Stock market returns, inflation and monetary regimes

Giuseppe Grande, Alberto Locarno and Massimo Massa*

Introduction

Since Fisher's initial contribution in the early thirties, several studies have looked at the ability of different assets to provide a hedge against inflation.[1] The Fisher hypothesis, relying on the idea that the monetary and real sectors of the economy are largely independent, states that expected asset returns should move one-to-one with expected inflation. In principle this hypothesis is applicable to any instrument that can serve to transfer wealth through time, but it should especially apply to assets representing physical capital, such as real estate and shares in the capital of a company. These assets should also provide a hedge against unexpected inflation. However, empirical studies have often concluded that the Fisher hypothesis is not well supported by the data; more surprisingly, its failure appears clearest for equities.

Theoretical as well as applied research has shown that the relation between stock prices and inflation is influenced by economic policy, and by monetary policy in particular. This paper focuses on the relation between stock returns, inflation and monetary policy. The working hypothesis is that the market interprets inflation differently according to a latent variable that captures the effects of shifts in the stance and the credibility of monetary policy, as well as those of changes in the institutional framework in which the central bank operates. Financial markets react differently to inflation news, depending on the monetary policy regime they perceive to be the prevailing one. When the central bank is thought to be strongly committed to price stability, even a small surge in inflation expectations induces the market to fear a strong monetary policy reaction, which would lead to higher interest rates, lower economic activity and lower expected dividends. As a consequence, stock prices drop, and the negative relationship between stock returns and expected inflation usually found in the literature obtains. This is essentially the so-called proxy hypothesis proposed by Fama (1981) and developed by subsequent studies, as will be explained in the next section.

The empirical framework adopted in this paper – applied to data on the Italian stock exchange covering the last twenty years – relies upon the present-value relation of Campbell and Shiller, and makes use of a Markov-switching model to identify regimes associated with different policy environments. The analysis focuses on the inflation information contained in stock returns, and does not address the issue of the possible effects of equity prices on real activity.[2]

After presenting a brief review of the main arguments put forward to explain the failure of the Fisher hypothesis in stock markets, we provide an initial assessment of the relation between asset returns and inflation in Italy in the second section. Then we present the VAR model with Markov switching and the decomposition of the βs of a portfolio according to the present-value relation. The last section discusses the methodological issues raised in the paper and sets out the main conclusions.

---

* Banca d'Italia, Research Department.

1 See Fama and Schwert (1977) and the survey by Rovelli (1984).

2 A careful comparison of the balance sheets of household and enterprise sectors in the major industrial countries can be found in Kneeshaw (1995).

1. Different explanations of the relation between stock returns and inflation

Various explanations of the failure of the Fisher hypothesis when applied to the stock market try to interpret empirical results in terms of spurious correlations and omitted variables. Some studies have also addressed the issue on theoretical grounds.

Fama (1981) argued that the sign on inflation is due to the fact that inflation acts as a proxy for omitted variables. Given that high inflation anticipates low growth and that there is a positive relationship between expected economic growth and stock prices, there should be a negative relationship between inflation and stock prices. According to Stulz (1986), an increase in expected inflation leads to a fall in the real wealth of households, which in turn lowers the real interest rate and the expected return on the market portfolio. Geske and Roll (1983) relate the high rates of inflation during recessions to counter-cyclical monetary policy actions. The central bank responds counter-cyclically to real activity shocks: a drop in real activity leads to a higher public deficit which, in turn, induces an increase in money growth to the extent that the debt is monetised. An unanticipated drop in stock prices signals this chain of events, with the counter-cyclical expansion of the money supply reinforcing the "proxy" mechanisms proposed by Fama.

The perception of a clear link between stock prices and monetary and fiscal policy induced Kaul (1987) to focus on the relationship between monetary regimes and the Fisher equation. In particular, he showed how the counter-cyclical monetary policy regime in the post-war period generated a strong negative relationship between stock returns and changes in expected inflation; conversely, the relationship was positive under pro-cyclical monetary policy regime in the thirties. Furthermore, Kaul (1990) found evidence that the negative relation between stock returns and changes in expected inflation in the post-war years is particularly strong during interest rate regimes. More recently, Balduzzi (1993) proposed a VAR decomposition that reinterprets the proxy hypothesis, showing that both inflation and stock returns tend to anticipate future interest rate changes, albeit in opposite directions. Groenwald et al. (1997) examine the matter within the framework of a small macroeconomic model and find that the reduced form for the interest rate equation is much more complex than that used by Fama and Schwert and requires a larger set of variables to be explicitly taken into account. Though they propose and estimate a more refined specification, they find that the negative sign of the correlation coefficient survives the extension to the full model.

Söderlind (1997) uses a modified version of a model by Fuhrer and Moore (1995) to show that the sign and size of the correlation between stock returns and inflation in a closed economy depend on the objective function of the central bank. Suppose that (i) inflation is persistent but can be controlled via a Phillips effect; (ii) output is negatively related to real interest rates through an IS-type relation; (iii) the interest rate is set by the monetary authorities; and (iv) there are exogenous inflation shocks. Under these assumptions, if the central bank wants to stabilise output, it will move the nominal rate so as to keep the real interest rate constant: the nominal interest rate then entirely reflects changes in expected inflation and the Fisher effect is complete. If the central bank targets inflation instead, it will use the nominal rate in order to allow the real rate to move as much as is required to stabilise expected and actual inflation. If this policy is successful, the nominal interest rate will be mainly correlated with the real rate and the Fisher relation will not be satisfied.

Focusing on econometric issues, Evans and Lewis (1995) reformulate the Fisher puzzle in terms of a time-varying model. They do not search for an economic rationale for the failure of the Fisher hypothesis for bond rates, but try to explain it in terms of small sample biases induced by the infrequency of shifts in the inflation process during the post-war period.

In sum, previous literature has pointed out that contemporaneous regressions of stock returns on inflation expectations, while simple and useful, do not shed light on the channels through which macroeconomic news affects asset prices. Moreover, the co-movements of inflation and stock prices are clearly influenced by monetary policy and, more generally, by the policy environment.

Concerning the first issue, the proxy hypothesis put forward by Fama can be interpreted as an attempt to remove the influence of future output growth; similarly Geske and Roll try to neutralise the effects of monetary policy by including money supply as an additional explanatory variable in simple regressions of stock returns on expected inflation. More generally, Groenewold et al. stress that once we interpret the Fisher relation as the reduced-form equation of a macroeconomic model, we must allow for a large number of additional variables affecting stock returns in addition to the rate of inflation, namely the exchange rate and government consumption. Regarding the second issue, Kaul (1987, 1990) acknowledges that the correlation between stock returns and inflation is altered by policy actions and suggests dividing the sample period according to shifts in the policy regime in order to allow a proper evaluation of the Fisher effect. Finally, Söderlind claims that "the Fisher effect [...] is probably not carved in stone, but is likely to depend on monetary policy". As mentioned above, if the central bank wants to stabilise output, movements in nominal rates will parallel movements in inflation, while if it aims to preserve price stability the yield curve will not provide meaningful information about inflation expectations.

A clear example of the shortcomings of reduced-form models of stock returns is provided by Campbell and Ammer (1993). They cite the case of the reaction of the stock market to news about industrial production. This association could reflect either a link with changing expectations about future cash flows or some correlation with movements in future discount rates, perhaps because both industrial production and stock prices respond to interest rate changes. The only way to distinguish these channels is to deal explicitly with the relations linking stock prices to future dividends and required returns. This is the approach adopted in this paper.

2. Asset returns and inflation: a first step in the empirical analysis

As a first step in the empirical analysis, we replicate the approach developed by Fama and Schwert to draw a general picture of the relationship between asset returns and inflation in Italy.

The Fisher theory of interest assumes that the monetary and real sectors of the economy are largely independent. Expected real returns are uncorrelated with expected inflation, being determined by non-financial factors such as productivity of capital, time preferences and risk aversion: expected asset returns therefore move one-to-one with expected inflation. However, in order to assess whether financial assets provide a hedge against inflation, it is also necessary to analyse how nominal returns react to unexpected inflation.

To address these issues in a consistent framework, Fama and Schwert begin with the following equation (see equation (3) in Fama and Schwert (1977)):

$$E(R_{jt} \mid \phi_{t-1}) = E(i_{jt} \mid \phi_{t-1}) + E(\pi_t \mid \phi_{t-1}) + \gamma_j [\pi_t - E(\pi_t \mid \phi_{t-1})] \tag{2.1}$$

where R_{jt} is the nominal return on asset j from time t-1 to time t, ϕ_{t-1} is the information set at t-1, π_t is the inflation rate from time t-1 to time t and i_{jt} is the equilibrium real return.

On the basis of equation (2.1) and having a measure of the expected inflation rate, $E(\pi_t \mid \phi_{t-1})$, tests of the joint hypothesis that markets are efficient[3] and expected real returns and inflation are uncorrelated can be obtained from the following regression model:

$$R_{jt} = \alpha_j + \beta_j \pi_t^e + \gamma_j (\pi_t - \pi_t^e) + \varepsilon_{jt} \tag{2.2}$$

---

[3] That is, agents' expectations are the best possible assessment of the expected value of random variables given available information.

where, for simplicity, $\pi_t^e \equiv E(\pi_t \mid \phi_{t-1})$. If the coefficient β is not significantly different from 1, then the Fisher hypothesis cannot be rejected and the asset provides a complete hedge against *expected* inflation; if $\gamma = 1$, then the asset is a complete hedge against *unexpected* inflation; finally, if both β and γ are not significantly different from 1, then the asset is a complete hedge against *actual* inflation, and ex-post real returns and inflation are uncorrelated.

Fama and Schwert point out that the relation between nominal returns and unexpected inflation is not the same for all assets: while it is generally believed that real estate, common stocks and human capital[4] are hedges against both anticipated and unanticipated inflation, short-term securities, with fixed nominal payments, are entirely exposed to nominal shocks.

Fama and Schwert estimate the regression (2.2) on monthly US data for the period between January 1953 and July 1971. Building on previous work by Fama, the return on Treasury bills with a residual maturity of one month is used as a proxy for the expected value of inflation. They find that: (i) Treasury bills and bonds provide a hedge against expected inflation; (ii) private residential real estate hedges against expected as well as unexpected inflation; and (iii) labour income shows a weak correlation with inflation. The most striking result is obtained for common stocks, whose nominal returns appear to be negatively related to expected and, probably, unexpected inflation.

A crucial role in this kind of test is played by the measurement of expected inflation. Santoni and Moehring (1994) claim that the puzzle shown by Fama for stock returns can be accounted for once inflation expectations are properly measured. Three proxies for expected inflation have been used in the literature:

1. the nominal return on Treasury bills (Fama and Schwert, Mishkin (1990), and Kaul (1990));[5]
2. survey data on inflation expectations (Bomberger and Frazer (1981));[6]
3. expected inflation defined on the basis of a set of previously specified variables. Balduzzi (1993), for example, explicitly defines expected inflation by inverting a rational-expectations version of the standard quantity theory equation.

We apply the approach suggested by Fama and Schwert to Italian data on five different assets:[7] 3, 6 and 12-month Treasury bills; Treasury bonds;[8] and the value-weighted Milan stock exchange index. With regard to inflation expectations, since none of the aforementioned approaches is without shortcomings or is uniformly superior to the others, we try different alternatives. We use both the Forum-ME survey data and the fitted values of the projection of inflation on its own lags and the percentage changes in the exchange rate and industrial production; an additional attempt is made on

---

4 They adopt the rate of change in per capita labour income as a proxy for the nominal return on human capital (this does not account for changes in capital values). If real labour income is to be independent of the price level, the measure must reflect inflation rate movements.

5 This choice is based on two hypotheses: (i) the expected real return on the short-term bill is constant through time and (ii) the market is efficient, so that the nominal return on the bill is equal to the constant expected real return plus the expected inflation rate; that is, it coincides with expected inflation apart from a constant factor.

6 Some shortcomings are inherent in the use of survey data, the main one being that the sample may not be representative of the whole economy. Also, it is certainly true that economists have better theories of how people take actions than they do of how they answer questions on surveys. Finally, unlike in a market where the participants back up their statements with money, it is less clear what it means when someone just expresses opinions about inflation or other variables.

7 The data used in the paper are described in Appendix 1.

8 Average yield on Treasury bonds with at least one year to maturity.

quarterly data by using the 3-month Treasury bill rate. The period covered runs from February 1979 to May 1997 for monthly data and from the second quarter of 1979 to the first quarter of 1997 for quarterly data.

The results for monthly data are reported in Table 1a. For each asset, two regressions are estimated according to the proxy considered for expected inflation. The estimated values of the parameters and their standard errors are shown in the first three columns, while the probabilities that β and γ are equal to one are shown in the last two columns.

The hypothesis that these assets are hedges against expected inflation is rejected with respect to Treasury bills and, to a lesser extent, Treasury bonds, but the estimated value of the parameter β is positive and markedly different from zero, suggesting that the assets provide at least a partial hedge against expected inflation. Conversely, the parameters associated with unexpected inflation are never significantly different from zero.

It is very important to observe that the test on β is actually a joint test of three hypotheses: lack of correlation between the expected values of the real rate and inflation; market efficiency; and the Fisher hypothesis. The above results must be interpreted with caution, because the rejection of the null hypothesis could be due to the fact that a fully developed market for government securities in Italy emerged only at the end of the eighties.[9]

A further warning is due because the effect of taxes on capital income has not been taken into account. Since financial assets are usually taxed, a change in inflation that is fully transmitted to nominal interest rates does not leave the lender with the same pre-inflationary real return: nominal returns have to move more than proportionately to leave the after-tax real rate unaffected. A proper treatment of this issue, which is complicated by the fact that tax incidence is not the same for all investors and assets, is beyond the scope of this study. Note, however, that the coefficient on inflation in equation (2.2) has to be greater than one if after-tax returns are to provide a complete hedge against inflation.

When the inflation forecast is measured by survey data or by the fitted values of a time-series model, the results obtained with quarterly data are very close to those found with monthly figures (Table 1b). When possible, the regressions on quarterly data also make use of the 3-month Treasury bill as a proxy for expected inflation. In this case, the hypothesis of a complete hedge against expected inflation cannot be rejected at standard confidence levels, while the estimated value for γ is still not consistent with the hypothesis of perfect coverage against unexpected inflation.

As regards stocks, in all models neither β nor γ are significant at standard confidence levels and the proportion of the variance of stock returns explained by the regression expectations is very low (about 2%). However, the estimated effect of inflation forecasts on stock returns is positive, as expected. This is an important difference with the results obtained on US data with similar methodologies.

To check for instability in the coefficients and to see how the estimated relation between nominal returns and expected inflation has moved through time, rolling regressions on a ten-year window, spanning the whole sample period, have been run on quarterly data; stability analysis has only been applied to the regressions that use the inflation forecasts of the Forum-ME survey. The estimated values for β and its confidence bands are plotted in Figure 1; the horizontal dashed line

---

9 A screen-based secondary market for government securities was introduced in May 1988 and grew quickly. The volume of transactions in Treasury bills on the secondary market has always been very thin. For this reason, the returns on Treasury bills used in the paper are those determined through competitive auctions on the primary market. It must be noted that until March 1989 the Treasury set a floor for the bid price, which often turned out to be binding; this constraint lessened the link between the average yield at auction and agents' expectations. In March 1989 the lower bound for bids was removed for all maturities; Grande (1994) provides evidence that the ability of the primary Treasury bills market to signal agents' expectations improved since that date.

indicates the points where the parameter is equal to 1: when the line is inside the confidence band, the hypothesis of the asset being a complete hedge against expected inflation cannot be rejected. This appears to be true for government securities since the eighties. However, the estimated β varies considerably over the period, and its standard error clearly shows a tendency to widen. The rolling estimates of the β parameter for stocks confirm the failure of this simple test of the Fisher hypothesis for the Italian stock exchange.[10]

Table 1a

Effects of expected and unexpected inflation on asset returns in Italy

Expected inflation proxy	α	β	γ	R^2	σ	H_0: β=1 complete hedge against expected inflation	H_0: γ=1 complete hedge against unexpected inflation
			(a) 3-month Treasury bills				
Forum-ME survey	0.006 (0.0003)	0.639 (0.0357)	-0.045 (0.0335)	0.608	0.002	0.00	–
AR model	0.007 (0.0002)	0.417 (0.0254)	-0.072 (0.0409)	0.552	0.002	0.00	–
			(b) 6-month Treasury bills				
Forum-ME survey	0.006 (0.0003)	0.637 (0.0367)	-0.041 (0.0344)	0.594	0.002	0.00	–
AR model	0.007 (0.0002)	0.420 (0.0257)	-0.077 (0.0414)	0.551	0.002	0.00	–
			(c) 12-month Treasury bills				
Forum-ME survey	0.006 (0.0002)	0.616 (0.0354)	-0.033 (0.0332)	0.596	0.002	0.00	–
AR model	0.007 (0.0002)	0.407 (0.0249)	-0.062 (0.0401)	0.549	0.002	0.00	–
			(d) Treasury bonds				
Forum-ME survey	0.007 (0.0002)	0.586 (0.0364)	-0.024 (0.0341)	0.561	0.002	0.00	–
AR model	0.008 (0.0002)	0.393 (0.0251)	-0.060 (0.0404)	0.528	0.002	0.00	–
			(e) Stocks				
Forum-ME survey	-0.003 (0.0113)	2.748 (1.576)	-0.130 (1.477)	0.006	0.069	–	–
AR model	0.001 (0.0085)	2.093 (1.046)	-0.966 (1.684)	0.011	0.069	–	–

Notes: Equation (2.2) is run on monthly data for the period 1979:2-1997:5. The statistic R^2 is adjusted for the degrees of freedom. σ is the standard error of the regression. Numbers in parenthesis are parameter standard errors. The last two columns show the probabilities of being wrong in rejecting the indicated hypotheses; they are reported only for those cases in which the estimated parameter is different from zero at a 5% confidence level. A description of the data is given in the Appendix.

---

[10] For almost the whole sample period, the hypothesis that the value of the parameter is equal to zero cannot be rejected.

All in all, the results in Table 1 confirm the findings in Fama and Schwert, though there is evidence that the relation between stock returns and expected inflation is positive in Italy as expected. These results signal that the Fisher hypothesis is not well supported by the empirical evidence, especially for stock returns.

In the rest of the paper we will try to model the relation between inflation and stock returns more accurately, taking into account the role played by policy regimes.

Table 1b

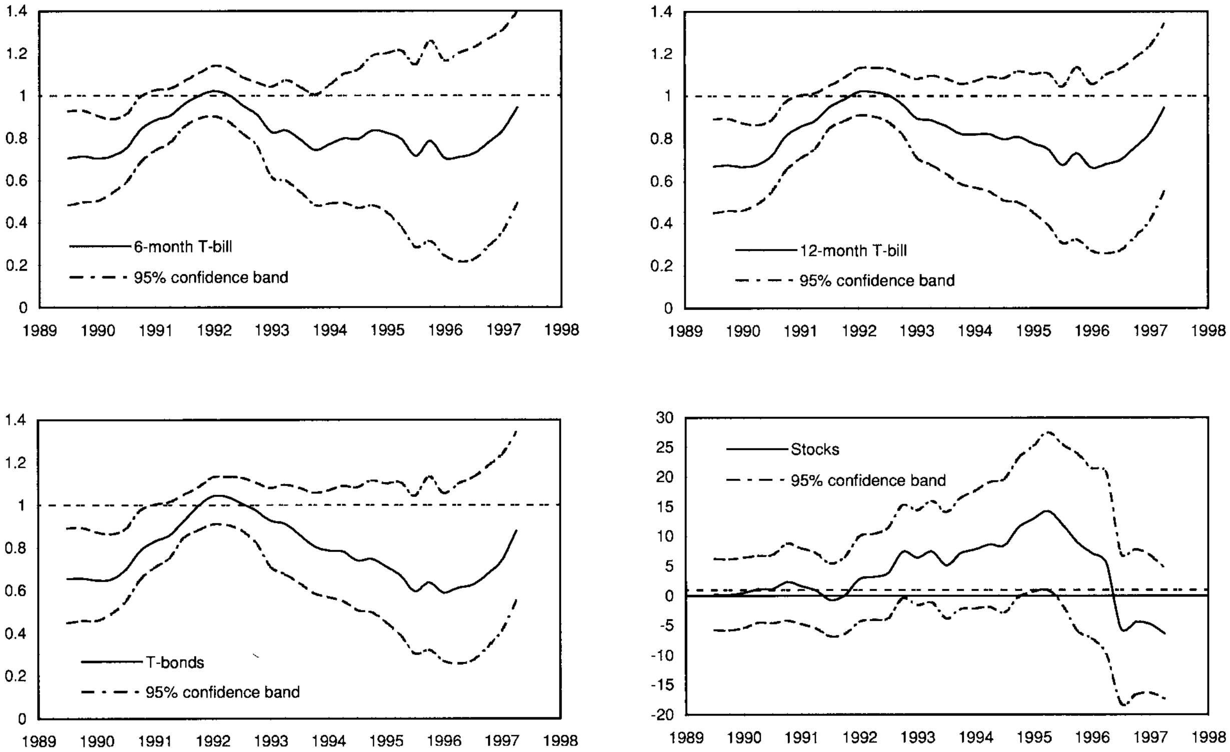
Effects of expected and unexpected inflation on asset returns in Italy

Expected inflation proxy	α	β	γ	R^2	σ	H_0: β=1 complete hedge against expected inflation	H_0: γ=1 complete hedge against unexpected inflation
(a) 6-month Treasury bills							
3-month Treasury bill	0.003 (0.0016)	0.944 (0.0487)	0.141 (0.0307)	0.847	0.003	25.71	0
Forum-ME survey	0.018 (0.0016)	0.689 (0.0757)	-0.124 (0.1042)	0.576	0.005	1.06	–
AR model	0.022 (0.0012)	0.458 (0.0480)	-0.102 (0.1060)	0.561	0.005	0.00	–
(b) 12-month Treasury bills							
3-month Treasury bill	0.004 (0.0014)	0.921 (0.0419)	0.135 (0.0264)	0.877	0.003	6.23	0
Forum-ME survey	0.018 (0.0015)	0.661 (0.0719)	-0.108 (0.0989)	0.583	0.005	0.12	–
AR model	0.022 (0.0011)	0.441 (0.0458)	-0.081 (0.1011)	0.565	0.005	0.00	–
(c) Treasury bonds							
3-month Treasury bill	0.005 (0.0014)	0.887 (0.0426)	0.124 (0.0269)	0.864	0.003	1.03	0
Forum-ME survey	0.020 (0.0015)	0.614 (0.0726)	-0.082 (0.0999)	0.547	0.005	0.00	–
AR model	0.024 (0.0011)	0.418 (0.0458)	-0.066 (0.1011)	0.537	0.005	0.00	–
(d) Stocks							
3-month Treasury bill	0.032 (0.0678)	1.194 (2.034)	2.248 (1.281)	0.019	0.128	–	–
Forum-ME survey	0.015 (0.0396)	1.427 (1.900)	2.754 (2.615)	0.018	0.128	–	–
AR model	-0.002 (1.310)	2.262 (1.181)	0.356 (2.618)	0.023	0.127	–	–

Notes: Equation (2.2) is run on quarterly data for the period 1979:II-1997:I. The statistic R^2 is adjusted for the degrees of freedom. σ is the standard error of the regression. Numbers in parenthesis are parameter standard errors. The last two columns show the probabilities of being wrong in rejecting the indicated hypotheses; they are reported only for those cases in which the estimated parameter is different from zero at a 5% confidence level. A description of the data is given in the Appendix.

Figure 1

Assets as hedges against expected inflation



Notes: Rolling regressions on a 10-year window running from 1979:II-1989:I to 1987:II-1997:I. The diagrams show the parameter in equation (2.2) associated with expected inflation. The proxy used for the latter is the inflation forecast of the Forum-ME survey.

3. Stock returns, inflation and monetary regimes in Italy

The model developed in this section builds on two considerations. First, the framework suggested by Fama and Schwert is not adequate for testing the Fisher effect. Being a restricted version of a reduced-form model, it does not provide any guidance on the selection of the relevant variables and runs the risk of identifying spurious correlations. Second, being dependent on the reaction function of the central bank, equation (2.2) is subject to structural instability. The literature surveyed in Section 1 largely supports these two claims.

The analysis is carried out by splitting the return on a stock or portfolio into two components: the riskless rate, proxied by the interest rate on 3-month Treasury bills, and the excess return. To explain excess returns we rely on the CAPM, while we use the present value relation along the lines suggested by Campbell and Shiller (1988a, b) to detect the channels through which macroeconomic factors affect βs and the market risk premia. This model has the advantage of relying on a sound theoretical basis, because it relates asset prices to their fundamental components. In particular, the decomposition by Campbell and Shiller allows us to express the innovation in the excess return of the stock market as a function of revisions in the expectations on the future values of dividends, excess returns, real interest rates and inflation.

In present value models a crucial role is played by assumptions about the way in which market participants forecast these fundamental variables. We assume that market participants approximate the evolution through time of the relevant state variables by means of a VAR process.

The effects of policy actions are accounted for by allowing the response of financial markets to news to depend on their perception of how the central bank responds to shocks to the economy. Unlike most studies, we do not explicitly define the monetary policy regimes themselves, but rather we try to infer them from market behaviour assuming that regime shifts are governed by an unobserved Markov process. That is, unlike Kaul (1990), we do not explicitly divide the sample period according to the monetary regimes, but rather model the latter as a latent variable in a Markov-switching model, thus allowing the data to speak for themselves. As long as we are able to approximate the way in which financial markets process information, we should succeed in providing a reasonable account of market expectations about policy actions. This is a standing feature of the paper for at least two reasons: first, it enables us to avoid an arbitrary splitting of the sample period; and second, since it does not require us to cluster the observations according to some pre-specified criterion, it does not confine attention to monetary policy but encompasses more general issues, such as credibility, changes in operating procedures and shifts in stance.

After having developed the VAR model with Markov-switching, we estimate the CAPM relation for five portfolios of Italian industry (manufacturing, services, banks, finance and insurance).[11] We then divide the β of each industry portfolio into the components related to the different state variables, following the methodology presented by Campbell and Mei (1993).

In this framework, risky assets provide a complete hedge against expected inflation if the following three conditions are satisfied: the nominal returns on short-term riskless rates move one-to-one with expected inflation; the β of a stock is not affected by anticipated changes in the price index; and the expected component of the excess return on the market portfolio is not correlated with expected inflation. These conditions also allow a test of the Fisher hypothesis, provided that it holds for the riskless asset.

The empirical framework can also deal with a more general assumption, i.e. that the Fisher hypothesis need not necessarily hold for the riskless asset. As will be shown in Section 3.2, the effect of expected inflation on nominal returns is estimated for every asset and the degree of coverage provided by stock returns could turn out to be different from that achieved on the short-term asset.

3.1 The Campbell and Shiller decomposition and the Markov-switching VAR

The model uses a log-linear approximation of the present value relation proposed by Campbell and Shiller. The basic equation links the unexpected stock excess return to changes in the rational expectation of future dividend growth, real interest rates, inflation and future excess returns. If e_{t+1} is the excess return on a stock held from the end of period t to the end of period t+1, d_{t+1} the log real dividend paid during period t+1, r_{t+1} the short-term riskless real interest rate and π_{t+1} the inflation rate, then the equation is:

$$e_{t+1} - E_t e_{t+1} = (E_{t+1} - E_t)\left\{ \sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - \sum_{j=0}^{\infty} \rho^j r_{t+1+j} - \sum_{j=0}^{\infty} \rho^j \pi_{t+1+j} - \sum_{j=1}^{\infty} \rho^j e_{t+1+j} \right\} \tag{3.1}$$

which can be also written in a more compact form as:

$$\tilde{e}_{i,t+1} = \tilde{e}_{di,t+1} - \tilde{e}_{r,t+1} - \tilde{e}_{\pi,t+1} - \tilde{e}_{ei,t+1} \tag{3.2}$$

Once the above asset return components have been computed, it is straightforward to derive the βs between innovations in stock excess returns and the state variables. This means that the

[11] Panetta and Zautzik (1990) show that the CAPM fits Italian stock market data quite well and that there is not much to gain in using a multi-factor model to explain excess returns on risky assets.

latter are used as factors, as in Chen, Ross and Roll (1976) and Fearson (1990). From (3.2), it follows that:

$$\beta_{i,m} \equiv \frac{Cov(\tilde{e}_{di,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{r,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{\pi i,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{ei,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} \equiv \beta_{di,m} - \beta_{r,m} - \beta_{\pi,m} - \beta_{ei,m} \tag{3.3}$$

If one is willing to assume that expectations of future returns are accurately described by the CAPM, then it is possible to substitute out the last term in (3.3). The decomposition of the overall β therefore becomes:

$$\beta_{i,m} = \frac{\beta_{di,m} - \beta_{r,m} - \beta_{\pi,m}}{1+\beta_{em,m}} \tag{3.4}$$

To become operational, the above formulae require a number of hypotheses about the mechanism driving expectations formation. The solution adopted by Campbell and Shiller is to assume that forecasts of excess returns can be approximated by a linear combination of a vector of state variables, x_t,[12] and that the law of motion of these variables can be adequately described by a VAR process.

We have made the further assumption that VAR coefficients are not constant throughout the sample period but rather are subject to occasional discrete shifts; the probability law governing these shifts is represented by a two-state Markov chain. In accordance with the literature surveyed above, we assume that only two regimes are allowed.

The state-space representation of the Markov-switching VAR is the following:

$$x_t = \Pi_{s_t} x_{t-1} + \tilde{x}_t \tag{3.5}$$

$$\xi_t = F\xi_{t-1} + \eta_t \tag{3.6}$$

where s_t is an unobserved random variable that takes the values 1 or 2 depending on in which regime the process is at time t; ξ_t is a two-element random vector, equal to $[1,0]$ if $s_t = 1$ and $[0,1]$ otherwise; $F \equiv \{p_{ij}\}_{i,j=1,2}$ is the transition matrix and p_{ij} is the probability that $s_t = j$ given that $s_{t-1} = i$. The assumption of a first-order centred VAR is not at all restrictive, since (3.5) has to be interpreted as the companion form representation of the process.[13]

[12] The VAR approximation of the mechanism of expectations formation faces at least two problems: first, expectations concern variables which are realised only over long periods of time; second, investors may have information which is not available to the econometrician or cannot be summarised by means of aggregated variables. The first problem can be handled by using the VAR expressions for multi-period forecasts, while the second does create difficulties. The only case where investors' superior information does not distort the analysis occurs when only one component of an asset price is time varying, since then the asset price itself contains all the relevant information about that component. In the general case, the VAR results must be interpreted cautiously, conditional on whatever information is included in the system.

[13] A tricky issue concerned with the analysis of the βs is the proper evaluation of the precision of the estimate. The approach suggested by Campbell is to treat the VAR coefficients and the elements of the covariance matrix of the residuals as parameters to be jointly estimated by GMM. The GMM parameter estimates are numerically identical to OLS ones, but GMM delivers a heteroskedastic-consistent covariance matrix V for the entire set of coefficients. Since the βs can be thought of as non-linear functions $f(\gamma)$ of the vector γ of parameters of the model, their variances turn out to be a quadratic form in the first derivative of $f(\gamma)$ and V. In the present set-up, this procedure is clearly unfeasible: the vector of

The use of a Markov process turns out to be useful on several grounds:

- As stressed forcefully by Sims (1982) and Cooley et al. (1984), it is at least doubtful whether changes in the policy framework should be characterised as permanent changes in the parameters of a reaction function, since genuine changes in regime are rare events. From past experience, economic agents know the menu of choices available to the policymakers and form expectations accordingly, taking into account all the possible outcomes. In other words, they have a probability distribution ranging over all possible policy rules and use it to forecast the behaviour of policymakers.
- A Markov-switching model is flexible enough to encompass once-and-for-all structural changes as well as period-by-period time-varying models. The first case corresponds to each state being a so-called absorbing state, which lasts forever once reached; the second can be approximated by assuming that there exist a large number of states. Any intermediate case can be obtained by appropriately choosing the parameters of the transition matrix.
- Finally, relying on a statistical procedure to split the sample period avoids arbitrary and unnecessarily restrictive assumptions. Monetary policy, which in the literature is usually considered responsible for regime shifts, may not be the only source of instability. Fiscal as well as incomes policies may play a similar role, not to mention the effects deriving from changes in the institutional framework within which economic agents operate. Focusing attention on only one source of instability may be unduly restrictive and could strongly bias the results. Using a statistical technique such as a Markov-switching model has the advantage of allowing the data to speak for themselves; furthermore, the interpretation of the odds attributed to a given regime in each time period provides a genuine test of the reliability of the method.

3.2 The results

The simple tests presented in Section 2 do not support the hypothesis that nominal yields on short-term government securities fully incorporate agents' inflation forecasts (Tables 1a and 1b); at a 5% confidence level, the Fisher hypothesis (together with market efficiency and the null of no correlation between expected inflation and the real rate) is almost always rejected.[14]

However, short-term assets provide partial insurance against expected inflation: the estimated effect on monthly data ranges between 0.41 and 0.64 and does not change significantly either with the maturity of the short-term asset or with the frequency of the data. However, these values are not stable throughout the estimation period.

The splitting of the sample period provided by the Markov-switching algorithm is shown in Figure 2. It is apparent from the graph that the second regime becomes the dominant one in the last quarter of 1988, after a two-year transition period. The interpretation of the change in regime can be clearly related to policy actions and changes in the institutional framework:

- after the realignment of the lira in January 1987, the exchange rate commitment became more credible and no other changes in the central parity of the Italian currency took place until the

---

estimated coefficients has more than 150 elements and the matrix of first derivatives has more than 20,000, not to mention the fact that it is not at all easy to find and differentiate the function relating the βs to the VAR parameters. The solution adopted in this paper is to consider the problem as a special case of the general issue of efficiently and consistently estimating second moments in a model with generated regressors (McKenzie and McAleer (1990) and Pagan (1984)). It is well known that the application of OLS to models with generated regressors will generally be inefficient and lead to inconsistent estimates of the standard errors of the regressor coefficients. A convenient way out of this problem, which has been adopted in this paper, is to allow for non-spherical errors and to use a GLS-type estimator; a simpler alternative is to compute the t-statistics by using a consistent estimate of the covariance matrix of the error term.

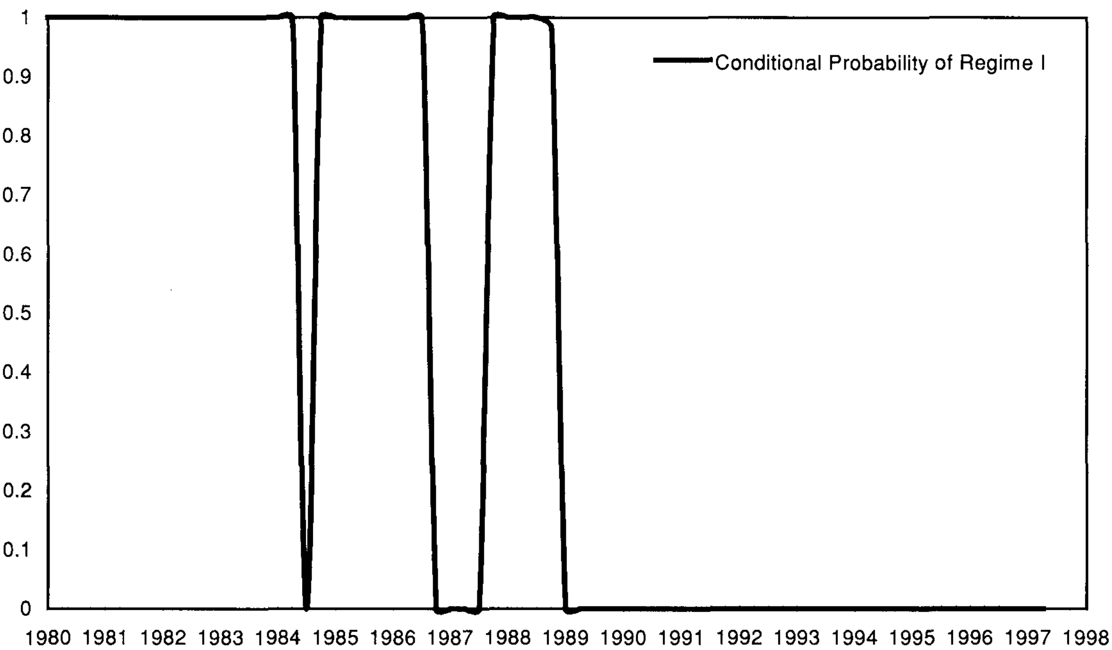
14 This result is robust to different measures of expected inflation and holds for both monthly and quarterly data.

exit of the lira from the Exchange Rate Mechanism of the EMS in September 1992. Between 1987 and 1990, capital movements were progressively liberalised to comply with the requirements set by the EC for the Single Market. The most important measure became effective on October 1988, when all capital movements, except those involving monetary instruments, were liberalised (see Passacantando (1996)). In January 1990, the fluctuation band was narrowed from 6 to 2.25% and the remaining capital controls were completely abolished by April of the same year;

- in May 1988, a screen-based secondary market for government securities was introduced. Between July of that year and March 1989, the floor price for Treasury bills in the primary market was abolished for all maturities. In February 1990, a screen-based market for interbank deposits was launched. In October, banks were allowed to mobilise part of their compulsory reserves. All of these reforms contributed to shifting the conduct of monetary policy from administrative controls to market-oriented procedures.
- incomes policy can also be a factor in a regime change. In the first half of 1984, wage increases were agreed on the basis of a planned rate of inflation rather than relying on a backward-looking indexation mechanism. In 1986, the wage indexation mechanism was further modified by reducing the overall degree of coverage and lowering the frequency of the adjustment (from 3 to 6 months).

Figure 2

Conditional probability of the Italian economy being in regime I



The smoothed probabilities associated with the two regimes also indicate that a reversal of the first occurred in the period 1987:III-1988:III. This shift coincides to a large extent with the reintroduction of controls on bank lending (from September 1987 to March 1988).

The estimates of the VAR model with Markov-switching are reported in Table 2. The effects of past values of inflation on the stock excess return provide a measure of the relationship between expected inflation and the premium requested on stocks. The coefficients indicate that past inflation does not contribute to explaining movements in the overall risk premium; neither in the first nor in the second regime does lagged inflation seem to affect the current excess return on the market portfolio.

Table 2

Double regime Markov-switching VAR model

Equation for market excess returns

	Regime I	Regime II
constant	-0.3599	-0.0836
	(0.1026)	(0.0627)
e_{t-1}	-0.0982	0.2178
	(0.1121)	(0.2123)
r_{t-1}	-9.5344	-1.2027
	(3.8742)	(6.2208)
dy_{t-1}	0.3949	27.1170
	(48.9280)	(55.8980)
π_{t-1}	2.7243	1.4835
	(1.9078)	(2.4867)
Δip_{t-1}	-2.0274	-1.1850
	(0.5814)	(0.3925)
e_{t-2}	0.3669	0.2252
	(0.1142)	(0.1264)
r_{t-2}	6.5005	7.2364
	(3.4563)	(6.5171)
dy_{t-2}	66.7410	-20.8520
	(16.635)	(35.8470)
π_{t-2}	-0.7223	-4.7145
	(10.6239)	(2.2791)
Δip_{t-2}	-0.3231	0.6242
	(0.8121)	(0.4036)

Equation for the short-term real riskless rate

	Regime I	Regime II
constant	0.01246	0.0032
	(0.0028)	(0.0022)
e_{t-1}	-0.0003	-0.0068
	(0.0028)	(0.0063)
r_{t-1}	0.2084	0.7772
	(0.1670)	(0.1640)
dy_{t-1}	-0.7231	0.0560
	(0.6479)	(1.2973)
π_{t-1}	-0.0047	-0.0484
	(0.0479)	(0.0687)
Δip_{t-1}	-0.0032	0.0183
	(0.0221)	(0.0161)
e_{t-2}	-0.0017	-0.0022
	(0.0025)	(0.0043)
r_{t-2}	0.3572	-0.2744
	(0.1024)	(0.1528)
dy_{t-2}	-0.2144	0.3691
	(0.6021)	(1.1270)
π_{t-2}	-0.0536	0.1040
	(0.0412)	(0.0722)
Δip_{t-2}	0.0533	-0.0130
	(0.0208)	(0.0122)

Equation for the dividend yield

	Regime I	Regime II
constant	0.0023	0.0007
	(0.0007)	(0.0004)
e_{t-1}	-0.0034	-0.0036
	(0.0014)	(0.0017)
r_{t-1}	-0.0099	-0.0281
	(0.0545)	(0.0423)
dy_{t-1}	0.7999	0.7405
	(0.2454)	(0.2965)
π_{t-1}	-0.0362	0.0090
	(0.0242)	(0.0129)
Δip_{t-1}	0.0152	0.0052
	(0.0059)	(0.0036)
e_{t-2}	-0.0004	-0.0011
	(0.0010)	(0.0012)
r_{t-2}	0.0060	-0.0156
	(0.0409)	(0.0325)
dy_{t-2}	-0.1163	0.1116
	(0.2649)	(0.2559)
π_{t-2}	0.0189	0.0474
	(0.0143)	(0.0126)
Δip_{t-2}	-0.0072	-0.0045
	(0.0056)	(0.0027)

Equation for the index of industrial production

	Regime I	Regime II
Constant	0.0401	0.0620
	(0.0271)	(0.0278)
e_{t-1}	-0.0341	0.0943
	(0.0292)	(0.0704)
r_{t-1}	-0.6866	3.9414
	(0.9705)	(1.8218)
dy_{t-1}	-4.5008	-18.7210
	(3.8494)	(12.6160)
π_{t-1}	0.9620	-0.9156
	(0.3391)	(0.6646)
Δip_{t-1}	0.0959	-0.2467
	(0.1354)	(0.1561)
e_{t-2}	0.0508	-0.0534
	(0.0262)	(0.0520)
r_{t-2}	0.1406	-6.0491
	(0.5458)	(1.8904)
dy_{t-2}	2.9878	13.8580
	(3.1308)	(12.0130)
π_{t-2}	-1.7657	1.0131
	(0.2922)	(0.6461)
Δip_{t-2}	0.0745	-0.2363
	(0.1517)	(0.1710)

Equation for the rate of inflation

	Regime I	Regime II
constant	0.0068	0.0081
	(0.0065)	(0.0032)
e_{t-1}	0.0025	-0.0047
	(0.0108)	(0.0105)
r_{t-1}	0.2634	-0.5306
	(0.359)	(.3026)
dy_{t-1}	4.6364	1.6690
	(1.8998)	(2.0778)
π_{t-1}	0.3537	0.1591
	(0.1944)	(0.0971)
Δip_{t-1}	-0.2553	0.0539
	(0.0378)	(0.0215)
e_{t-2}	0.0061	-0.0118
	(0.0104)	(0.0057)
r_{t-2}	-0.7168	0.0644
	(0.3204)	(0.2291)
dy_{t-2}	-3.8665	-0.4569
	(1.8501)	(1.9652)
π_{t-2}	0.3333	0.0482
	(0.1615)	(0.1012)
Δip_{t-2}	-0.1055	0.0425
	(0.0751)	(0.02164)

Notes: The VAR model is estimated on quarterly data, for the period 1979:4-1997:1. Numbers in parentheses are coefficient standard errors, calculated according to the formulas suggested in Hamilton (1996). The variables are defined as follows: e_t is the excess return on the market portfolio, r_t the riskless short-term rate, dy_t the dividend yield, π_t is the rate of inflation, and Δip_t is the first difference of the logarithm of the index of industrial production.

Table 3

Italian stock exchange sub-indexes:
Campbell and Shiller's decomposition of the βs with respect to the market portfolio

	Manufacturing		**Services**		**Credit**		**Finance**		**Insurance**	
	Regime I	Regime II	Regime I	Regime II	Regime I	Regime II	Regime I	Regime II	Regime I	Regime II
β dividends	1.6392	2.2592	1.6215	2.0408	1.6808	1.9176	1.8999	2.1656	1.9771	1.7804
	(15.859)	(48.784)	(19.127)	(20.718)	(7.877)	(16.386)	(20.795)	(42.875)	(18.247)	(20.464)
β real rate	0.89145	1.0185	0.89059	1.0185	0.88596	1.0184	0.88437	1.0183	0.87843	1.0180
	(27.460)	(40.603)	(27.341)	(40.612)	(26.718)	(40.494)	(26.512)	(40.241)	(25.765)	(39.284)
β inflation	-0.00774	-0.01091	-0.00776	-0.01091	-0.00783	-0.01093	-0.00786	-0.01097	-0.00795	-0.01112
	(-5.426)	(-18.233)	(-5.426)	(-18.232)	(-5.427)	(-18.240)	(-5.427)	(-18.258)	(-5.429)	(-18.314)
β future excess returns	-0.11964	0.06833	-0.17944	0.04022	-0.09882	0.00822	-0.13035	0.09012	-0.08024	-0.02533
	(-4.455)	(3.361)	(-6.580)	(-0.983)	(-2.210)	(-0.464)	(-3.218)	(5.328)	(-2.108)	(-0.593)
β total	0.87516	1.1833	0.91808	1.0735	0.9015	0.9183	1.1538	1.0681	1.1869	0.79889
	(7.675)	(22.629)	(18.481)	(13.055)	(5.407)	(7.819)	(19.350)	(17.819)	(16.374)	(13.769)

Notes: According to the present value relation, the β of a sub-index with respect to the market portfolio can be decomposed as follows (see equation (3.3)):

$\beta_{total} = \beta_{dividend} - \beta_{real\ rate} - \beta_{inflation} - \beta_{future\ excess\ return}$. Numbers in parentheses are t-statistics.

The two VAR models provide two sets of residuals, which have been used to compute two sets of βs for five industry portfolios, one for each regime (see the last row in Table 3); each β has then been identified as the combination of four components: the real rate, inflation, dividends and excess returns (see equation (3.3)).

Two general remarks on the interpretation of the βs are in order. Since the five portfolios sum up to the whole market and the β for the market is one, the overall βs, shown in the last row of the table, increase in some cases and decrease in others when moving from one regime to another. A further warning is due: since the VAR is not identified, innovations in the state variables are not orthogonalised and the *ceteris paribus* clause cannot be applied in interpreting the βs. This means that the residuals of the VAR equations do not identify exogenous, idiosyncratic shocks to the state variables, but rather represent the unexpected components in the state variables with respect to the previous period information set.

The share of the β of a portfolio attributed to news about the future short-term real interest rate measures the main channel whereby monetary policy affects stock prices, while the β related to inflation provides a quantitative assessment of the effect of inflation innovations on stock excess returns: if the latter did not exert a significant influence on stock excess returns, the value of β_π would be very low and barely significant.

In most cases, all the βs show the same sign across portfolios and the same ranking (in absolute value) across regimes. Compared with the results in Campbell and Mei, all βs show the expected sign: positive for cash-flow and real rate and negative for inflation and future excess return.[15] The differences in the βs across regimes are substantial and statistically significant,[16] showing the existence of tight links between policy actions and market behaviour and supporting the sample splitting induced as a Markovian latent variable.

The dividend component is positive and by far much larger than the other components. In Campbell and Mei, by contrast, the cash-flow βs are always smaller than those related to future excess returns. The size of the dividend component may be overstated, because it is computed as a residual; indeed, one might suspect that the harder portfolio returns are to forecast, the more important the dividend component becomes. But this cannot be the whole story for at least two reasons. First, as is observed by Cambell and Mei, there is no incontrovertible evidence that the fit of the regressions for portfolios' excess returns, as measured by the adjusted multiple correlation coefficient, is negatively related to the size of the residual dividend component. Second, even if $\tilde{e}_{di,t+1}$ is large, there is no guarantee whatsoever that β_d is also large, since most of the variation in cash-flows could be idiosyncratic.

It is worth stressing that β_d changes dramatically between the two sub-samples. The significant increase in the second regime may reflect factors peculiar to the Italian market. Until the mid-eighties companies mostly raised funds by borrowing from banks, thanks to a cheap credit; only rarely did bond or equity issues represent a significant source of financing. In the nineties, owing in part to higher real interest rates and banks' restructuring, an increasing number of companies turned to the international capital markets and thus had an incentive to pursue a dividend policy more akin to those in countries with more developed stock markets. As a result, dividends themselves have become a binding constraint for companies, influencing their investment projects. Another event may have strengthened this process. Starting in the late eighties, small and medium-sized firms have been listed on the Milan stock exchange. Because their capacity to borrow in international capital markets is

---

[15] However, only in few cases is the latter positive.

[16] Two different tests have been computed: in the first case, it has been assumed that residuals are homoskedastic while, in the second, time-varying second moments have been allowed.

limited, they have been forced to pay a great deal of attention to dividend policy. At the same time, two other developments may have altered the sensitivity of stock returns to cash-flow news: the introduction of new financial intermediaries, namely mutual funds, and changes in the tax code.

The estimated values for the βs associated with the real interest rate are positive and increase between the first and the second regime;[17] but the magnitude is greater than that computed on US data. The sign of the correlation is as expected for two reasons. First, since stock prices are forward-looking, they can react to information that is used by the central bank for the conduct of monetary policy. Second, changes in asset prices may have a direct impact on aggregate demand, via both consumption and investment expenditure; whenever the central bank is not confident that this appreciation is fully justified by changes in fundamentals, it may choose to intervene to avoid excessive price and output variability. The increase in β_r in the second regime reinforces this interpretation, as the second half of the sample is characterised by a more restrictive monetary policy stance; increased sensitivity of stock returns to real rates in a tighter monetary environment is also one of the main implications of Söderlind's model.

The estimated effects of news about inflation are negative. Although their size is of second-order compared with the cash-flow and real rate components, they are all highly significant, showing that unexpected inflation exerts some influence on the excess return required on stock portfolios. This evidence is consistent with the Mundell-Tobin effect: upward revisions of agents' inflation forecasts result in a rebalancing of portfolios from money to other assets. In moving from the first to the second regime, the variance of inflation innovations decreases (as is to be expected when monetary policy assigns more weight to inflation targets), while β_π increases. This finding can be related to the greater openness of the Italian economy in the nineties, which has increased the costs of inflationary shocks for most of Italy's listed companies.

The β component associated with future excess returns is generally small and in most cases not statistically significant. This contrasts with the evidence presented by Campbell and Mei for US data, in which, on average, this component is the largest. A possible explanation may be the weak persistence of Italian stock returns, which stands in stark contrast to the US data.

These results provide a first clue about the influence of inflation on stock returns. However, the assumption that portfolio sensitivity to systematic risk is constant within each regime and not allowed to respond to changes in inflation may be unwarranted. This may introduce a bias in the measure of the Fisher effect. In order to test time variation in the overall βs, we have replicated the analysis of Ferson and Schadt (1996) by regressing the innovation in each portfolio's excess returns on the innovation on the markets excess return and the cross products of the latter with each element of the Campbell and Shiller decomposition.[18] The results of the estimates tend to reject time variation in the βs, thus providing additional support to the previous findings.

All in all, the evidence supports the claim that in the last twenty years Italian stocks have not provided a better hedge against inflation than government securities, even when the effects of policy actions on market expectations are taken into account.

---

[17] The sensitivity of the market return to inflation and real interest rate news is approximately the same for all portfolios: marginal differences are due to the discount factor, which is related to the yield ratio. As equation (A.11) in Appendix 2 points out, only the parameter ρ is different across portfolios.

[18] All variables, except the innovation on the excess return on the market portfolio, have been lagged once, in order to ensure that they represent commonly available information.

Summary and conclusions

This paper builds on two main ideas.

(i) Testing the Fisher hypothesis by simple projections of nominal returns on expected inflation is misleading, since those regressions are reduced-form models, powerfully affected by changes in both policy actions and in the institutional framework. The sign and size of the parameter associated with expected inflation can take any value, depending on which variables are added among the regressors. Moreover, the menu of omitted variables is endless, since, in principle, any variable appearing in a structural macro model can be relevant to changes in nominal returns. A structural model is therefore the proper framework within which to analyse the correlation between returns and inflation;

(ii) As is clearly pointed out in the literature on inflation and stock returns, monetary policy must be dealt with to provide a proper account of the relevance of the Fisher hypothesis to the stock market. However, given that the potential sources of instability in the relation between asset returns and inflation are not limited to monetary policy but also include fiscal and income policies as well as changes in the institutional environment, imposing the splitting of the sample among different regimes on the grounds of a priori evaluations would not appear to be the safest and most valuable modelling strategy. The alternative proposed in this paper is to model regime shifts as a stochastic latent variable, with non-sample information not used in setting up the model, but rather in interpreting the results. The advantage is that while no information is discarded, results are not biased by untested assumptions and due attention can be paid not just to monetary policy but to other policy factors.

As a first step in the empirical analysis, we run simple tests of the Fisher hypothesis for Italian Treasury bills and bonds on a sample covering the last twenty years. We apply the same test to equities, to check whether the negative relationships between inflation and stock yields found by Fama and Schwert for US data applies to Italian data as well. In line with evidence for other countries, we find that government securities provide only a partial hedge against expected inflation, while the estimated relationship for common stocks proves inconclusive, due to instability in the parameters.

Within the framework of the CAPM and the log-linear present-value model suggested by Campbell and Shiller, we then test for the influence of inflation on the excess returns required by investors in equities, with a separate analysis of portfolio βs and factor risk. To generate innovations in the state variables, we assume that financial markets form expectations about the relevant macroeconomic variables by means of a VAR model and that the parameters of the expectation formation mechanism change across policy regimes. Finally, we compute βs for a number of industry portfolios and use this decomposition to make inferences about policy actions and the Fisher effect.

The main conclusions of this section are the following:

- sample evidence indicates a shift in the policy environment in the second half of the eighties, when the exchange rate commitment became more binding, monetary control was definitively based on market instruments, and incomes policies became stricter;
- the evidence on short-term assets does not support the Fisher hypothesis, but expected inflation is widely incorporated in short-term interest rates;
- there is no evidence of an influence of inflation forecasts on the market excess return, though the evidence is less clear-cut for the second regime;
- inflation does not seem to have a significant influence on stock excess returns and no time variation in the βs induced by movements in inflation was detected.

Once all the channels through which inflation affects stock returns are taken into account, it turns out that in the last two decades stocks have not significantly outperformed government securities as hedges against inflation.

Appendix 1: Data description

Industrial production: the index of industrial production refers to manufacturing, marketable services and energy. It is collected monthly and adjusted for the number of working days; seasonal adjustment is by an X11-Arima filter.

Inflation: log of the first difference of the cost of living index, net of tobacco products. The index uses a basket of 290 items, which refer to more than five hundred goods and services; data are collected monthly in the capitals of the twenty Italian regions.

Forum-ME survey of inflation expectations: since 1952, the Italian magazine *Mondo Economico* (ME) has conducted surveys on expectations of inflation. The respondents are selected within four main categories of economic agents: managers and executives in industrial, financial and commercial sectors, and business economists. The surveys are conducted by means of an anonymous mail questionnaire and answers have to fall into one of a number of pre-selected intervals (the lower and upper ones being of course open intervals). Until 1981, the survey was run twice a year and covered a six-month forecasting horizon; thereafter, it has been quarterly, with a corresponding shortening of the time frame.

Treasury bills: allotment rate at end-of-month auctions gross of withholding tax. Until June 1981, the Bank of Italy was committed to act as residual buyer for unsubscribed bills. Competitive-bid auctions replaced uniform price auctions in May 1983 for 3-month bills, in May 1984 for 6-month bills and in February 1988 for 12-month bills; for competitive-bid auctions, the yield is the weighted average allotment rate. A floor price for each auction was fixed by the Treasury until June 1988 for 3-month bills and February 1989 for 6 and 12-month bills.

Treasury bonds: average yield of the BTPs with at least one year to maturity traded on the Italian stock exchange, gross of withholding tax.

Dividend yield: total dividends paid over the previous year relative to the current stock price; the latter is computed on the basis of end-of-month closing prices. Data refer to shares of Italian companies listed on the Italian stock exchange.

Stock returns: holding period returns computed on the basis of value-weighted portfolio indexes; Italian listed companies.

Appendix 2: An approximate present-value model with a stochastic discount factor

The model suggested by Campbell and Shiller is a modified version of the present value equation in real terms, relating unexpected returns to changing expectations of future cash flows, real interest rates and excess return.[19] Since the model is derived from a dynamic accounting identity, it is not conditional on any particular asset pricing model; but if one is willing to impose a theoretical structure, it is possible to cancel future required returns and to relate unexpected excess returns to future cash flows and real interest rates only.

The model is derived from the Gordon present value relation, by disposing of the assumption concerning the constancy of the discount factor. Though the relaxation of this hypothesis improves the accuracy of the model, it creates problems of its own, since time-variability of stock returns introduces non-linearities. To overcome this, Campbell and Shiller propose taking logs and linearising the present value relation. The approximate equation is then solved forward, imposing a "no rational bubble" terminal condition.

Starting from the definition of gross stock returns and taking logs, we have:

$$\begin{aligned}\log(1+H_{t+1}) &\equiv \log(P_{t+1}+D_{t+1})-\log(P_t)=\log(P_{t+1})+\log\left(1+\frac{D_{t+1}}{P_{t+1}}\right)-\log(P_t)\\ &= p_{t+1}-p_t+\log(1+\exp(d_{t+1}-p_{t+1})) \qquad \text{(A.1)}\end{aligned}$$

where H_{t+1}, D_{t+1} and P_{t+1} are, respectively, the real return, the dividend and the price of the stock or portfolio we are considering (by the standard convention, logs of variables are denoted with lower-case letters). The last term on the right-hand side is a nonlinear function of the log dividend-price ratio, which can be approximated around the mean using a first-order Taylor expansion:

$$\begin{aligned}\log(1+\exp(d_{t+1}-p_{t+1})) &\approx \log(1+\exp(\bar{d}-\bar{p}))+\frac{\exp(\bar{d}-\bar{p})}{1+\exp(\bar{d}-\bar{p})}\left[(d_{t+1}-p_{t+1})-(\bar{d}-\bar{p})\right]\\ &= -\log(\rho)-(1-\rho)\log\left(\frac{1}{\rho}-1\right)+\rho p_{t+1}+(1-\rho)d_{t+1}-p_t \qquad \text{(A.2)}\end{aligned}$$

where $\rho\equiv\frac{1}{1+\exp(\bar{d}-\bar{p})}\equiv\frac{\bar{P}}{\bar{P}+\bar{D}}$ (the bar indicates sample means). ρ is a number close to 1 and plays the role of a weighting factor. The reason is intuitive: the dividend is much smaller than the stock price, so a given percentage change in the dividend component must have a much smaller effect than the same variation in the price. Substituting (A.2) into (A.1) and solving forward yields:

$$p_t=-\frac{\log(\rho)}{1-\rho}-\log\left(\frac{1}{\rho}-1\right)+\sum_{j=0}^{\infty}\rho^j\left[(1-\rho)d_{t+1+j}-h_{t+1+j}\right] \qquad \text{(A.3)}$$

where the definition $\log(1+H_{t+1})\equiv h_{t+1}$ has been used. This equation is to be interpreted as a dynamic accounting relation, obtained by approximating an identity; it holds *ex post* but also *ex ante*, once future realisations of dividends and returns are replaced by their expected values:

---

[19] A thorough treatment of the present-value relation can be found in Chapter 7 of Campbell et al. (1997).

$$p_t = -\frac{\log(\rho)}{1-\rho} - \log\left(\frac{1}{\rho} - 1\right) + E_t \sum_{j=0}^{\infty} \rho^j \left[(1-\rho)d_{t+1+j} - h_{t+1+j}\right] \tag{A.4}$$

Rearranging (A.1) so that the rate of return is the left-hand variable and substituting (A.4) for both p_t and p_{t+1}, we can write asset returns as linear combinations of revisions in expectations:

$$h_{t+1} - E_t h_{t+1} = (E_{t+1} - E_t)\left\{\sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - \sum_{j=1}^{\infty} \rho^j h_{t+1+j}\right\} \tag{A.5}$$

This equation links the unexpected real stock return in period $t+1$ to changes in the rational expectation of future dividend growth and future stock returns. Equation (A.5) must be interpreted as a consistency condition for expectations; it states that if the unexpected stock return is negative, then either expected future dividend growth must be lower or expected future stock returns must be higher, or both. The discount factor ρ indicates that the further in the future the expectation of a change in returns is, the smaller is the change in today's stock price.

For many purposes it is convenient to work with excess stock returns. If the log real interest rate on a riskless short-term security is r_{t+1}, then the excess return is just $e_{t+1} \equiv h_{t+1} - r_{t+1}$. Substituting this expression into (A.5) provides the following consistency condition:

$$e_{t+1} - E_t e_{t+1} = (E_{t+1} - E_t)\sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - (E_{t+1} - E_t)\sum_{j=0}^{\infty} \rho^j r_{t+1+j} - (E_{t+1} - E_t)\sum_{j=1}^{\infty} \rho^j e_{t+1+j} \tag{A.6}$$

In this paper, we have used a slightly modified version of this equation, obtained by taking a present value relation expressed in nominal rather than real terms as a starting point. In this case, (A.6) becomes:

$$e_{t+1} - E_t e_{t+1} = (E_{t+1} - E_t)\left\{\sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - \sum_{j=0}^{\infty} \rho^j r_{t+1+j} - \sum_{j=0}^{\infty} \rho^j \pi_{t+1+j} - \sum_{j=1}^{\infty} \rho^j e_{t+1+j}\right\} \tag{A.7}$$

which can also be written in a more compact form as:

$$\tilde{e}_{i,t+1} = \tilde{e}_{di,t+1} - \tilde{e}_{r,t+1} - \tilde{e}_{\pi,t+1} - \tilde{e}_{ei,t+1} \tag{A.8}$$

(The meaning of these terms is evident, by comparing (A.8) with (A.7).)

The excess return on a portfolio is assumed to be predictable by means of a projection on a vector of state variables x_t:

$$e_{i,t+1} = a_i^{'} x_t + \tilde{e}_{t+1} \tag{A.9}$$

where $a_i^{'}$ is a vector of projection coefficients and $\tilde{e}_{t+1}$ is the unexpected component of the excess return.

To become operational, the above formulas require some hypotheses concerning the mechanism that drives expectation formation. The solution adopted by Campbell and Shiller is to assume that the law of motion of the state variables can be adequately described by a VAR process:

$$x_{t+1} = \Pi x_t + \tilde{x}_{t+1} \tag{A.10}$$

where $\tilde{x}_{t+1}$ is the innovation in the state vector. To allow for higher order processes or deterministic components, one must suitably augment the dimension of the vector of state variables. The first three elements of x_t are the excess return on the market, the real return on a short-term Treasury bill and the rate of inflation; the other components are selected from variables that are known to the market by time t and that have been shown in the literature to have some explanatory power for future returns:[20] for example, the dividend yield, the slope of the term structure and the default spread. Given the VAR model, revisions in rational expectations of the state variables are provided by the expression:

$$(E_{t+1} - E_t)x_{t+1+j} = \Pi^j \tilde{x}_{t+1} \tag{A.11}$$

Equation (A.11) enables us to compute the right-hand terms in (A.6) and (A.7). If i_j indicates the vector that picks the j-th component of $\tilde{x}_{t+1}$, the following equations hold:

$$\tilde{e}_{em,t+1} = \rho a_1^{'}(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{ei,t+1} = \rho a_i^{'}(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{r,t+1} = i_2^{'}(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{\pi,t+1} = i_3^{'}(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{di,t+1} = \tilde{e}_{i,t+1} + (i_2^{'} + \rho a_i^{'})(I - \rho\Pi)^{-1}\tilde{x}_{t+1} \tag{A.12}$$

The component associated with innovation in the path of dividend growth is computed as a residual and is therefore likely to be overstated. However, the sign of the bias is uncertain, since it will depend on the covariances between omitted and included variables.

Once the above asset return components have been computed, it is straightforward to derive the βs between innovations in stock excess returns and in the state variables. This means that the latter are used as factors, as in Chen, Ross and Roll (1976) and Fearson (1990). From (A.7), it follows that:

$$\beta_{i,m} \equiv \frac{Cov(\tilde{e}_{di,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{r,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{\pi i,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{ei,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} \equiv \beta_{di,m} - \beta_{r,m} - \beta_{\pi,m} - \beta_{ei,m} \tag{A.13}$$

If one is willing to assume that expectations of future returns are well described by a simple CAPM, then the last term in (A.13) can be substituted out. The decomposition of the overall β thus becomes:

$$\beta_{i,m} = \frac{\beta_{di,m} - \beta_{r,m} - \beta_{\pi,m}}{1 + \beta_{em,m}} \tag{A.14}$$

---

[20] See for instance Fama (1988), Fama and French (1988) and Boldrin et al. (1995).

References

Balduzzi, P. (1993): "Stock returns, inflation, and the 'proxy hypothesis': a new look at the data". New York University, *mimeo.*

Boldrin, M., L. J. Christiano and J. D. M. Fisher (1995): "Asset pricing lessons for modelling business cycles", *mimeo.*

Bomberger, W. A. and W. J. Frazer Jr. (1981): "Interest rates, uncertainty and the Livingston data". *The Journal of Finance*, Vol. 36, No. 3, pp. 661-75.

Campbell, J. and J. Ammer (1993): "What moves the stock and bond markets? A variance decomposition for long-term asset returns". *The Journal of Finance*, Vol. 48, pp. 3-37.

Campbell, J., A. W. Lo and A. C. MacKinlay (1997): *The econometrics of financial markets*, Princeton University Press, Princeton.

Campbell, J. and J. Mei (1993): "Where do betas come from? Asset price dynamics and the sources of systematic risk". *NBER Working Paper*, No. 4329.

Campbell, J. and R. Shiller (1988a): "The dividend-price ratio and expectations of future dividends and discount factors". *Review of Financial Studies*, Vol. 1, pp. 195-227.

Campbell, J. and R. Shiller (1988b): "Stock prices, earnings and expected dividends". *The Journal of Finance*, Vol. 43, pp. 661-76.

Chen, N., R. Roll and S. Ross (1986): "Economic forces and the stock market". *Journal of Business*, Vol. 59, pp. 386-403.

Cooley, T., S. F. LeRoy and N. Raymon (1984): "Econometric policy evaluation: note". *American Economic Review*, Vol. 74, No. 3, pp. 467-70.

Cooley, T. (1989): "Using production based asset pricing to explain the behavior of stock returns over the business cycle". *NBER Working Paper*, No. 3212.

Evans, M. D. D. and K. Lewis (1995): "Do expected shifts in inflation affect estimates of the long-run Fisher relation?" *The Journal of Finance*, Vol. 50, No. 1, pp. 225-53.

Fama, E. F. (1981): "Stock returns, real activity, inflation and money". *American Economic Review*, Vol. 71, No. 4, pp. 545-65.

Fama, E. F. and G. W. Schwert (1977): "Asset returns and inflation". *Journal of Financial Economics*, No. 5, pp. 115-46.

Fama, E. F. and K. R. French (1988): "Business cycles and the behavior of metals prices". *The Journal of Finance*, Vol. 43, No. 5, pp. 1075-93.

Fama, E. F. (1988): "Dividends yields and expected stock returns". *Journal of Financial Economics*, No. 22, pp. 3-25.

Ferson, W. E. (1990): "Are the latent variables in time-varying expected returns compensation for consumption risk?" *The Journal of Finance*, Vol. 45, No. 2, pp. 397-429.

Ferson, W. E. and R. W. Schadt (1996): "Measuring fund strategy and performance in changing economic conditions". *The Journal of Finance*, Vol. 51, No. 2, pp. 425-61.

Fuhrer, J. and G. Moore (1995): "Monetary policy trade-offs and the correlation between nominal interest rates and real output". *American Economic Review*, Vol. 85, pp. 219-39.

Geske, R. and R. Roll (1983): "The fiscal and monetary linkage between stock returns and inflation". *The Journal of Finance*, Vol. 38, No. 1, pp. 1-33.

Grande, G. (1994): "La curva dei rendimenti dei BOT come misura dei tassi futuri attesi". Banca d'Italia, *Temi di discussione*, No. 232.

Groenewold, N., G. O'Rourke and S. Thomas (1997): "Stock returns and inflation: a macro analysis". *Applied Financial Economics*, Vol. 7, pp. 127-36.

Hamilton, J. D. (1994): *Time series analysis*, Princeton University Press, N. J.

Hamilton, J. D. (1996): "Specification testing in Markov-switching time-series models". *Journal of Econometrics*, No. 70, pp. 127-57.

Kaul, G. (1987): "Stock returns and inflation: the role of the monetary sector". *Journal of Financial Economics*, Vol. 18, pp. 253-76.

Kaul, G. (1990): "Monetary regimes and the relation between stock returns and inflationary expectations". *Journal of Financial and Quantitative Analysis*, Vol. 25, No. 3, pp. 307-21.

Kneeshaw, J. T. (1995): "Non-financial sector balance sheets in the monetary policy transmission mechanism". BIS, *Financial structure and the monetary policy transmission mechanism*, Basle.

McKenzie, C. R. and M. McAleer (1990): "On efficient estimation and correct inference in models with generated regressors: a general approach". *Australian National University Working Papers*, No. 211.

Mishkin, F. S. (1990): "The information in the longer maturity term structure about future inflation". National Bureau of Economic Research, *mimeo.*

Mitchell, D. W. (1985): "Expected inflation and interest rates in a multi-asset model: a note". *The Journal of Finance*, Vol. 40, No. 2, pp. 595-9.

Pagan, A. (1984): "Econometric issues in the analysis of regressions with generated regressors". *International Economic Review*, Vol. 25, No. 1, pp. 221-47.

Panetta, F. and E. Zautzik (1990): "Evoluzione e performance dei fondi comuni mobiliari italiani". Banca d'Italia, *Temi di discussione*, No. 142.

Passacantando, F. (1995): "Building an institutional framework for monetary stability: the case of Italy (1979-1994)". *BNL Quarterly Review*, No. 196, March.

Rovelli, R. (1984): "Expected inflation and the real interest rate: a survey of current issues", in *Giornale degli Economisti*, pp. 671-97.

Santoni, G. J. and H. B. Moehring (1994): "Asset returns and measured inflation". *Journal of Money, Credit and Banking*, Vol. 26, No. 2, pp. 232-48.

Sims, C. A. (1982): "Policy analysis with econometric models". *Brookings Papers on Economic Activity*, pp. 107-52.

Söderlind, P. (1997): "Monetary policy and the Fisher effects". *CEPR Discussion Paper*, No. 1610.

Stulz, R. M. (1986): "Asset pricing and expected inflation". *The Journal of Finance*, Vol. 41, pp 209-23.

Tobin, J. (1965): "Money and economic growth". *Econometrica*, Vol. 33, pp. 671-84.