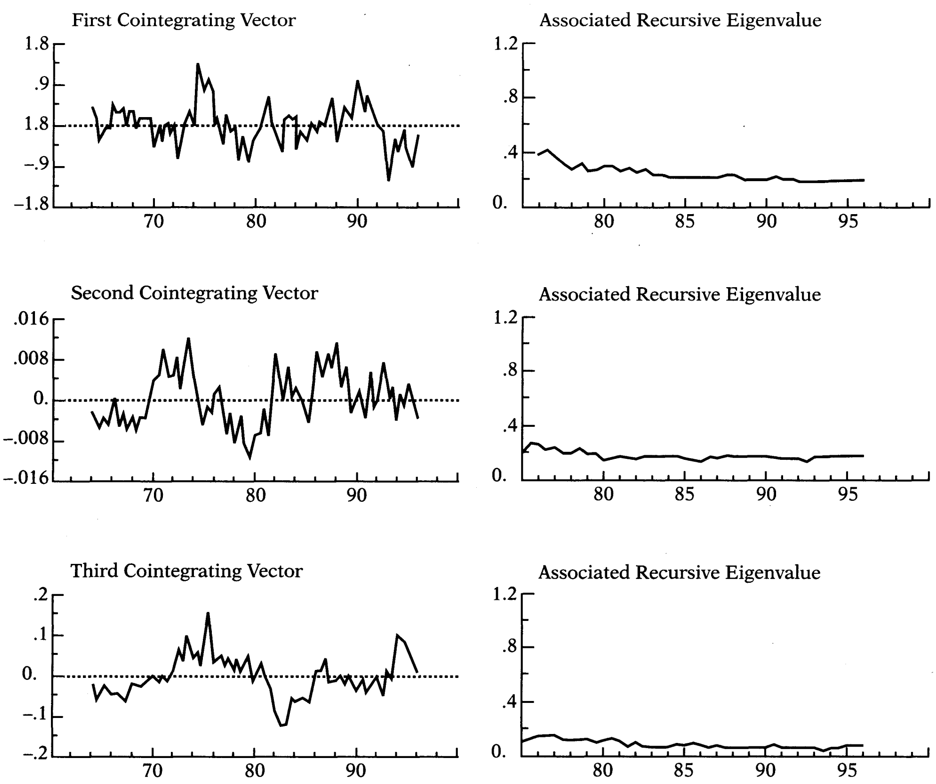

Fig. 4 - Unrestricted Cointegrating Vectors and Associated Recursive Eigenvalues

2.3. *The Cointegrated VAR System*

Prior to formulating testable structural hypotheses on the long-run relations, the conditional four-variable VAR system in (2) is subjected to a battery of diagnostic tests to assess its adequacy as a characterization of the time-series properties of the data. As Table 2A shows, no evidence of deviations from residual normality and of residual serial correlation is detected and only in the equation for $\Delta\pi$ the residuals display ARCH behaviour. Moreover, recursive estimation of the conditional system supports parameters' stability, leading to the conclusion that (2) is a congruent representation of the data generating process.

As it is well-known, a feature of the Johansen's procedure is that in the presence of multiple cointegrating vectors the estimates of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ cannot be immediately interpreted in terms of underlying behavioural parameters – long-run responses and short-run adjustment coefficients –; in fact, the estimated columns of $\boldsymbol{\beta}$ form an arbitrary base for (in our case) a three-dimensional cointegration subspace[5].

To provide an economically meaningful solution to this long-run identification problem, we formulated and tested several structural hypotheses on the cointegrating vectors, ending up with the following characterization of the long-run relations among the variables:

i) a relation between the real stock price and the inflation rate, suggested by the evidence presented in Fig. 2;

ii) a relation linking the output index to the real money stock and to the inflation rate;

iii) finally, a relation between the nominal interest rate and the rate of inflation; moreover, a cointegrating vector with coefficients $(1, -1)$ is specified, which amounts to imposing stationarity of the real interest rate. In the light of the already noted ambiguous results of the univariate unit-root tests for the real rate, we feel justified in imposing this additional long-run restriction on the data and formally testing it in a multivariate context[6].

This set of hypotheses imposes three overidentifying restrictions on the estimated $\boldsymbol{\beta}$ matrix and yields, after normalization of the resulting restricted vectors, three estimated coefficients, reported in Table 2B[7]. The associated standard errors show that all coefficient estimates are statistically well determined and the value of the likelihood ratio statistic (6.20 for an $\chi^2(3)$ distribution) does not reject the imposed set of restrictions (the p-value of the test is 0.10)[8].

The estimated *restricted* cointegrating vectors display several notable features. First, a strong negative effect of inflation on real stock prices is found, with an increase by one percentage point in the annual rate of inflation de-

[5] For any non-singular matrix $\boldsymbol{\xi}$ we have $\boldsymbol{\Pi} = (\boldsymbol{\alpha}\boldsymbol{\xi}^{-1})(\boldsymbol{\xi}\boldsymbol{\beta})$. The estimated columns of $\boldsymbol{\beta}$ may be then rearranged, with corresponding modifications of $\boldsymbol{\alpha}$, obtaining the same long-run matrix $\boldsymbol{\Pi}$.

[6] Furthermore, the stationarity of the real rate may be difficult to detect because of the large temporary declines of $i - \pi$ in the mid-1970s due to sharp inflationary shocks not immediately reflected in nominal interest rates. In what follows we note also the (qualitatively very similar) results obtained when the stationarity restriction on the real rate is not imposed.

[7] In the reported results, a constant term has also been restricted to enter the cointegrating vectors.

[8] Moreover, when the weak exogeneity of the real money stock is tested jointly with the above restrictions on the cointegrating vectors, the test statistic is 12.1 (with a p-value of 0.10 for an $\chi^2(6)$ distribution).

TABLE 2 - *The Cointegrated VAR System*

A. *Univariate Statistics on VAR Equations*

EQUATION	σ	R^2	AR $\chi^2(4)$	Norm.$\chi^2(2)$	ARCH$F(4,89)$	Het.$F(38,58)$
Δs	0.096	0.41	0.31 (0.98)	1.87 (0.39)	0.15 (0.96)	0.46 (0.99)
Δi	0.002	0.43	4.34 (0.36)	2.75 (0.25)	1.61 (0.18)	0.75 (0.82)
$\Delta \pi$	0.006	0.62	3.41 (0.49)	4.79 (0.09)	3.90 (0.01)	1.31 (0.18)
Δy	0.026	0.38	3.35 (0.50)	4.07 (0.13)	2.54 (0.04)	0.43 (0.99)

σ is the standard error of the regression. AR is the value of the Lagrange multiplier test for residual autocorrelation up to the fourth order; Norm. is the Jarque-Bera test for residual normality; ARCH is the test for autoregressive conditional heteroscedasticity (Engle, 1982) up to the fourth order; Het. is the F test for residual (unconditional) heteroscedasticity (White, 1980).

B. *Restricted Cointegrating Vectors*

	s	i	π	y	m	*const.*
$\boldsymbol{\beta}_1'$	1	0	46.09 (4.10)	0	0	–3.49 (0.15)
$\boldsymbol{\beta}_2'$	0	0	1.48 (0.65)	1	–0.71 (0.02)	5.12 (0.33)
$\boldsymbol{\beta}_3'$	0	1	–1	0	0	–0.001 (0.005)

($\boldsymbol{\beta}'$ matrix, standard errors of estimated coefficients in parentheses)

Estimated coefficients in the three cointegrating vectors have been normalized on s, y and i respectively.

LR test of restrictions on cointegrating vectors:
$\chi^2(3) = 6.20$ p-value: 0.10

C. *Loadings of the Restricted Cointegrating Vectors*

	$\boldsymbol{\beta}_1'\mathbf{x}_{t-1}$	$\boldsymbol{\beta}_2'\mathbf{x}_{t-1}$	$\boldsymbol{\beta}_3'\mathbf{x}_{t-1}$
Δs_t	–0.175 (0.042)	0.032 (0.306)	–1.988 (1.630)
Δi_t	–0.001 (0.001)	0.009 (0.005)	–0.091 (0.027)
$\Delta \pi_t$	0.003 (0.003)	–0.061 (0.019)	0.252 (0.101)
Δy_t	–0.016 (0.011)	–0.220 (0.079)	0.321 (0.422)

($\boldsymbol{\alpha}$ matrix, standard errors in parentheses)

D. *Correlation Matrix of VAR Residuals*

	Δs	Δi	$\Delta\pi$	Δy
Δs	1			
Δi	–0.202	1		
$\Delta\pi$	0.008	0.073	1	
Δy	–0.001	0.149	0.191	1

termining an 11.5% decrease in stock prices in the long-run. This is qualitatively similar to the result obtained by Blanchard (1993) for the US, where a 1% increase in inflation causes a 0.48% decrease in real prices in 1929-1953 and a 3.93% decrease for the more recent 1953-1992 period. Second, inflation has a negative effect also on the level of industrial production (an increase by 1% in annual inflation brings about a decrease in industrial output by nearly 0.4%). Output is also positively influenced by real money supply with a long-run elasticity of 0.71[9].

The graphs in the first column of Fig. 5 show the restricted cointegrating vectors, normalised on s, y and i respectively. In the graphs in the second column, each vector is split into two components: the actual value of the variable chosen for normalization (*e.g.*, s_t for the first vector) and the "fitted" value obtained using the estimated coefficients reported in Table 2B (*e.g.*, $3.49 - 46.09\pi_t$ for the first vector). The comparison of actual and fitted values suggests the presence of large and persistent deviations from the long-run equilibrium relations, particularly in the case of the interest rate-inflation cointegrating vector (the real interest rate). However, the differences between actual and fitted values from the estimated cointegrating relations are determined not only by deviations from long-run equilibrium but also by the short-run dynamics of the whole system. The latter effect may be eliminated by applying the estimated cointegrating coefficients (the elements of $\boldsymbol{\beta}$) to the residuals from a regression of the endogenous variables in levels ($\mathbf{x}$) onto the terms in (2) capturing the short-run dynamics and the deterministic elements (*i.e.*, Δz_{t-i}, Δm_t, Δm_{t-i}, d). The variables so obtained, $\mathbf{v}_t$, are then used to construct the cointegrating vectors "corrected" for the short-run

[9] When real rate stationarity is not imposed, estimation yields a coefficient of 0.55 (with a standard error of 0.22) on inflation in the third cointegrating vector. As for the other coefficients, the estimates imply a 9.5% drop in stock prices and a 0.8% decline in output in the face of an increase of annual inflation by 1%, whereas the output elasticity to money balances is unchanged at 0.71.

dynamics of the system: $\boldsymbol{\beta}'\mathbf{v}_t$[10]. These vectors, plotted in the first column of Fig. 6, give a clearer picture of the deviations from the long-run equilibrium. As in the preceding figure, in the second column the three vectors are split into their "actual" (v_s, v_y and v_i respectively) and "fitted" components.

Stock prices display huge fluctuations around their long-run relation with the inflation rate (first cointegrating vector) especially during the 1970s. Industrial production also shows deviations from its equilibrium relation with inflation and the real money stock (second vector) roughly consistent with the pattern of cyclical expansions and recessions occurred over our sample period. Also persistent fluctuations in the real interest rate (third vector) are detected in the mid-1970s, following the upsurge in inflation, and in the late 1980s and early 1990s, when the inflation rate followed a declining path.

To be validly interpreted as temporary deviations from the long-run equilibrium relations in the system, the cointegrating vectors must bring about error-correcting responses in (some of) the endogenous variables. From the estimated elements of the loadings matrix $\boldsymbol{\alpha}$, reported in Table 2C, we see (first column) that real stock prices display a statistically significant correcting reaction to deviations from the stock price-inflation relation. Moreover, the dynamic behaviour of the stock return index does not seem to be directly influenced by past disequilibia in the other two long-run relations (first row of the table). Movements in the real interest rate induce error-correcting responses both of the nominal interest rate and inflation, whereas deviations from the long-run relation among production, inflation and the money stock affect both inflation and output dynamics (second column).

Before studying a structural version of the model we evaluate the robustness of the above results to a change in the sample period. Indeed, starting the estimation period in 1963 might seem a very peculiar choice since at that time the Italian stock market was close to its historical peak. One might think that such a choice could be responsible for the lack of a long-run relation between the stock index and output, as the latter grew, with varying intensity, over the 1970s and 1980s, while the former remained substantially stagnant. A longer sample, including the 1950s too, characterised by high output growth rates and increases in real stock prices (Fig. 7), can provide a test on the robustness of the above results. When we re-estimate the system over the longer 1953-1995 period it yields very similar conclu-

[10] More precisely, this procedure naturally follows from the first step of the Johansen estimation method. Here the vectors $\Delta\mathbf{z}_t$ and $\mathbf{x}_{t-1}$ in (2) are regressed to $\Delta\mathbf{z}_{t-i}$ ($i = 1, ...4$), $\Delta\mathbf{m}_{t-i}$ ($i = 0, ..., 4$), $\mathbf{d}_t$. The obtained residual series, $\mathbf{v}_{0t}$ and $\mathbf{v}_t$ respectively, are then used in maximum likelihood estimation of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ from the (reduced rank) linear regression: $\mathbf{v}_{0t} = (\boldsymbol{\alpha\beta}')\mathbf{v}_t + \boldsymbol{\zeta}_t$

sions on the long-run properties of the series, confirming that the stock index is cointegrated with inflation only[11].

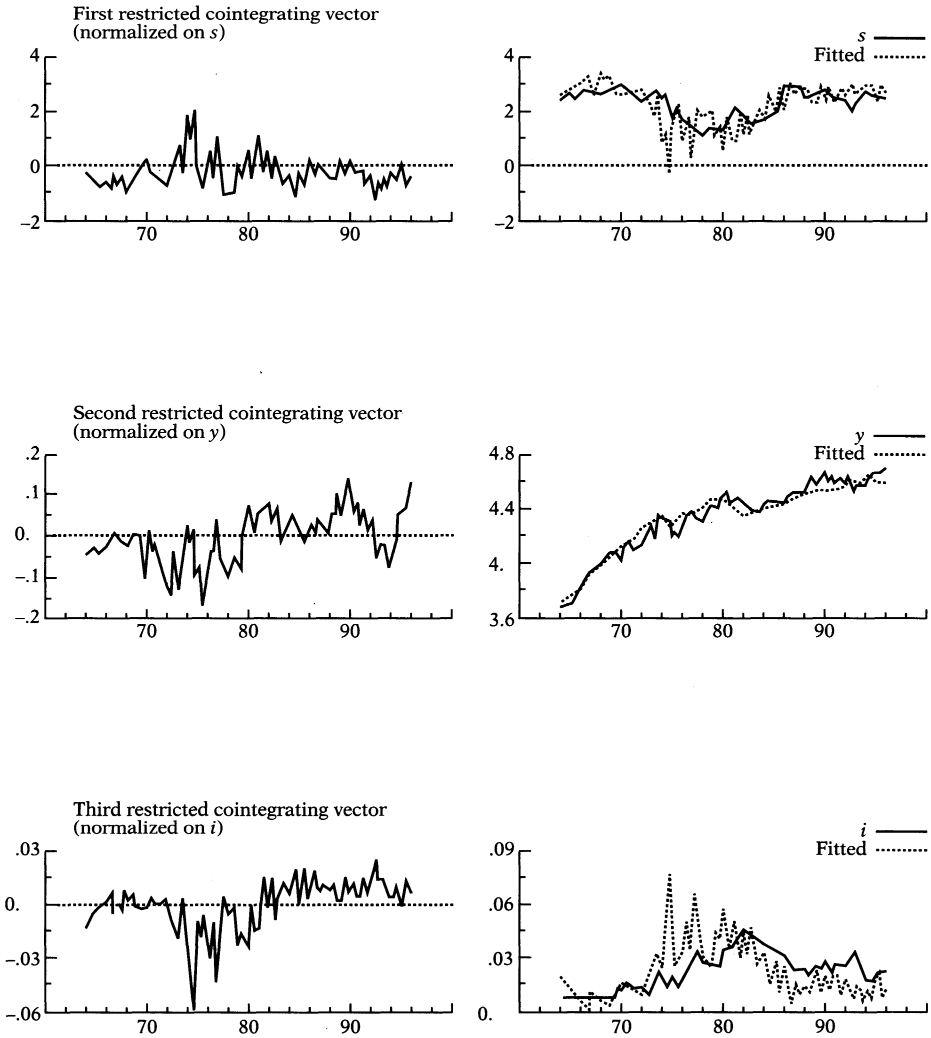

Fig. 5 - Restricted Cointegrating Vectors

[11] Our longer sample period starts in 1953(1), when the first observation of the industrial production index is available. Since no homogeneous series for the money stock is available over this longer sample, the system is estimated without the real money variable. The possibility of including money too (which plays an important role in the long run) in the analysis from the 1960s is the main reason which justifies our focus on the sample starting in 1963.

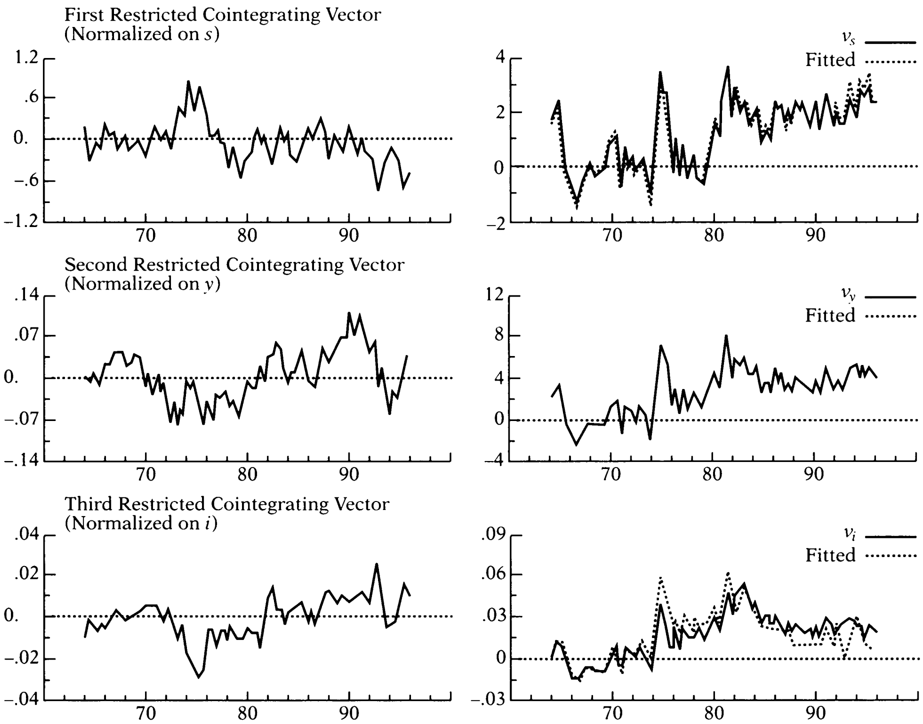

Fig. 6 - Restricted Cointegrating Vectors Corrected for Short-Run Dynamics

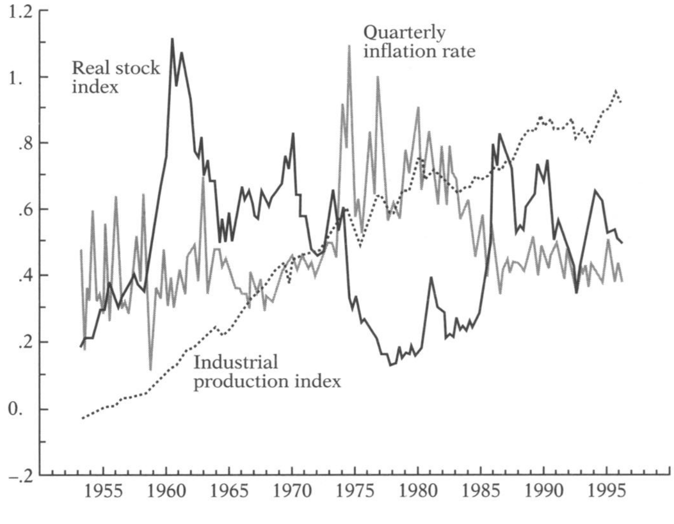

Fig. 7 - Real Stock Index, Quarterly Inflation Rate and Industrial Production Index: Italy 1953(1)-1996(1)

2.4. *"Structural" VAR Estimation and Dynamic Analysis*

Having analysed the long-run features of the data and provided a data-supported economic identification of the equilibrium relations linking stock prices, the interest rate, inflation and output, we are now able to simulate the VAR system in order to assess the dynamic response of all endogenous variables to structural disturbances. In so doing we have to face a second identification problem, this time involving the estimated residuals from (2), $\mathbf{u}$.

The nature of the problem is easily illustrated by recalling the reduced form nature of the VAR in (2). Such formulation of the system may be thought of as the reduced form of the following structural model:

$$\mathbf{A}\,\Delta\mathbf{z}_t = \sum_{i=1}^{4} \mathbf{B}_i^* \Delta\mathbf{z}_{t-i} + \boldsymbol{\alpha}^*\,\boldsymbol{\beta}'\mathbf{x}_{t-1} + \sum_{i=0}^{4} \mathbf{C}_i^* \Delta m_{t-i} + \mathbf{G}^*\mathbf{d}_t + \boldsymbol{\epsilon}_t \tag{3}$$

where the following definitions apply:

$$\begin{aligned} &\mathbf{B}_i^* = \mathbf{A}\,\mathbf{B}_i \quad (i = 1, ...4), \qquad && \boldsymbol{\alpha}^* = \mathbf{A}\,\boldsymbol{\alpha} \\ &\mathbf{C}_i^* = \mathbf{A}\,\mathbf{C}_i \quad (i = 0, ...4), && \mathbf{G}^* = \mathbf{A}\,\mathbf{G} \\ &\boldsymbol{\epsilon}_t = \mathbf{A}\,\mathbf{u}_t \quad E(\boldsymbol{\epsilon}_t\boldsymbol{\epsilon}_t') = \boldsymbol{\Sigma} && \end{aligned} \tag{4}$$

In (3) the elements of matrix $\mathbf{A}$ capture the contemporaneous (intra-quarter) relations among the variables in $\Delta\mathbf{z}$, and the specification of the (restricted) cointegrating vectors, $\boldsymbol{\beta}'\mathbf{x}_{t-1}$, is the same as in the reduced form (2). $\boldsymbol{\epsilon}$ is a vector of orthogonal shocks, the "structural" disturbances, with a diagonal covariance matrix $\boldsymbol{\Sigma}$. Each element of $\boldsymbol{\epsilon}$ may be then interpreted as an innovation in one specific endogenous variable of the system, uncorrelated with contemporaneous innovations in other variables. The dynamic response of the whole system to these innovations may finally be analysed using the traditional impulse response function and forecast error variance decomposition techniques.

Economic hypotheses may be imposed and tested on the matrix $\mathbf{A}$ in order to identify the structural disturbances. Technically, the reduced form VAR (2) yields ten estimated elements of the covariance matrix of the residuals $\mathbf{u}$ (the associated correlation matrix is reported in Table 2D). Therefore, at most ten parameters in $\mathbf{A}$ and $\boldsymbol{\Sigma}$ may be estimated: since the latter matrix contains the four standard errors of the structural disturbances, at most six elements of the contemporaneous matrix $\mathbf{A}$ may be obtained. Giannini (1992) and Hamilton (1994) formally establish necessary and sufficient conditions for identification, providing likelihood-ratio tests for the overidentifying restrictions if less than six elements of $\mathbf{A}$ are estimated.

Initially, we estimated the structural form of the system adopting a simple lower-triangular form for the **A** matrix and trying different orderings of the variables in **z**. In so doing we did not impose any overidentifying restrictions on **A**, and no test for the validity of the assumed recursive structure of the system is available[12]. The resulting patterns of dynamic responses to the disturbances in $\boldsymbol{\epsilon}$ are similar across different orderings of the endogenous variables, a result in accordance with the low values of the VAR residuals correlations in Table 2D. However, the statistical significance of only some elements in the estimates suggested the following form of the relation between reduced form and structural disturbances:

$$\begin{pmatrix} 1 & a_{12} & 0 & 0 \\ 0 & 1 & 0 & a_{24} \\ 0 & 0 & 1 & a_{34} \\ 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} u_s \\ u_i \\ u_\pi \\ u_y \end{pmatrix} = \begin{pmatrix} \varepsilon_s \\ \varepsilon_i \\ \varepsilon_\pi \\ \varepsilon_y \end{pmatrix} \tag{5}$$

In (5) the reduced form residual in the output equation is identified as a shock to production, uncorrelated with all other structural disturbances in the system, and having within-quarter effects on both the interest rate and inflation, captured by a_{24} and a_{34} respectively. Moreover, the stock price index contemporaneously reacts to the disturbance in the nominal interest rate (a_{12}).

This identification scheme is imposed on the **A** matrix and the resulting estimates (obtained by the makimum-likelihood method described in Giannini, 1992, and Hamilton, 1994) are reported in Table 3, together with the estimated values of the structural disturbances standard errors. A production shock tends to raise both the rate of inflation and the nominal interest rate, and disturbances to the latter variable (due either to structural interest rate shocks or to the effect of output shocks) have a sizeable negative effect on stock prices. On the other hand, no intra-quarter effect of inflation on real stock prices is detected. The set of overidentifying restrictions imposed on matrix **A** is not rejected, the likelihood ratio test yielding a p-value of 0.89.

In order to obtain a more complete picture of the dynamic interactions between stock prices and the other macroeconomic variables in the system, we can finally perform a dynamic analysis on the cointegrated structural VAR, with the identification scheme in (5), simulating the response of the system to (one-standard deviation) structural shocks in each endogenous variable. The resulting estimated impulse response functions, up to forty quar-

[12] This solution for the identification problem, based on a Choleski decomposition of the covariance matrix of the reduced form residuals, is common in traditional VAR analyses (*e.g.*, Sims, 1980).

TABLE 3 - *Structural VAR Estimation*

$u_s = -13.054$ (5.20)	$u_i + \varepsilon_s$	$\sigma_s = 0.094$ (0.006)
$u_i = 0.010$ (0.005)	$u_y + \varepsilon_i$	$\sigma_i = 0.002$ (0.0001)
$u_\pi = 0.045$ (0.020)	$u_y + \varepsilon_\pi$	$\sigma_\pi = 0.006$ (0.001)
$u_y =$	ε_y	$\sigma_y = 0.026$ (0.002)

(standard errors in parentheses)

Test of overidentification restrictions: $\chi^2(3) = 0.62$ p-value: 0.89

ters, are depicted in Fig. 8A and Fig. 8B, with 95% confidence intervals[13]. Two main features of the results are worth noting:

i) From panels A, B and D of Fig. 8, structural disturbances to stock prices, the nominal interest rate and production are not relevant in the long-run (none of the impulse responses at the longest horizon is statistically significant), in accord with our specification of the long-run equilibrium relations. However, in some cases they have a quantitatively important effect in determining the dynamics of the system at relatively short- and medium-term horizons. For example, an output disturbance determines, on impact, a reduction in the real interest rate which is readily reversed by opposite movements of the nominal rate and inflation in the following quarter. Subsequently, the nominal interest rate persists at a higher level for several quar-

[13] For the simulation analysis of the system to be valid, the contemporaneous rate of change of real money balances, Δm_t, must be *strongly* exogenous for the short-run parameters of the system. To test the *weak* exogeneity for those parameters, following ENGLE - HENDRY (1993), we formulated a time-series model for Δm_t using lags of all other variables. When added to the conditional VAR (2) the estimated residuals from the marginal model for Δm_t are found insignificant, ensuring the weak exogeneity of real money balances for the short-run parameters. However, estimation of the time-series model for Δm_t shows that this variable is marginally Granger-caused (at the 5% but not at the 1% significance level) by past acceleration in the inflation rate ($\Delta\pi$). This finding may cast doubt on the *strong* exogeneity of real money. Therefore, though presenting the simulation results for the conditional system, we also checked that their main qualitative features did not change when money is introduced as an additional endogenous variable, as in the complete system (1).

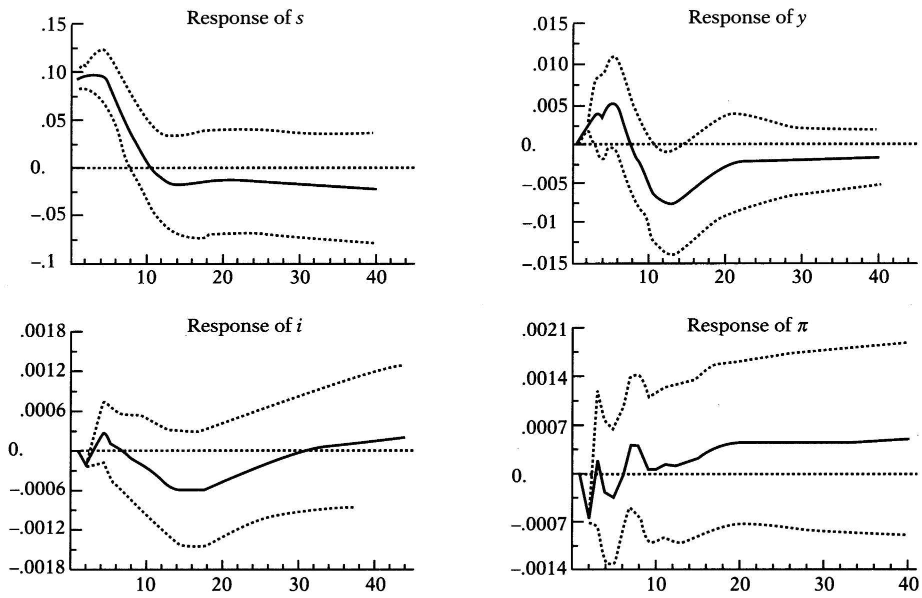

Fig. 8A - Impulse Response Functions to a Shock in s
(Forty Quarters, with 95% Confidence Intervals)

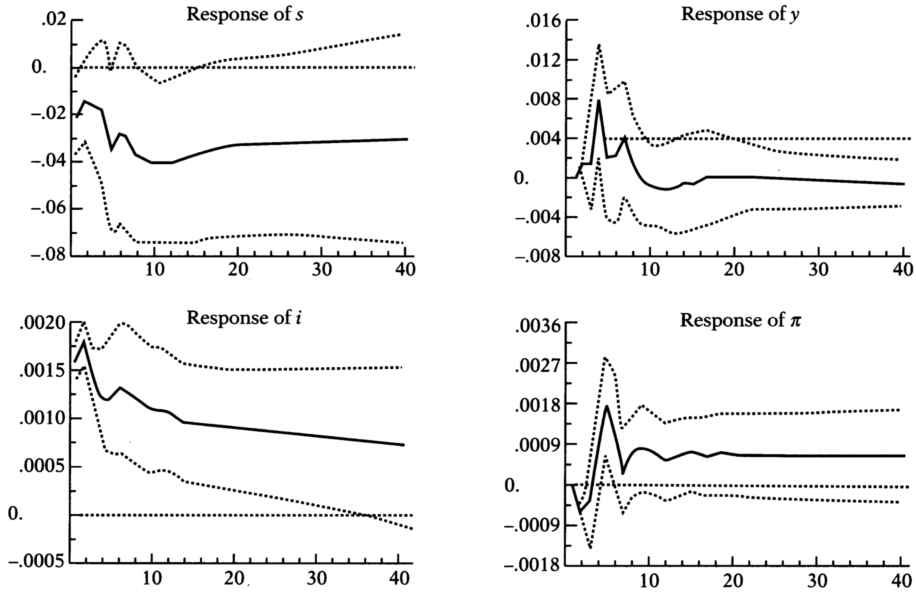

Fig. 8B - Impulse Response Functions to a Shock in i
(Forty Quarters, with 95% Confidence Intervals)

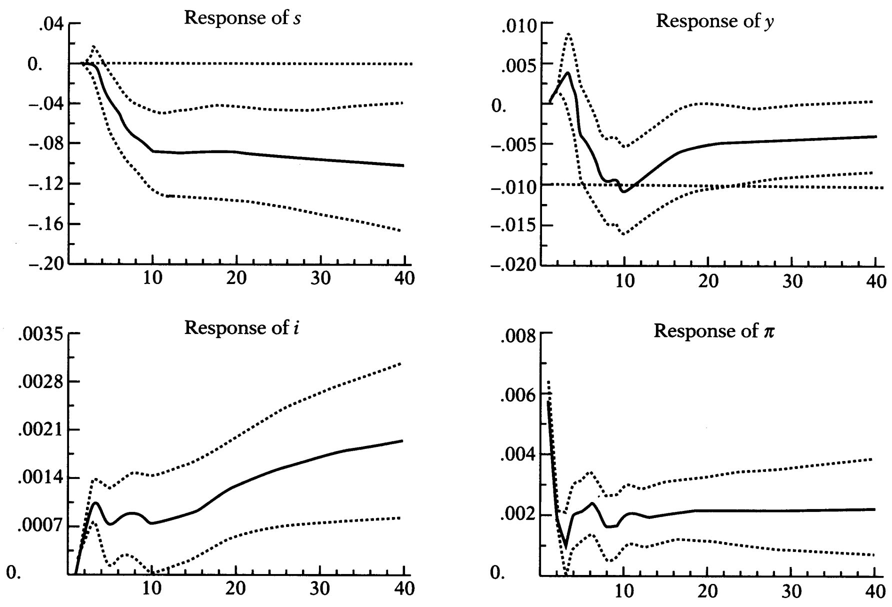

Fig. 8C - Impulse Response Functions to a Shock in π
(Forty Quarters, with 95% Confidence Intervals)

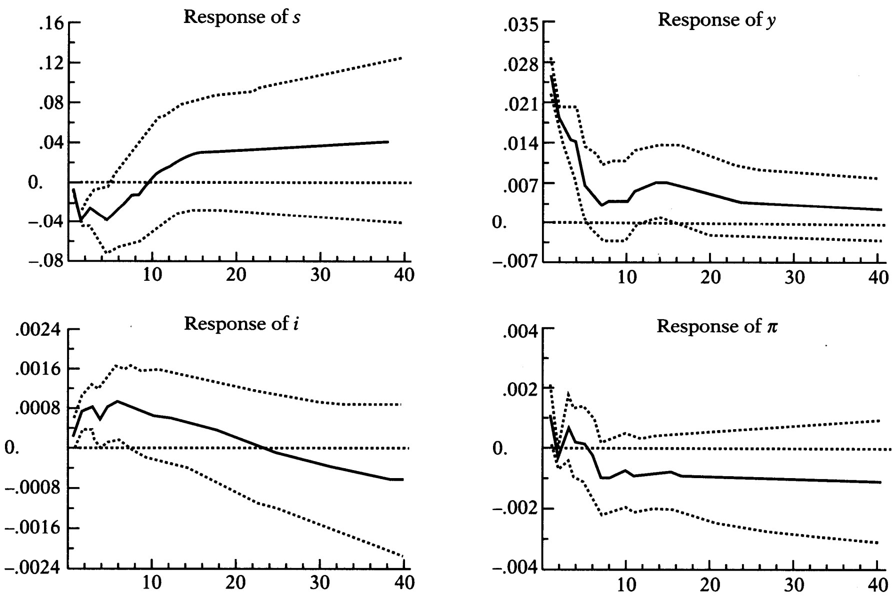

Fig. 8D - Impulse Response Functions to a Shock in y
(Forty Quarters, with 95% Confidence Intervals)

ters whereas the inflation rate converges rapidly to its initial level; the ensuing (temporary) increase in the real interest rate is mirrored by a decline in stock prices which lasts for four to six quarters.

ii) An inflation shock (panel C) has long-run, statistically significant, effects on stock prices, the nominal interest rate and output. The impulse response function of inflation to an own-shock is informative about the transitory or permanent nature of inflation disturbances. In the quarter following an inflation shock of about 0.6% (the estimated standard error of ε_π), the level of the inflation rate is about 0.2% higher than the initial level and remains, though with fluctuations, around that level at longer horizons. The response of output, slightly positive at the beginning, becomes negative afterwards, reaching a minimum ten quarters after the occurrence of the inflation shock. The nominal interest rate increases sharply in the quarter after the inflation disturbance and then gradually adjusts to the higher inflation rate, to restore the invariance of the real rate in the long run. Parallely, the stock index decreases steadily, reaching its lower long-run level ten quarters after the inflation shock.

Additional information on the relative importance of the various shocks can be gathered from the forecast error variance decomposition. For each endogenous variable z_i, from the moving average representation of the system (2) the variance of the forecast error $z_{i,t+k} - E_t z_{i,t+k}$ for horizons $k = 1, ..., 40$ is estimated. The fraction of these variances attributable to each of the variance of the structural shocks reflects the relative importance of the various disturbances in determining the overall variability of the endogenous series. Fig. 9A and Fig. 9B plot the results obtained (with 95% confidence intervals) for stock prices and the interest rate. Again, several points must be noted:

i) at short- and medium-term horizons the bulk of the variability in the real stock index is attributable to its-own shock, which, at a three-year horizon, still accounts for about 50% of the variability in s. The relative importance of inflation disturbances is increasing with the horizon, reaching 60% for forty-quarter forecasts;

ii) as the horizon increases, inflation shocks become more important also for the variability of the nominal interest rate (and industrial output), with slowly decreasing fractions attributed to the own-shocks to i and y respectively;

iii) at all horizons, most of the variability in the inflation rate is due to the own-shock to π, with almost negligible contributions from the other variables in the system.

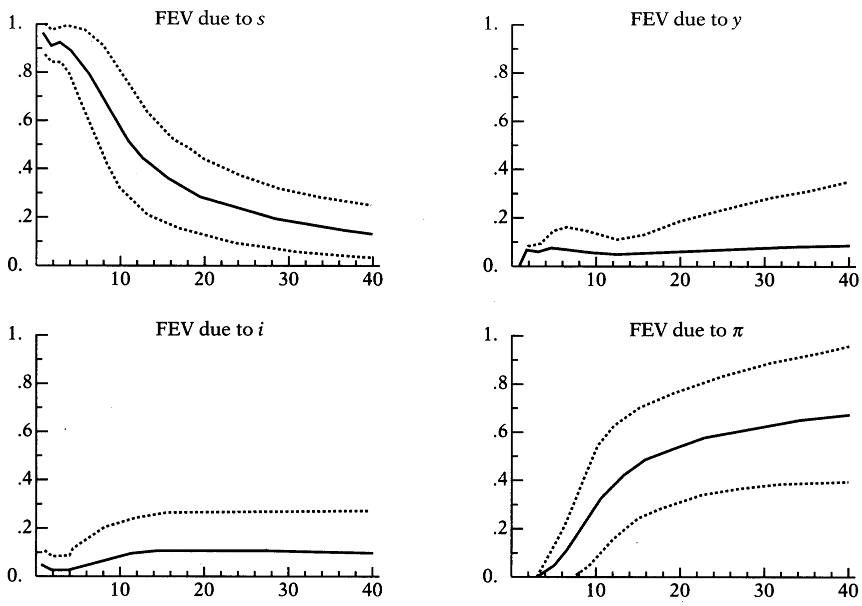

Fig. 9A - Forecast Error Variance Decomposition of s
(Forty Quarters, with 95% Confidence Intervals)

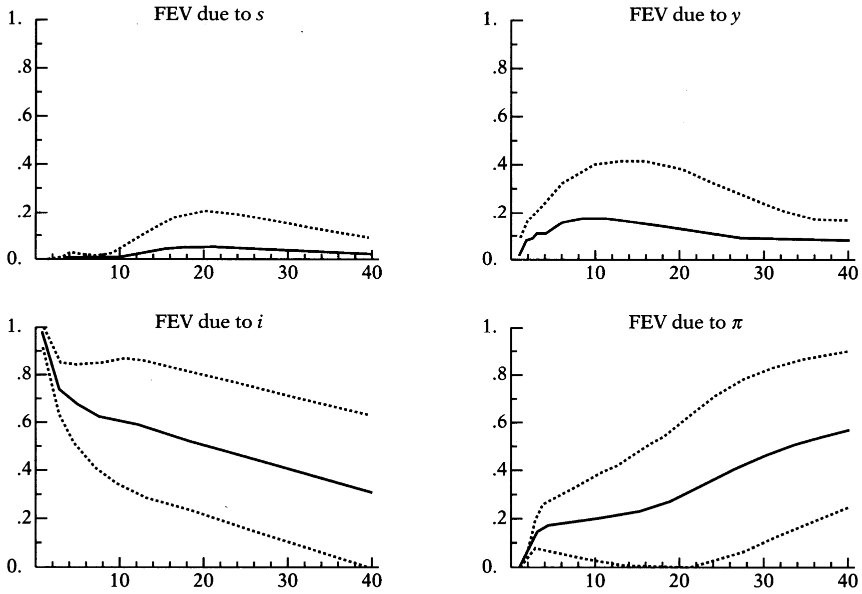

Fig. 9B - Forecast Error Variance Decomposition of i
(Forty Quarters, with 95% Confidence Intervals)

3. IMPLICATIONS FOR ASSET PRICING THEORIES

Until now we have made no use of theoretical asset pricing models. We have let the data tell the story of the relation between the stock market and macroeconomy. In this section we interpret our empirical findings in terms of the present value model, according to which stock prices represent the present discounted value of future dividends and expected returns. The linearised version of the present discounted model proposed by Campbell - Shiller (1988)[14] is:

$$\delta_t = E_t \sum_{j=0}^{\infty} \rho^j (h_{t+j} - g_{t+j}) + k \tag{6}$$

where k is a constant of linearization, δ_t is the log of the dividend-price ratio, h_{t+j} is the real rate of return on stocks, g_{t+j} is the rate of growth of real dividends. h_t can be thought of as the sum of the real rate of return on short-term assets r_t and a risk premium and can be written as:

$$h_t = r_t + rp_t \tag{7}$$

We have seen that the results about the real interest rate are somewhat conflicting: univariate unit-root tests give a result which is on the border of non-stationarity while the maximum-likelihood test performed within the multivariate system does not reject the hypothesis of stationarity. If one assumes that the dividend-price ratio is stationary in the long run, the present discounted value model allows to describe two alternatives which are both compatible with our empirical findings.

In the first, the Fisher relation does not hold and both the riskless real rate and the risk premium are non-stationary variables related to the inflation rate:

$$r_t = \beta\pi_t + \varsigma_t$$

$$rp_t = \gamma\pi_t + \upsilon_t$$

[14] In the US case CAMPBELL - SHILLER (1988) have performed various tests of this relationship under the assumption (not rejected by the data) that the level of the real stock index is cointegrated with the index of real dividends, with a $(1, -1)$ cointegrating vector. The resulting (log) dividend-price ratio is therefore stationary and consistent with the time-series properties of the real rate of interest and the rate of growth of real dividends. We cast the empirical model in terms of the index augmented of the distributed dividends (in the US case we find cointegration between this variable with inflation and industrial production). Therefore, our quantitative results are not directly comparable with those of Campbell and Shiller because of different definitions of the variables.

where $\beta \neq 0$ and $\gamma \neq 0$ and ς_t and υ_t are stationary series. It follows that the discounting term described in (7) is:

$$r_t + rp_t = (\beta + \gamma)\pi_t + \upsilon_t + \varsigma_t$$

Given the $I(1)$ nature of the inflation rate, the discount rate is stationary only if $\gamma = -\beta$. This model would imply that inflationary periods are associated with a lower real interest rate and a larger risk premium in such a way as to maintain cointegration between stock prices and dividends.

In the second alternative, both the real interest rate and the risk premium are stationary variables. This may be more appealing from an economic point of view and may be a plausible generalization out-of-sample.

A final issue we briefly address is whether the above results are consistent with some of the theories on the relationship between inflation and the stock market described in Section 1. We cannot perform a formal testing, but can at least evaluate the broad compatibility of some of the models with the Italian data.

On the one hand, the model by Fama would seem to be compatible with the impulse response function for the inflation shock, since there is a persistent negative relation between economic activity and a shock to inflation. This is exactly the hypothesis of Fama who considers the reaction of an efficient stock market to news about macroeconomic variables. However, it seems that this is not consistent with the zero covariance between the stock index and inflation within the quarter. An efficient forward-looking financial market should immediately react to an inflation shock which is associated with such strong effects on the real economy. On the other hand, the Modigliani-Cohn hypothesis seems to be more consistent with the data, since from the impulse response functions we observe a gradual increase in the nominal interest rate and a gradual decrease in the stock market, as if agents responded to the level of the nominal interest rate.

Conclusions

We have analysed the relation between the stock market and some macroeconomic variables on the basis of a cointegrated structural VAR. Including additional macroeconomic variables like industrial production, the rate of interest and real money supply helps to improve the precision of the estimated long-run relations. The main finding is a strong negative long-run relationship between the rate of inflation and the level of the real stock index.

However, the absence of a long-run relationship between stock prices and an output measure may suggest caution in the interpretation of the results, especially for out-of-sample extrapolations. In fact, from the estimated long-run relation, a prolonged period of stable inflation should see no change in the real value of the stock index. In our opinion, this suggests the existence of a potential problem with the Italian experience in the considered sample, due to the importance of the major inflation shocks of the 1970s and 1980s. The magnitude of these shocks might have been so large as to hide other long-run relations with important macroeconomic variables such as output. It is very likely that, as the performance of the Italian economy tends to become more similar to that of other developed countries, the fundamental relations between the stock market and macroeconomic variables will tend to emerge, as it is the case, for example, of the US. This should represent a warning to those willing to rely on historical evidence of the Italian economy for extrapolation over the long run.

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