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The Stability of the Relation between the Stock Market and Macroeconomic Forces

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THE STABILITY OF THE RELATION BETWEEN THE STOCK MARKET AND MACROECONOMIC FORCES

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Abstract

This paper identifies the macroeconomic factors that influence Italian equity returns and tests the stability of their relation with securities returns. In the sixteen-year period that has been analyzed the relation between stock returns and the macroeconomic factors is found to be highly unstable: not only are the betas of individual securities virtually uncorrelated over time, but a high percentage of the shares experience a reversal of the sign of the estimated loadings. This result is not confined to single periods or to a small group of shares, but holds in different sub-periods and for securities in all risk classes. These findings suggest that empirical analysis of asset pricing should carefully investigate the specification of the return generating process and the stability of the risk measures.

JEL classification: G12, E44.

Keywords: arbitrage pricing theory, return generating process, stock market factors, factor loadings.

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1. Introduction[1]

The relation between the stock market and macroeconomic forces has been widely analyzed in the finance and macroeconomic literature - see, for example, Fama (1981), Friedman (1988), Keran (1971), Nelson (1976). The linkages between equity prices and variables such as money supply, inflation and industrial production are of crucial importance not only in analyzing equity returns, but also in understanding the connections between expected returns and the real economy.

Starting with the work of Chen, Roll and Ross (1986b), the literature on the Arbitrage Pricing Theory (APT) has given new impetus to research on the macroeconomic determinants of equity returns. Research has concentrated mainly on the significance of the risk premia attached to each macroeconomic factor, providing considerable evidence that state variables such as industrial production growth, default risk premia and yield spreads between long and short-term government bonds are important in explaining equilibrium asset prices - see Chan, Chen and Hsieh (1985), Hamao (1988), McElroy and Burmeister, (1988) for the US and Roma and Schlitzer (1996) for the Italian stock market.

By contrast, the previous studies have generally paid little attention to the stability of securities' risk measures. However there is no reason, at least in principle, for which the sensitivities of equity returns to the macroeconomic factors should not change over time. The changes might be due to economy-wide factors (take for example the process of globalization, which in the last decade has reshaped most economies and industries) or to firm-specific factors, such as an increase in leverage or the development of a new line of business. The limited research on the stability of risk measures in a multifactor framework is even more striking if one considers the large number of studies which have investigated stability in a market model framework[2].

Assessing the stability of securities' risk characteristics is of fundamental importance for empirical applications of multifactor pricing models. The list of applications that use macroeconomic state variables includes the analysis of the performance of seasoned equity issues (see Eckbo, Masulis and Norli, 2000), the evaluation of the profitability of insider trades (see Eckbo and Smith, 1998), the evaluation of portfolio performance (see, for example Elton, Gruber and Blake, 1995, for an investigation of the performance of bond mutual funds and Connor and Korajczyk, 1991, for an analysis of the performance of equity funds), the forecasting of the correlation structure of share prices (see, for example, Eun and Resnick, 1992), the estimation of the cost of capital (see, for example, Goldenberg and Robin, 1991).

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2 See Blume (1971, 1975), Sharpe and Cooper (1972), Fabozzi and Francis (1978), Ball and Kothary (1989), Collins et al. (1987), Chan and Chen (1988) for the US market, Dimson and Marsh (1983) for the UK market and Brooks, Faff, and Lee (1992) for the Australian market.

This paper identifies the macroeconomic state variables that influence Italian equity returns and tests the stability of their relation with securities returns. No attempt is made to test pricing relationships. Since different economies are likely to be idiosyncratic to some degree in selecting the relevant macroeconomic factors our attention is not confined to variables used in previous research on other countries, but rather variables which might have a specific relevance for the Italian stock market. In accordance with the rational expectations and market efficiency hypotheses, the innovations in the macroeconomic series are estimated; the variables included in the return generating process are then selected on the basis of their ability to predict the factor scores estimated using Exploratory Factor Analysis (EFA). The stability of the risk measure for individual securities is then analyzed.

The results of the analysis suggest that Italian stock returns are influenced by innovations in the slope of the term structure, by unexpected changes in the rate of inflation, in industrial production and in the price of crude oil imports, and by changes in the lira/US dollar exchange rate. Other variables, such as a revision of the expected rate of inflation, the surprise in the default risk premium or unexpected shifts in the money supply, and the rate of change of *per capita* consumption are not significant in explaining securities returns. The relation between stock returns and the relevant macroeconomic systematic factors is highly unstable: not only are the betas of individual securities virtually uncorrelated over time, but in the sixteen-year period which has been analyzed a large proportion of the shares also experienced a reversal of the sign of the estimated betas. This result is not confined to single periods or to a specific group of shares but is detected in different sub-periods and for securities in all risk classes.

These findings suggest that applications of asset pricing models which employ macroeconomic state variables should be regarded cautiously, and that the empirical analysis on asset pricing should take a step backward and carefully analyze the specification of the return generating process and in particular the stability of securities risk measures.

The paper is organized as follows. Section 2 reviews the previous literature. Section 3 recalls the theoretical framework. Section 4 describes the macroeconomic variables. The selection of the systematic factors is described in Section 5. Section 6 presents the test on the stability of the risk measures for individual securities. The main conclusions are summarized in Section 7. The data sources and the estimation of the factor scores are described in an Appendix.

2. Related research

Several methods have been used in the previous literature to measure the stability of individual securities risk measures. The most common approach is to estimate betas over successive periods and then calculate the correlation coefficient between successive estimates or to construct transition matrices showing how the estimates change over time - see, for example, Blume (1971), Sharpe and Cooper (1972) and Dimson and Marsh (1983). This implicitly assumes that the true underlying betas are constant within each estimation period but vary stochastically between successive non-overlapping estimation periods.

Alternatively, the stability hypothesis has been tested by modeling the beta variation explicitly. Examples of this approach include Fabozzi and Francis (1978), who apply the Hildret and Houck random coefficient model, Sunders (1980), who considers a random walk and an AR(1) model for the beta changes, Collins et al. (1987), who test the ARMA model for beta variations.

While several papers have addressed the problem of the stability of the beta coefficient in a market model framework, little attention has been devoted to analyzing the stability of the risk measures in multifactor models. This lack of attention is striking given that multifactor asset pricing models are widely used in the financial literature and the issue of beta stability is crucial for any economic application of such models.

Among the rare exceptions are the papers of McQueen and Roley (1990), Chang and Pinegar (1990) and Ferson and Harvey (1991). McQueen and Roley (1990) find that the impact of fundamental macroeconomic news on stock returns depends on the state of the economy: in particular, positive news on real activity when the economy is already strong results in lower stock prices, while during a recession the same surprise is associated with higher prices. Using the Chen, Roll and Ross (1986b) macroeconomic state variables, Chang and Pinegar (1990) find that the factor betas differ between January and other months. Ferson and Harvey (1991) use prespecified economic factors similar to Chen, Roll and Ross (1986) to analyze the behavior of the risk premia over time. They decompose the predictable part of portfolio returns to assess the relative importance of time-varying risk and time-varying risk premia, and find that most of the predictable variation in the portfolio returns can be attributed to changes in the market price of beta risk, while time variation of the betas appears small.

In this paper the stability of the relation between Italian equity returns and macroeconomic factors is analyzed using the methodology suggested by Dimson and Marsh (1983). Formal tests of stability of the factor loadings are reported, although no attempt is made to model loading changes.

3. Theoretical framework

The APT, derived by Ross (1976, 1977), assumes that the uncertainty about securities payoffs depends on a small number of common factors, and that the return generating process is such that the difference between realized and expected returns is a linear function of the common factors plus a random error:

$$(1) \qquad \tilde{R}(t) - E_{t-1}\tilde{R}(t) = B\tilde{f}(t) + \tilde{\varepsilon}(t)$$

where $\tilde{R}(t) = \left[\tilde{R}_1(t),\ldots.,\tilde{R}_N(t)\right]'$ is the vector of returns on individual securities, $E_{t-1}\tilde{R}(t) = E_{t-1}\left[\tilde{R}_1(t),\ldots,\tilde{R}_N(t)\right]'$ is the vector of expected returns, B is the NxK matrix of the risk exposure of each of the N assets to each of the K sources of systematic risk (for example, the element b_{ij} is the sensitivity of asset i to factor j), $\tilde{f}(t) = \left[\tilde{f}_1(t),\ldots,\tilde{f}_k(t)\right]'$ is the vector of innovations in the k systematic risk factors and $\tilde{\varepsilon}(t) = \left[\tilde{\varepsilon}_1(t),\ldots,\tilde{\varepsilon}_N(t)\right]'$ is the vector of the N firm-specific risk terms. Moreover, it is assumed that the $\tilde{\varepsilon}(t)$ terms are serially uncorrelated and independent of the factors $\tilde{f}(t)$ with $E[\tilde{\varepsilon}(t)] = 0$ and $E[\tilde{f}(t)] = 0$. By the diversification argument implicit in the APT, only economy-wide factors (i.e. variables which influence a large number of securities) may be systematic factors in the APT.

The APT does not provide any indications about which variables should appear on the right hand side of equation (1), since the systematic factors are not identified and the existence of the linear relation between the factors and securities returns is merely an assumption of the model. A useful benchmark suggested by financial theory is the so-called dividend discount model, which states that securities prices must equal the discounted value of expected dividends, i.e.

$$P_t = \sum_{j=o}^{\infty} E_t \left[\frac{D_{t+j}}{1 + r_{t,t+j}} \right] \tag{2}$$

where D_t is the dividend at date t, $r_{t,t+j}$ is the interest rate between date t and date t+j and E_t denotes expectations at time t. Following equation (2), the macroeconomic factors have been selected from among the variables which influence expected cash-flows or discount rates.

4. The macroeconomic factors

This section describes the state variables that are used in the empirical analysis. No claim is made that all the macroeconomic variables which influence stock returns are included; however, the variables that are analyzed are of some economic interest and many of them have been widely used in the financial literature. In some cases the list of variables has been constrained by data availability. The variables which have been employed are shown in Table 1, while the data sources are reported in the Appendix.

The rational expectations and market efficiency assumptions require the identification of unexpected changes in the series. Particular attention is paid to the timing of the arrival of information: financial variables are generally measured precisely and can be observed in real time, while information on non-financial variables is often released with substantial delay. Although other hypotheses have been examined, this paper assumes that stock prices are influenced by the announcements about the macroeconomic factors, although the information embedded in the announcement might refer to previous periods. This problem is discussed in greater detail below, especially with reference to the industrial production series.

4.1 Industrial production

Since the numerator of equation (2) is related to current and future economic activity, proxies for the innovations in the rate of growth of industrial production are employed, following an approach which has been adopted in previous studies.[3]

In Italy monthly industrial production figures are released by ISTAT (the Italian National Institute of Statistics) with approximately 45 days of delay: for example, the data on the level of industrial production in January is released around mid-March. Given such delays, one has to consider how to model the relation between news on industrial production and equity returns. In this paper the hypothesis is made that stock market returns are influenced by the *announcement* of the most recent figures on industrial production, although they refer to month t-2. This implies, for example, that in March investors formulate their investment decisions on the basis of the news on the January figures for industrial production. Given this hypothesis, stock returns and the industrial production series are made contemporaneous by lagging the latter by two periods.

A simple estimator for the innovation in industrial production that has been used in the previous literature is the rate of growth of the series

$$MP_t = \log IP_{t-2} - \log IP_{t-3}$$

where IP_{t-2} is industrial production in period t-2, announced at t, and MP_t is the logarithmic rate of growth of industrial production during month t-2. However, the monthly growth rates for industrial production are strongly serially correlated (see below), so that the information that they contain cannot be considered unexpected. In this paper we have therefore used an econometric model suggested by Bodo, Signorini and Cividini (1991) to estimate the series of the expected values of IP on the basis of electricity consumption. The following equation was fitted to the monthly industrial production series:

$$IP_t = a_0 + \sum_{i=1}^{11} a_i S_{it} + b_1 t + b_2 t^2 + b_3 t^3 + \sum_{k=1,2,5,7,8} c_k EL_{k,t} + d_1 Temp_t + d_2 Temp_t^2 + e_1 IP_{t-12} + \varepsilon_t \tag{3}$$

where the S_i variables are monthly dummies, t is a time trend, EL_k is electricity consumption in each month in 5 selected areas of Italy[4], $Temp_t$ is the average temperature in

[3] See for example Chen, Roll and Ross (1986a, b), Hamao (1988), Shanken and Weinstein (1987).

[4] Rome, Milan, Turin, Palermo and Cagliari.

each month[5]. Bodo, Cividini and Signorini (1991), Schlitzer (1993) and Marchetti and Parigi (1998) analyze extensively the forecasting properties of equation (3) and compare it with other methods, including ARIMA models, concluding that for monthly forecasts its performance is superior to that of the alternative methods. In addition to its high explanatory power, an attractive feature of equation (3) is that the RHS variables are promptly observable, because the data on electricity consumption in month t are available in the first week of month t+1. After fitting equation (3), the following innovation variable has been calculated:

$$UMP_t = \log IP_{t-2} - E(\log IP_{t-2} | \mathrm{t}-1)$$

where $E(IP_{t-2}|t-1)$ is the expectation of IP_{t-2} conditional on the information available at time t-1 and UMP is the forecast error for the most recent figure on industrial production available to investors in month t (i.e. production in month t-2). The expectation of the production series has been calculated estimating for each month a rolling regression of equation (3), using the last 48 months of data. For example, the forecast error which is relevant for the investment decisions which were taken in March 1985 is estimated by calculating the log of the difference between the level of industrial production in January 1985 (released in March 1985) and the value of industrial production in January 1985 obtained using equation (3). Using UMP implies that what matters for the stock market is the difference between the latest figures on industrial production announced by ISTAT for month t-2 and the expectation of IP_{t-2} formed on the basis of equation (3).

However, since other information may become available in the interim, there may be limited information in the release of the production figures by ISTAT. One might therefore hypothesize that the relevant figures for investors' decisions in month t are those regarding production in the *same month*. Two proxies were therefore calculated for the unexpected growth of industrial production in month t. If agents anticipate future movements of the relevant variables, stock returns should lead changes in the factors; accordingly, the first proxy is the actual change of industrial production in month t (although the information has not been released yet):

$$MP1_t = \log IP_t - \log IP_{t-1}.$$

The second proxy is the revision in the expectation of industrial production in month t, i.e.:

$$UMP1_t = E(\log IP_t | t) - E(\log IP_t | t-1).$$

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5 The trend, the temperature and the desegregation by geographic areas are included in the regression to account for electricity consumption not related to industrial production.

When new information becomes available, investors revise the expectation of industrial production in month t formulated in the previous month. Using UMP1 is equivalent to assuming that returns in month t are influenced by the revision in investors' forecast about industrial production in the same month: for example, this implies that in March investors formulate their investment decisions on the basis of the difference between their expectation for industrial production in March formulated in February and the same expectation revised in March on the basis of a broader set of information. The term $E(logIP_t|t\text{-}1)$ has been estimated using rolling regressions of equation (3) up to month t-2 and then calculating a two-step-ahead forecast with an autoregressive model[6]; similarly, the term $E(logIP_t|t)$ has been calculated running a rolling regression of equation (3) up to month t-1 and then calculating a one-step-ahead forecast.

4.2 Inflation

Changes in the expected rate of inflation influence both the numerator and the denominator of equation (2). The expected value of firms' future cash flows might be influenced by revisions in expected inflation if inflation has real effects - for example, redistributing resources among different sectors of the economy - which are larger when average inflation is higher. In this case a change in expected inflation will have a systematic effect on share prices. The influence on share prices of the surprise in the inflation process is also analyzed: if the effect of inflation is not neutralized by changes in nominal cash flows or in the discount rate, unanticipated inflation will be systematically followed by adjustments in share prices.[7]

The inflation rate is calculated as the monthly logarithmic change in the cost of living index (CLI), whose value for each calendar month is released by ISTAT in the last week of the same month (see the Appendix). The proxy for expected inflation has been estimated by fitting an

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6 The industrial production forecasts were obtained by estimating the following seasonal ARIMA model using maximum likelihood:

$$(1-L)(1-L^{12})IP_t^* = c + (1-\Theta_1 L)(1-\Theta_{12}L^{12})\mu_t ,$$

where IP* is the fitted value of industrial production obtained using equation (3), c is a constant, L is the lag operator and μ_t is a white noise error. UMP1 was also re-estimated by fitting an autoregressive model on the industrial production series directly, rather than on the fitted values obtained with equation (3). This modification did not imply any qualitative difference in the regression analysis reported below and the results are not reported.

7 Schwert (1981) analyzed the impact of inflation surprises, and found that on average stock returns fall in response to unexpected inflation. The negative relation between inflation surprises and stock prices has been attributed to the deterioration of inflationary expectations which is caused by a positive inflation surprise - see also Fama and Schwert (1977). Alternatively, the negative impact of inflation surprises on stock prices could be due to the anticipation of more restrictive monetary policies and higher real rates.

ARIMA model[8] to the CLI over a rolling period and calculating a one-step-ahead forecast for every month using the previous 4 years of data on monthly inflation for each estimate. Subsequently, unanticipated inflation (UI) was calculated as

$$UI_t = I_t - E(I_t | t-1)$$

where I is the monthly logarithmic change in the CLI.

The change in expected inflation was computed as:

$$DEI_t = E(I_{t+1} | t) - E(I_t | t-1).$$

4.3 Interest rates

To capture the risk reflected in unexpected changes in the denominator of equation (2), interest-rate-related variables are also examined. Since stock prices reflect the value of all future cash flows, the discount operator in (2) is influenced by modifications in the term premium and in the risk premium. Therefore, proxies are examined for unexpected shifts in the slope of the term structure and for innovations in the spread between the bank lending rate paid by high and low-quality borrowers, a proxy for the default risk premium.

i) Term structure. The unanticipated change in the slope of the term structure has been proxied with the unexpected return on long bonds, measured by the difference between the holding-period return on a portfolio of long-term government bonds in month t ($LHPR_t$) and the yield on short-term Treasury bills at the end of the previous month (SR_{t-1}):

$$UTS_t = LHPR_t - SR_{t-1}.$$

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8 The ARIMA model used to fit monthly inflation is the following:

$$(1-L)(1-L^{12})I_t = c + (1-\Theta_1 L)(1-\Theta_{12}L^{12})\mu_t$$

where I is the monthly logarithmic change in the CLI, c is a constant, L is the lag operator and μ_t is a white noise error. The estimates were obtained using maximum likelihood.

For a risk neutral investor, the ex-ante value of UTS should be zero in equilibrium. A positive value for UTS implies that the change in the price of long-term bonds has been greater than its "expected value", so that their yield has decreased and the yield curve has become flatter than anticipated.[9] As Chen, Roll and Ross (1986b) suggest, under the assumption of risk neutrality the expected value of (UTS) should equal zero, so that this variable should allow one to isolate the pure term structure effect from the effect of changes in risk aversion.

ii) Risk premia. In Chen, Roll and Ross (1986a, b) the impact of unexpected changes in risk premia on equity returns has been captured using the difference between the return on government bonds and that of low grade bond portfolios. Unfortunately, in Italy there are no data on corporate bonds or on company grading. Data on bank lending rates were therefore employed: the first indicator which has been calculated is the change in the spread between the minimum bank lending rate (MINBR) and the average bank lending rate (AVEBR):

$$UPR_t = PR_t - PR_{t-1}$$

where $PR_t = AVEBR_t \text{ - } \mathrm{MINBR}_t$.

A change in UPR can be interpreted as a shift in the degree of risk aversion, which is implicit in the discount applied to future cash flows. The data refer to sight loans: this choice was made not only because sight loans are by far the largest component of external financing to the corporate sector in Italy, but also because banks can change the interest rate on these loans at any time, ensuring that the rates reflect changes in market conditions.[10]

4.4 International factors

International factors have a strong impact on the competitiveness of Italian exports, and their inclusion proxies for future economic growth opportunities. The effects of surprises in the lira/US dollar exchange rate and in the price of crude oil, which is the most important production input imported by Italian producers, were therefore considered.

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9 A positive value for UTS does not imply that the term structure has actually become flatter, but only that it has become flatter than expected. In fact, a positive UTS might exist, contemporaneously with a fall in SR_t, so that the slope of the yield curve could remain unchanged or even become steeper.

10 The change in the spread between the bank lending rate paid by corporate borrowers in the bottom and top deciles of the distribution of bank lending rates was also analyzed. However, this variable, which is available only on a quarterly basis, displayed serial correlation and provided no additional information.

i) Exchange rates. The proxy which has been used to capture the effect of unexpected changes in exchange rates on stock returns is the rate of change in the lira/US dollar exchange rate (DLUSD):

$$DLUSD = \log SX_t - \log SX_{t-1}.$$

The decision to use the US dollar exchange rate is motivated mainly by the fact that the US dollar was the most important currency (in terms of its relevance for Italian international trade) that was not part of the EMS, so that its fluctuations should reflect market forces more accurately than other exchange rates included in the EMS.[11]

ii) *Oil prices.* For each month the expected value of oil prices (expressed in US dollars) was estimated fitting a MA(1) process over a rolling period of four years and calculating a one-step-ahead forecast for every month. Subsequently, unanticipated changes in the series (UOILG) were calculated:

$$UOILG = \log OILG_t - E(\log OILG_t | t-1) .$$

4.5 Money growth

The impact of the weekly money stock announcements on securities returns has been analyzed by Cornell (1983), Pearce and Roley (1983) and Ulrich and Watchel (1981). The consensus finding is that unexpected money growth is associated with lower stock prices.[12]

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[11] Two alternative variables were considered to capture the effects of exchange rate surprises and were subsequently omitted from the analysis. The first is the difference between the rate of change during month t in the lira/US$ spot exchange rate and the 1-month forward exchange rate recorded in month t-1 for delivery in month t. This variable was omitted because it is strongly serially correlated (and therefore cannot be considered an innovation) and was not significant in explaining the variation in the factor scores estimated using EFA (see below). The second variable considered is the log change in the terms of trade (the ratio between the price index of export goods and the price index of import goods). It was omitted because the delay with which it becomes available to the market changes from month to month, so that it is impossible to compute the unexpected change. However, the stability tests reported below have been replicated substituting the change in the terms of trade for DLUSD, and no qualitative modification of the results was detected (see below).

[12] This finding has been interpreted in two different ways. First, the decline in stock prices could be due to the fact that in response to an unexpected increase in money supply agents anticipate tighter monetary policy and higher interest rates. The second interpretation is based on the possibility that an increase in the money supply causes a deterioration in inflationary expectations, and thus a fall in stock prices.

From the early eighties to the mid-nineties, at the end of each year the Bank of Italy announced the target monthly rates of growth in monetary and financial aggregates for the following year. The announcement included a range with a top and bottom limit for money growth. For example, at the end of 1992 the Bank of Italy announced for 1993 a target rate for M2 growth equal to 5 per cent, with an upper limit of 6 per cent and a floor of 4 per cent.[13] The variable which we have used to proxy unexpected shifts in the money supply, UMS, is the change in the divergence between the actual and planned (announced) rate of growth in financial aggregates:

$$UMS_t = \Delta\left(MG_t - AMG_t\right)$$

where Δ is the difference operator, AMG_t is the monthly projected rate of growth over the following year in monetary aggregates that is announced each year and MG_t is the actual rate of growth in monetary aggregates in month t. Until December 1984 the growth of credit aggregates (total lending to the private sector) is considered, while from 1984 to 1994 the rate of expansion in M2 is considered. The figures are seasonally adjusted and expressed as monthly growth rates.[14] Since the figures on money growth were released with a delay of one month for most of the period considered in the analysis, the UMS series has been lagged by one period.

4.6 Consumption

The inclusion of the percentage change in real per capita consumption (CG) among the state variables is motivated by the consumption-based asset pricing theories (see Lucas, 1978 and Breeden, 1979), which suggest that securities risk and expected return are a function of their covariance with consumption. An index of real consumption was divided by an index of population to obtain the series of real *per capita* consumption. The logarithmic change in the series was then taken to obtain consumption growth. Since monthly data on consumption in Italy are not available, the monthly figures for CG are estimated on the basis of the quarterly data, employing the Chow and Lin (1971) method.[15]

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13 In the first part of the eighties stronger emphasis was given to the rate of growth in total bank lending, while since the middle of the eighties monetary aggregates (M2) have been given greater attention.

14 The data on this variable are only available since 1981.

15 The Chow and Lin (1971) method is a GLS procedure to estimate a series at higher frequencies. The method assumes that the (unobservable) higher frequency series are linearly related to a set of variables which is observable at the desired frequency. The following procedure was used to estimate the monthly figures for CG: first, a monthly estimate of GNP was obtained using the Chow and Lin procedure and the monthly series of industrial production as a reference variable. The Chow and Lin procedure was then used to obtain monthly consumption, using the estimated monthly figures of GNP as a reference variable.

4.7 *Stock market returns*

We also analyze the relation between the factor scores and the return on stock market indices. This choice is based on the assumption that the stock market reacts in advance to unanticipated shifts in the macroeconomic variables; therefore, the inclusion of the market indices should provide some insight into the robustness of the relations between the factor scores and the macroeconomic state variables. Returns were computed for both the equally-weighted index of all shares listed at the Milan Stock Exchange (EWMSE) and the value-weighted market index (VWMSE).

4.8 *Statistical properties of the series*

Table 2 shows the correlation coefficients among the macroeconomic factors.[16] The correlations among the macroeconomic variables are generally small. Higher values are recorded between the raw series and their unexpected component (e.g. MP and UMP), or between the stock market indices EWMSE and VWMSE (equal to 0.96). Higher correlation coefficients show up by construction among CG and the contemporary industrial production variables (MP1 and UMP1), which have been used to estimate the monthly consumption figures. In contrast with previous results obtained for the US by Chen, Roll and Ross (1986b) and for Japan by Hamao (1988), the correlation between the term structure factor (UTS) and the risk premium factor (UPR) is very low. This difference probably depends on the fact that both Chen, Roll and Ross and Hamao employed the return on the long-term bond series to construct both variables, thereby inducing high correlation between these factors. The way in which UPR has been constructed in this work allows us to avoid this problem.

Table 3 shows summary statistics and the autocorrelation coefficients[17] of the series at various lags, together with the values of the Box-Pierce Q-statistic for lags from 1 to 24. As expected, the monthly production variables MP and MP1 show high serial dependence (the Q-statistic is significant at the 1 per cent level). However, the degree of serial correlation drops significantly when the procedures described above are used to estimate the unexpected production variables UMP and UMP1.

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16 The figures in Tables 2 and 3 were also analysed for two sub-periods, but since no appreciable differences emerged only the figures for the entire sample are shown.

17 The autocorrelation coefficients have been calculated as $\rho_k = \frac{c_k}{c_0}$ where ρ_k is the autocovariance of the series at lag k (and therefore ρ_0 is the variance of the input series). The coefficients have been estimated using the Box and Jenkins formula: $c_k = \frac{1}{N-k}\sum_{t=1}^{N-k}(x_t - \bar{x})(x_{t+k} - \bar{x})$.

5. The selection of the macroeconomic factors

Statistical methods do not provide an unequivocal criterion to select the systematic factors which influence stock returns from among the potential candidates. In the previous literature the pervasive factors have therefore been selected either on the basis of empirical considerations, such as their explanatory power in predicting equity returns (see Chen, Roll and Ross, 1986a and Chen and Jordan, 1993) or on the basis of *a priori* beliefs (see McElroy and Burmeister, 1988, Shanken and Weinstein, 1987, He and Ng, 1994). This paper follows the first approach, selecting the macrovariables that are more closely related to the factors influencing securities returns on the basis of their ability to explain the factor scores estimated using Exploratory Factor Analysis (EFA). Although this method is still arbitrary, it has the advantage of reducing the degree of arbitrariness and allowing one to compare the results with previous studies.

The following three-step procedure has been employed: (a) a 5-factor model was estimated[18] using maximum likelihood EFA, in order to calculate the time series of the monthly factor scores for each of the 5 factors (see the Appendix for the details on the methodology used to fit the factor model); (b) each series of the factor scores was regressed on a subset of the macroeconomic factors and (c) the significance of each macrovariable in predicting the stock market factors was evaluated on the basis of an F-test on the joint significance of the coefficient of the macrovariable across the 5 regressions (one for each factor). For example, the p-values shown in the first row (Model 1) of Table 4 were computed by estimating the following set of five equations:

$$Factor\ i = a_i + d_{i1} MP + d_{i2} UI + d_{i3} DEI + d_{i4} UTS + d_{i5} UPR + e_i \qquad i = 1,2,3,4,5$$

and testing the significance of the coefficients of each macrovariable across the five equations simultaneously. For example, for the MP variable the following null hypothesis was tested:

$$H_0:\quad d_{11} = d_{21} = d_{31} = d_{41} = d_{51} = 0\ .$$

This choice was motivated by the possibility that each variable could be significant in any single regression but irrelevant when the regressions are considered jointly. The p-values of the significance tests are shown in Table 4.

The selection procedure was performed by successive steps, in part on the basis of prior beliefs. The search was started by analyzing the significance of the variables which are more closely related to equation (2). Model 1 therefore includes the monthly growth rate of industrial production (MP), unexpected inflation (UI), the revision of inflation expectations (DEI) and the interest rate variables (the change in the slope of the term structure, UTS, and the change in the

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[18] Applying the cross-validation technique suggested by Conway and Reinganum (1988), Panetta (1996) finds that 5 factors are sufficient to explain Italian equity returns.

risk premium, UPR). Both MP and UTS are strongly significant (at the 5 per cent and 1 per cent levels, respectively); among the inflation variables, only UI appears significant, while DEI is not (the p-value of the F-test is equal to 0.23).

The coefficient of the risk premium variable (UPR) is not significant (its p-value is equal to 0.99).[19] This result differs from those presented in the previous literature, in which UPR is often found to be significant. This difference could depend on two factors. First, our definition of the risk premium variable might allow us to distinguish the effects of UPR from those of UTS more precisely than in previous studies: for example, in Chen, Roll and Ross (1986b) and in Hamao (1988) UTS and UPR are both a function of the return on government bonds, so that their significance might be caused by the correlation between equity returns and returns on government bonds. Our definition, which separates the effects of each variable, allows us to estimate more precisely the effect of UTS, which is found to be strongly significant, and that of UPR, which is not significant. However, another explanation for the lack of significance of UPR might be that it largely reflects other factors unrelated to stock returns, such as changes in the degree of competition in the banking sector or the dynamics of leverage for different groups of firms[20], which obtain bank credit on different terms.

In the next regression (Model 2) the contemporaneous production variable MP1 was substituted for lagged production. However, MP1 was not significant (its p-value is equal to 0.74). This suggests that stock returns are influenced by the *announcement* of the ISTAT figures on production, although they refer to month t-2. However, this result could be partially determined by the strong degree of serial correlation detected in MP. Unanticipated production variables were therefore employed, but the results were similar to the previous ones: the effect of the lagged monthly production variable (UMP) became significant at 1 per cent (see Model 3), while the innovation in contemporaneous production (UMP1) remained not significant (see Model 4), although its fit improved relative to MP1.

In the next regression we include the surprise in oil prices (UOILG), the rate of change of the spot lira/US dollar exchange rate (DLUSD) and CG, the rate of growth in real *per capita* consumption (see Model 5). DLUSD and UOILG display strong significance. This result is not surprising: in a country as highly dependent on international trade and oil imports as Italy, both variables have a strong influence on the future growth and profitability of Italian firms. Conversely, the hypothesis that consumption is related to the factor scores is strongly rejected (the p-value is equal to 0.86).[21]

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[19] Model 1 was re-estimated substituting UPR with the spread between the bank rate paid by corporate borrowers in the bottom and top deciles of the distribution of bank lending rates. However, no appreciable changes emerged.

[20] However, this criticism also applies to other measures of UPR used in the literature. For example, Chen, Roll and Ross estimate UPR as the difference between the return on government bonds and the return on bonds with a rating of less than BAA: this measure might reflect changes in the way bonds are rated, changes in the leverage of low-quality firms, or changes in the relative liquidity of the secondary markets for government and junk bonds.

[21] A further regression based on quarterly data (not shown) confirmed that this result is not related to the procedure employed to estimate monthly consumption

The money supply variable UMS was included in Model 7, but was not significant. This result, which is not affected by the introduction of lagged values of UMS (the results are not shown), might reflect the fact that the influence of monetary policy on stock returns is already captured by the interest rate and inflation variables, whose significance deteriorates when UMS is included among the regressors. However, the result could be due to a poor relation between our proxy and the true unexpected shifts in the stance of monetary policy.

Finally, Model 8 includes only the significant macrovariables: lagged unexpected monthly production (UMP), the inflation surprise (UI), the innovation in the term structure (UTS), the unexpected change in the oil price (UOILG) and the change in the lira/US dollar spot exchange rate (DLUSD). The coefficients of all the independent variables are very significant, and no changes emerge with respect to previous results.

In the last two regressions the equally and value weighted indices of stock prices are included among the regressors (see Models 9 and 10). Although the coefficients of both indices are highly significant[22], the differences in the results are not dramatic (only for UI does the p-value increase slightly above 10 per cent in both regressions).[23]

Summarizing the results of this section, close relations emerge between the factor scores and the macroeconomic state variables. The strongest relations are those detected for the innovation in the term structure (UTS), the unexpected change in industrial production (UMP), unexpected inflation (UI), unexpected changes in oil prices (UOILG) and the change in the lira/US dollar spot exchange rate (DLUSD). Viceversa, no relation was detected between the factor scores and variables such as the change in expected inflation (DEI), the surprise in the risk premium (UPR), consumption growth (CG) and unexpected shifts in monetary policy (UMS).

6. Stability analysis

The stability analysis was performed using the following procedure: first, four non-overlapping periods of 48 months (period 1, 2, 3 and 4) were examined, beginning in January 1979 and ending in December 1994. For every share that was continuously listed in any two

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22 The market indices are significant only for the first factor and have no effect on the others. An inspection of the results of the regressions for each single factor (not reported) shows that when the first factor is regressed on the macroeconomic factors and on the stock market index, the coefficient for the latter is approximately equal to 1 and highly significant. Moreover, the high explanatory power of the regression - the R^2 is close to 1 - is almost totally determined by the inclusion of the indices. These findings are not surprising, since the factor scores are simply the return on a portfolio of shares, like the return on the indices; a high correlation is therefore to be expected.

23 A further check of the robustness of the results was performed by estimating the factor scores imposing a Varimax rotation in the exploratory factor analysis and replicating the regression which includes the significant macrovariables (Model 8 in Table 4). All the state variables were also significant in this regression, and the results were similar to those reported in the Table.

subsequent time periods the loadings on each factor were estimated running the following multiple regression:

$$R_{it} = c_i + \beta_{i1}UTS_t + \beta_{i2}DLUSD_t + \beta_{i3}UMP_t + \beta_{i4}UOILG_t + \beta_{i5}UI_t + \varepsilon_t$$

where R_{it} is the return on share i in month t and the right-hand-side variables are those defined previously. Correlation coefficients were then computed between successive loading estimates (the β) both for individual shares and for portfolios. The shares were then ranked in ascending order on the basis of their loading estimates in the estimation period and were assigned to quintile. Portfolio regression tendencies were then evaluated by monitoring variations in quintiles loadings from the estimation period to the subsequent prediction period. Finally, transition matrices were computed to show the proportion of shares which move from one quintile in the estimation period to another in the prediction period. The results are shown in Tables 5-8.

The relation between equity returns and the macroeconomic factors is generally highly unstable for all the state variables. Only a small proportion of shares falls in the same risk quintile after four years (the proportion ranges from 18 per cent for DLUSD to 24.4 per cent for UTS - see Table 5) or in the same or an adjacent quintile (the proportion ranges from 51.6 per cent for the unexpected change in oil prices to 59.2 per cent for UTS). For all the macroeconomic factors the Pearson independence test fails to establish any dependency of the risk quintile in period t on the quintile in period t+1.

The correlations between risk measures for individual securities are generally low; in some cases they are negative (see Table 6). No significant improvement is detected by considering portfolios rather single securities: although when the number of shares expands from 1 to 15 an increasing pattern shows up for DULUSD and UMP, the correlations between the loadings of the portfolios over time are always very low in absolute value: for 15-security portfolios, the average value of the estimated correlation for the 5 macrovariables is equal to 12.5 per cent, ranging from -16.9 per cent for UOILG to 39 per cent for UMP. Moreover, all the state variables display negatively correlated loadings in at least one of the sub-periods.

Portfolio risk measures do not show any tendency to revert towards particular values (see Table 7): the estimates of the coefficients in each four-year period are almost totally unrelated to their value in the following period. This pattern is detected for all risk quintiles and in each pair of sub-periods. Moreover, for a high proportion of shares the sensitivities to the macroeconomic state variables change sign from the estimation period to the control period: as Table 8 shows, on average the proportion of sign reversals over two successive periods is close to 50 per cent (see panel A of the Table), ranging from 44 per cent for UI to 55 per cent for UOILG. This finding is also common to all risk classes (see Panel B of the Table).

These results differ considerably from those obtained in previous studies in a market model framework, both for individual securities and for portfolios. For the US Francis (1979) found that the proportion of shares falling in the same (same or adjacent) risk quintile is 40 per cent (79 per cent) in the period 1961-71. For the UK the corresponding figures are 33 per cent (72 per cent) - see Dimson and Marsh (1983). Moreover, in market model studies the correlations among β estimates for subsequent periods are always found to be positive, approaching unity as the number of securities in the portfolio increases.

Statistical tests were also performed on the stability of the risk measures. As already mentioned, no effort is made to model the changes in the loadings; rather, the analysis concentrates on formally assessing the degree of general instability of the risk parameters. As before, the sample period was divided into four non-overlapping subperiods of 48 months each, and stability was tested for all individual securities that were continuously listed for any two subsequent sub-periods. The stability hypothesis was tested using the Wald test suggested by Honda and Ohthani (1986):

$$W = (\hat{q}_1 - \hat{q}_2)' V^{-1} (\hat{q}_1 - \hat{q}_2) \tag{4}$$

where $\hat{q}_1$ and $\hat{q}_2$ are the estimates of the loadings in period t and t+1 and $V = \mathrm{var}(\hat{q}_1) + \mathrm{var}(\hat{q}_2)$. Under the null hypothesis that the two estimates have the same expected value, the statistic (4) has a chi-squared distribution with degrees of freedom equal to the number of coefficients for which the stability is tested. The advantage of the test is that for reasonably large samples it is valid even if the error variances differ in the two sub-periods.[24]

Table 9 reports summary statistics on the proportion of rejections of the null hypothesis of stability of the risk measures for individual shares for the 1 per cent, 5 per cent and 10 per cent significance level. The results of the Wald tests confirm the presence of a non-trivial degree of instability for the loadings of UTS (in the first two pair of periods), for the loadings of UOILG (between period 2 and 3) and for the loadings of DLUSD (in periods 3 and 4). However, the phenomenon is less pronounced than one would expect from the evidence presented in the previous tables. An inspection of regression results for the individual securities reveals that although the estimates of the risk measures change considerably over different sub-periods, the coefficients are estimated imprecisely, so that it is very difficult to detect any structural break in the coefficients.

This finding suggests that the results discussed so far might depend merely on the lack of precision of the estimates. In order to verify this hypothesis, a number of checks were performed.

First, the analysis was replicated by splitting the sample period into only two sub-periods: if the lack of precision of the previous estimates depends on the fact that the length of the sub-periods in which the regression coefficients have been estimated (48 months each) is too short, then using two periods of 96 months each should reduce the noise. The results for the entire sample (reported in Table 10) do not differ from the previous findings: from the estimation period to the subsequent control period the estimated values of the risk measures change substantially (see Panel A of the Table), display low correlations (see Panel B; for UMP and UI the correlation is negative); on average, almost 50 per cent of the loadings of the single shares

[24] See Greene (1993), p. 215.

exhibit a sign reversal (see Panel C). Also in this case, the results of the Wald test partly confirm the results (see Panel D; the estimated risk measures exhibit high standard errors).

A second check of the robustness of the results was conducted by estimating the risk measures on the basis of univariate regressions: this choice was prompted by the fact that the changes in each single estimate of the risk measures could be related to shifts in the correlations between the factors or by the fact that new factors become relevant in successive sub-periods. However, the results (reported in Table 11) are similar to those reported previously; differences emerge for the correlation between successive estimates of the sensitivity to UI (which, in comparison with the results reported in Table 10, remains small but becomes positive) and for the proportion of shares which experience a sign reversal (which decreases to approximately 40 per cent).

Another possible source of concern about the robustness of the results is that they could reflect an erroneous specification of the model. In fact, although the statistical procedure used in this paper provides an coherent framework to select the macrovariables, it is not possible to eliminate completely arbitrariness from the selection procedure. Therefore, further checks of the results consisted in replicating the analysis by changing the factor proxies. First, the stability analysis was replicated by substituting the unanticipated growth of lagged production (UMP) for actual growth (MP), which showed some significance in predicting the factor scores. Second, the regressions were replicated by using the factor proxies employed by Chen, Roll and Ross, (1986b)[25]. However, in both cases the results (not shown) did not differ from the previous ones, confirming that the risk measures of Italian equities change considerably over time.

Finally, we replicated the regressions by splitting the observations into odd and even months: if the changes in the estimated risk measures reflect true instability rather than estimation noise, this method of splitting the data should reduce the instability of the estimates. The results obtained using odd and even months, shown in Table 12, are consistent with the hypothesis that the problem is true instability rather than estimation error: in fact, the stability of the loadings improves for all factors except the lira/dollar exchange rate: in particular, the loadings measured for the even months are close to those of the odd months and the proportion of sign reversals decreases considerably by comparison with the results of Table 10; furthermore, the correlation coefficients between the risk measures for odd and even months are always positive and higher in absolute value than those in Table 10.

7. Conclusions

Despite the extensive literature on testing the stability of risk measures in a market model framework, only a very small number of papers has analyzed the stability of the linkages between the macroeconomic state variables and equity returns. In an APT framework this is particularly disturbing, since the theory does not identify the macroeconomic factors. As Fama (1991) noted: "*Since multifactor models offer at best vague predictions about the variables that are important in returns and expected returns, there is the danger that measured relations*

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[25] The factor proxies used to replicate Chen, Roll and Ross (1986b) are MP, DEI, UI, UPR and UTS.

between returns and economic factors are spurious, the result of special features of a particular sample (factor dredging). Thus, the Chen, Roll and Ross tests, and future extensions, warrant extended robustness checks." In fact, an important check of robustness of any asset pricing model consists in analyzing the robustness of the return generating process implied by the model, and in particular the stability of the factor loadings.

This paper has analyzed the stability of relations between macroeconomic variables and Italian equity returns. Analyzing sixteen years of returns, the paper has shown that instability is indeed a serious problem. The sensitivities to the macroeconomic variables are highly unstable for both individual securities and portfolios; moreover, the instability is not limited to a single time period or to shares in a particular risk class, but has been detected in each of the subperiods that have been considered and for shares in all risk classes. In the multifactor framework, the consequences of instability are much greater than in the market model framework: for a large proportion of securities the estimated risk measures change sign during the sample period.

Our results suggest that the changes in the sensitivity to the macroeconomic factors are due to true instability, rather than to noise determined by estimation error. The instability of the risk measures could have several causes. First, instability might arise from different responses of stock returns to the economic fundamentals in different phases of the cycle: during a recession, an unexpected rise in economic activity would likely cause an increase of stock prices, while during an expansion it could be interpreted negatively, generating inflationary fears and a fall in share prices. Second, mispecifications in the return generating process might cause instability in the estimated risk measures even if the "true" structural risk parameters were stable. This might happen if the linear relation hypothesized by the APT is not correct: linear relationships can approximate non-linear relations adequately only for short time intervals, while for longer intervals a linear relationship would cause instability in the parameter estimates. Mispecifications of the return generating process might also arise if a relevant variable which is not orthogonal to the variables included on the regression is omitted or if the systematic factors change over time.

The instability of the relation between securities returns and macroeconomic factors can lead to severe bias in the risk measures. If models in which the return generating process is driven by macroeconomic variables are used for any application that requires estimates of expected returns (e.g. portfolio selection or cost of capital evaluation), such instability can lead to misleading conclusions. The finding that the loadings are unstable also has strong implications for the analysis of the factor structure of securities returns: in fact, all previous studies on the number of pervasive factors in securities returns have assumed the stability of the factor loadings. Our results suggest that further research should be devoted to investigate the causes and the consequences of the instability of the risk measures.

Appendix: Data sources and the estimation of the factor scores

This Appendix indicates the sources from which the series described in the text were drawn and describes the methodology used to estimate the factor scores.

Share prices

The data used is monthly returns for all shares listed on the Milan Stock Exchange (MSE) for the period December 1978 to December 1994, drawn from the Bank of Italy share price database. The average number of shares included in the sample in each month is 252, ranging from 159 in 1980 to 337 in 1991. Returns are defined as

$$R_t = \ln\left(P_t + d_t\right) - \ln P_{t-1}$$

where d_t is the dividend paid during month t and P_t is the price of the share, adjusted for changes in the capital structure due to scrip issues or right issues, etc. The returns on an equally-weighted (EWMSE) and on a value-weighted (VWMSE) market index have been calculated based on all shares listed in the market in each month.

Industrial production

The industrial production series are those published monthly by ISTAT and are corrected to take account of the different number of working days in each month. The electricity series are those released by ENEL (the Italian state-owned electricity supplier), corrected to take account of the different number of working days in each month.

International factors

Exchange rates. The lira/US dollar spot and forward exchange rates were drawn from the Bank for International Settlement (BIS) database. Data refer to contracts registered on the last day of each month. Forward rates are contracts with a 1-month maturity.

Oil prices. Oil prices are average dollar prices recorded during each month for the three grades of oil with the largest weight in the basket of Italian crude oil imports (Brent, Dubai and WTI). The figures were drawn from the IFS database.

Inflation

Inflation is calculated on the basis of the cost of living index, published monthly by ISTAT.[26] Although the final figure is released with one month delay, precise anticipations

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[26] ISTAT calculates also a second price index at monthly frequency, the consumer price index (CPI). However, the CPI is released only after a few months delay.

based on prices recorded in the 9 largest cities are announced in the last week of the same month.

Term structure

The data were collected from the monthly statistics published by the Bank of Italy in the *Supplemento al Bollettino Statistico.* The holding period return on long-term government bonds (LHPR) is calculated as the monthly percentage change in the capitalization index of BTPs (*Buoni Poliennali del Tesoro*, Italian government long-term fixed coupon bonds) net of withholding tax; since the Bank of Italy only started to publish data on the BTP index in 1983, BTP prices and coupon payments for the previous period were collected from the *Bollettino Ufficiale* of the Milan Stock Exchange.

The short-term interest rate (SR) is the return on BOTs *(Buoni Ordinari del Tesoro*, Italian government T-bills) net of withholding tax; since the shortest maturity of BOTs is three months, the returns were compounded back to yield one-month rates.

Risk premia

The monthly data on the average and monthly rates have been taken from the monthly series published by the Bank of Italy. The quarterly data on bank lending rates paid by corporate borrowers in the bottom and top deciles of the distribution of bank lending rates drawn from the database of the *Centrale dei Rischi* (Central Credit Register, CCR), a department of the Bank of Italy which collects quarterly data on the interest rates charged by a group of banks accounting for over 80 per cent of total bank lending in Italy.

The estimation of the factor scores

The factor scores of the 5-factor benchmark were estimated using maximum likelihood Exploratory Factor Analysis (EFA). First, the covariance matrix S of the monthly returns of the 90 shares which were continuously listed on the Milan Stock Exchange from December 1978 to December 1994 was calculated. EFA was then used to decompose the covariance matrix into the 90 x 5 factor loading matrix Γ and the 90 x 90 residual risk matrix Φ:

$$S = \Gamma\Gamma' + \Phi\,.$$

The 90x1 vector of portfolio weights w_j for each of the 5 benchmark portfolios was then estimated using the Minimum Idiosyncratic Risk Portfolios (Mirp) procedure suggested by Lehmann and Modest (1988):

$$\begin{aligned}
&\underset{\mathrm{wj}}{\mathrm{Min}}\ w_j'\Phi w_j \qquad j = 1,\ldots..,5 \\
&s.t.\ \ w_j'\Gamma_k = 0 \quad \text{for } \mathrm{j} \neq \mathrm{k} \\
&\qquad w_j'\iota\ = 1
\end{aligned}$$

where ι is a vector of ones and Γ_k is the vector of the loadings of each security to factor k (i.e. the kth column of the loading matrix Γ). The portfolio weights were then multiplied by the monthly excess returns on the securities to determine the monthly time series of the returns on the 5 benchmark portfolios (the factor scores).

Table 1

Glossary of the macroeconomic factors

Symbol	Variable
MP	Lagged percentage change in monthly industrial production
MP1	Percentage change in monthly industrial production
UMP	Lagged unexpected change in monthly industrial production
UMP1	Unexpected change in monthly industrial production
UI	Unexpected inflation
DEI	Change in expected inflation
UTS	Term structure factor (unexpected return on long bonds)
UPR	Change in risk premium (change of the spread between the average and the minimum bank rate)
UOILG	Unexpected change in oil prices (in US dollars).
DLUSD	Percentage change in the lira/US dollar spot exchange rate
CG	Percentage change in real per capita consumption
UMS	Unexpected money supply growth (announced minus actual growth of monetary aggregates)
EWMSE	Return on an equally-weighted index of the Milan Stock Exchange (MSE)
VWMSE	Return on a value-weighted index of the MSE

Table 2

Correlation matrix of the macroeconomic factors

MP is the lagged percentage change in monthly industrial production. MP1 is the percentage change in monthly industrial production. UMP is the lagged unexpected change in monthly industrial production. UMP1is the unexpected change in monthly industrial production. UI is unexpected inflation. DEI is the modification in expected inflation. UTS is the unexpected return on long bonds. UPR is the unexpected change in the risk premium (change of the spread between the average and the minimum bank rate). UOILG is unexpected change in oil prices (in US dollars). DLUSD is the percentage change in the lira/US dollar spot exchange rate. CG is the percentage change in real per capita consumption. UMS is the unexpected money supply growth (announced minus actual growth of monetary aggregates). EWMSE is the return on an equally-weighted index of the Milan Stock Exchange (MSE). VWMSE is the return on a value-weighted index of the MSE. The correlations are estimated on a sample of 192 observations (monthly data from January 1979 to December 1994). The UMS series starts in January 1981. The data sources are described in the Appendix.

State Variables	MP	MP1	UMP	UMP1	UI	DEI	UTS	UPR	UOILG	DLUSD	CG	UMS	EWMSE	VWMSE
MP	1													
MP1	-0.044	1												
UMP	0.500	-0.028	1											
UMP1	-0.057	0.358	-0.069	1										
UI	-0.023	0.008	-0.070	0.051	1									
DEI	0.044	-0.202	-0.016	-0.106	0.023	1								
UTS	-0.039	-0.008	0.022	0.075	0.046	0.052	1							
UPR	-0.186	0.081	-0.106	0.030	-0.028	0.030	0.035	1						
UOILG	0.033	0.055	0.119	-0.009	0.028	0.060	0.045	-0.047	1					
DLUSD	0.000	0.035	-0.085	0.039	-0.070	0.054	-0.023	0.041	0.080	1				
CG	-0.035	0.839	0.001	0.408	0.042	-0.206	-0.019	0.072	0.066	0.019	1			
UMS	-0.002	0.123	0.035	0.046	0.055	0.072	0.063	0.003	0.076	-0.105	0.119	1		
EWMSE	-0.150	0.027	-0.158	0.150	-0.087	0.042	0.222	0.008	0.014	0.089	-0.005	0.075	1	
VWMSE	-0.159	0.006	-0.125	0.127	-0.113	0.008	0.245	-0.007	-0.010	0.110	-0.013	0.051	0.960	1

Table 3

Summary statistics of the macroeconomic factors

MP is the lagged percentage change in monthly industrial production. MP1 is the percentage change in monthly industrial production. UMP is the lagged unexpected change in monthly industrial production. UMP1is the unexpected change in monthly industrial production. UI is unexpected inflation. DEI is the modification in expected inflation. UTS is the unexpected return on long bonds. UPR is the unexpected change in the risk premium (change of the spread between the average and the minimum bank rate). UOILG is unexpected change in oil prices (in US dollars). DLUSD is the percentage change in the lira/US dollar spot exchange rate. CG is the percentage change in real per capita consumption. UMS is the unexpected money supply growth (announced minus actual growth of monetary aggregates). EWMSE is the return on an equally-weighted index of the Milan Stock Exchange (MSE). VWMSE is the return on a value-weighted index of the MSE. The statistics are estimated on a sample of 192 observations (monthly data from January 1979 to December 1994). The UMS series starts in January 1981. The symbols ρ_i are autocorrelation coefficients at lag i calculated as $r_k = c_k / c_0$ where r_k is the autocovariance of the series at lag k (and therefore r_0 is the variance of the input series); the coefficients have been estimated using the Box and Jenkins formula $c_k = \frac{1}{N-k}\sum_{t=1}^{N-k}(x_t - \bar{x})(x_{t+k} - \bar{x})$. Q(24) is the Box-Pierce statistic for the first 24 lags. The symbol * indicates that the Q-statistic is significant at the 1 per cent level. Data sources are described in the Appendix.

State Variables	No. of Obs.	Mean	S. D.	Min.	Max.	ρ_1	ρ_2	ρ_3	ρ_6	ρ_{12}	ρ_{24}	Q(24)
MP	192	0.13	2.03	-11.35	6.65	-0.387	-0.042	0.080	0.159	-0.031	-0.151	* 74.8
MP1	192	0.14	2.03	-11.35	6.65	-0.383	-0.043	0.080	0.165	-0.033	-0.200	* 71.9
UMP	192	-0.06	1.68	-5.6	5.37	0.112	0.038	0.123	0.091	0.124	-0.175	34.2
UMP1	192	-0.04	1.70	-9.87	6.92	-0.048	0.077	0.090	0.045	-0.055	-0.076	19.8
UI	192	-0.01	0.35	-1.36	1.47	0.117	-0.184	-0.066	0.051	-0.120	-0.010	31.9
DEI	192	-0.02	0.121	-0.46	0.55	-0.042	-0.197	-0.058	-0.005	-0.025	0.046	28.4
UTS	192	0.02	0.75	-3.47	4.04	0.180	0.133	0.174	0.080	-0.091	0.063	42.1
UPR	192	-0.00	0.02	-0.20	0.04	0.029	0.001	0.032	0.025	-0.037	-0.059	10.5
UOILG	192	-0.21	8.41	-31.27	41.85	-0.103	-0.047	0.027	-0.116	-0.022	0.037	31.4
DLUSD	192	0.25	3.33	-7.78	13.83	0.117	0.091	0.072	-0.120	-0.002	0.033	22.8
CG	192	0.20	2.42	-11.65	8.87	-0.439	-0.025	0.092	0.241	-0.018	-0.191	* 96.3
UMS	168	0.08	0.64	-2.03	2.23	-0.040	0.105	-0.014	0.074	0.224	0.214	38.3
EWMSE	192	1.41	6.73	-17.37	26.67	0.223	0.046	0.171	0.062	0.052	-0.032	37.4
VWMSE	192	1.34	6.99	-19.74	25.63	0.154	0.004	0.166	0.077	0.031	-0.019	31.0

Table 4

Time series regressions of the factor scores on the macroeconomic factors

MP is the lagged percentage change in monthly industrial production. MP1 is the percentage change in monthly industrial production. UMP is the lagged unexpected change in monthly industrial production. UMP1is the unexpected change in monthly industrial production. UI is unexpected inflation. DEI is the modification in expected inflation. UTS is the unexpected return on long bonds. UPR is the unexpected change in the risk premium (change of the spread between the average and the minimum bank rate). UOILG is unexpected change in oil prices (in US dollars). DLUSD is the percentage change in the lira/US dollar spot exchange rate. CG is the percentage change in real per capita consumption. UMS is the unexpected money supply growth (announced minus actual growth of monetary aggregates). EWMSE is the return on an equally-weighted index of the Milan Stock Exchange (MSE). VWMSE is the return on a value-weighted index of the MSE. The table shows the p-values obtained running time series regressions of each of the 5 factors estimated using maximum likelihood factor analysis on the macroeconomic factors and calculating an F-test for the null hypothesis that the coefficient of each RHS variable is jointly equal to zero in the five regressions. For example, the value in the MP box in Model 1 has been obtained by estimating the regression simultaneously

$Factor\ i = a_i + d_{i1}YP + d_{i2}UI + d_{i3}DEI + d_{i4}UTS + d_{i5}UPR + e_i$ for each factor i=1,...5

and computing the p-value for the null hypothesis that each coefficient is equal to zero in all regressions contemporaneously, i.e.:
H_0: $\delta_{11} = \delta_{21} = \delta_{31} = \delta_{41} = \delta_{45} = 0$. The regressions have been estimated on monthly data from January 1979 to December 1994. Data sources are described in the Appendix.

	Production factors				Inflation factors		Interest rate factors		International factors		Cons.	Money supply	Stock market indices	
	MP	MP1	UMP	UMP1	UI	DEI	UTS	UPR	UOILG	DLUSD	CG	UMS	EWMSE	VWMSE
Model 1	0.042				0.041	0.233	0.000	0.990						
Model 2		0.740			0.084	0.409	0.000	0.987						
Model 3			0.001		0.081	0.234	0.000	0.995						
Model 4				0.145	0.051	0.248	0.000	0.997						
Model 5			0.024		0.050	0.303	0.000	0.992	0.033		0.915			
Model 6			0.047		0.074	0.322	0.000	0.993	0.040	0.032	0.865			
Model 7			0.020		0.224	0.821	0.000	0.991	0.012	0.028	0.680	0.948		
Model 8			0.012		0.040		0.001		0.033	0.011				
Model 9			0.145		0.165		0.009		0.029	0.029			0.000	
Model 10			0.031		0.137		0.009		0.006	0.021				0.000

Table 5

Transition Matrices for Risk Measures on the Milan Stock Exchange

The matrices were computed by estimating in each sub-period risk measures for individual stocks by running the following multiple regression:

$$R_{i,t} = c_i + b_{i1}UTS_t + b_{i2}DLUSD_t + b_{i3}UMP_t + b_{i4}UOILG_t + b_{i5}UI_t + e_{it}$$

where R_{it} is the return on share i in month t, UMP is the lagged unexpected change in monthly industrial production, UI is unexpected inflation, UTS is the unexpected return on long bonds, UOILG is unexpected change in oil prices (in US dollars), DLUSD is the percentage change in the lira/US dollar spot exchange rate. For each pair of periods (t, t+1) the table shows the proportion of shares which change risk quintile from the estimation period t to the subsequent control period t+1. The last column of the table shows the proportion of shares that for each pair of periods (t,t+1) remains in the same or adjacent risk quintile from the estimation period t to the control period t+1. The chi-square test is the Pearson independence test on the null hypothesis that the initial and subsequent ranking of the individual securities' risk measures are independent (See Mood, Graybill and Boes, 1974) and is distributed as a chi-square with 16 degrees of freedom. The null hypothesis is evaluated at the 5 per cent confidence level. In total 417 observations were used to estimate these matrices (113 shares in periods 1 and 2, 120 in period 2 and 3 and 184 in period 3 and 4). Data sources are described in the Appendix.

	Percentage in Risk Class in Period t+1						
Risk Class in Period t	High	2	3	4	Low	Same or adjacent	Chi-square Statistic
			Term structure spread (UTS)				
High	28.8	21.3	13.8	16.3	20.0	50.1	
2	20.7	20.7	23.2	20.7	14.6	64.6	
3	13.6	23.5	24.7	22.2	16.0	70.4	Not Reject
4	15.9	13.4	19.5	24.4	26.8	70.7	
Low	19.5	22.0	18.3	17.1	23.2	40.3	
			Change in the lira/US$ exchange rate (DLUSD)				
High	25.0	25.0	10.0	18.8	21.3	50.0	
2	22.0	19.5	20.7	17.1	20.7	62.2	
3	18.5	22.2	16.0	24.7	18.5	62.9	Not Reject
4	23.2	11.0	25.6	15.9	24.4	65.9	
Low	9.8	23.2	26.8	24.4	15.9	40.3	
			Unexpected monthly production (UMP)				
High	23.8	15.0	26.3	17.5	17.5	38.8	
2	18.3	19.5	22.0	20.7	19.5	59.8	
3	19.8	27.2	16.0	17.3	19.8	60.5	Not Reject
4	24.4	17.1	14.6	25.6	18.3	58.5	
Low	12.2	22.0	20.7	19.5	25.6	45.1	
			Unexpected change in oil prices (UOILG)				
High	20.0	15.0	22.5	23.8	18.8	35.0	
2	22.0	20.7	15.9	22.0	19.5	58.6	
3	13.6	24.7	21.0	19.8	21.0	65.5	Not Reject
4	19.5	25.6	22.0	14.6	18.3	54.9	
Low	23.2	14.6	18.3	20.7	23.2	43.9	
			Unexpected inflation (UI)				
High	16.3	25.0	16.3	22.5	20.0	41.3	
2	20.7	22.0	19.5	17.1	20.7	62.2	
3	21.0	24.7	18.5	14.8	21.0	58.0	Not Reject
4	19.5	17.1	19.5	25.6	18.3	63.4	
Low	20.7	12.2	25.6	20.7	20.7	41.4	

Table 6

Correlation Coefficients between Risk Measures on the Milan Stock Exchange

The correlation coefficients were estimated using the following procedure: (a) in each subperiod the sensitivity of individual shares to each macroeconomic factor was estimated by running the following multiple regression:

$$R_{i,t} = c_i + b_{i1}UTS_t + b_{i2}DLUSD_t + b_{i3}UMP_t + b_{i4}UOILG_t + b_{i5}UI_t + e_{it}$$

where R_{it} is the return on share i in month t, UMP is the lagged unexpected change in monthly industrial production, UI is unexpected inflation, UTS is the unexpected return on long bonds, UOILG is unexpected change in oil prices (in US dollars), DLUSD is the percentage change in the lira/US dollar spot exchange rate; (b) securities were then ranked by their loading estimates and portfolios of *n* securities were then formed by including the first *n* securities in portfolio 1, the next *n* securities for portfolio 2, and so on; (c) the correlation between the portfolio sensitivities to each factor between each pair of periods (t,t+1) was then calculated. In total 417 observations were used to estimate these matrices (113 shares in periods 1 and 2, 120 in period 2 and 3 and 184 in period 3 and 4). Data sources are described in the Appendix.

No. of Securities in Portfolio	Periods 1 and 2	Periods 2 and 3	Periods 3 and 4	Average
		Term structure spread (UTS)		
1	-17.6	7.2	21.4	3.7
2	-25.3	10.9	27.7	4.4
4	-34.9	13.2	36.7	5.0
7	-23.5	17.1	49.8	14.4
10	-35.1	25.2	47.7	12.6
15	-61.2	25.8	65.1	9.9
		Change in the lira/US$ exchange rate (DLUSD)		
1	6.5	17.9	-2.3	7.3
2	11.4	27.4	-3.3	11.8
4	15.2	37.4	-5.5	15.7
7	6.8	53.3	-6.6	17.8
10	18.2	55.6	-13.7	20.0
15	27.1	75.8	-21.7	27.0
		Unexpected monthly production (UMP)		
1	10.4	2.6	-1.8	3.7
2	13.9	3.4	-2.6	4.9
4	26.9	11.7	0.1	12.9
7	19.4	39.1	0.7	19.7
10	24.3	35.3	-5.3	18.1
15	45.4	75.9	-4.6	38.9
		Unexpected change in oil prices (UOILG)		
1	-14.8	22.2	-5.9	0.5
2	-18.1	31.6	-8.6	1.6
4	-33.8	40.7	-10.6	-1.2
7	-37.2	68.1	-16.8	4.7
10	-37.2	64.8	-28.2	-0.2
15	-69.7	61.5	-40.4	-16.2
		Unexpected inflation (UI)		
1	4.8	1.3	0.5	2.2
2	9.2	3.2	0.8	4.4
4	15.2	3.4	-0.2	6.1
7	30.2	-2.9	-5.7	7.2
10	31.3	1.7	-9.3	7.9
15	31.7	2.1	-24.5	3.1
No. of Obs.	113	120	184	139

Table 7

Stability and Regression Tendency of Risk Estimates Grouped by Quintiles

For each macroeconomic factor the quintiles were formed by (a) estimating in each subperiod the sensitivity of individual shares to each macroeconomic factor with the following multiple regression:

$$R_{i,t} = c_i + b_{i1}UTS_t + b_{i2}DLUSD_t + b_{i3}UMP_t + b_{i4}UOILG_t + b_{i5}UI_t + e_{it}$$

where R_{it} is the return on share i in month t, UMP is the lagged unexpected change in monthly industrial production, UI is unexpected inflation, UTS is the unexpected return on long bonds, UOILG is unexpected change in oil prices (in US dollars), DLUSD is the percentage change in the lira/US dollar spot exchange rate; (b) ranking the shares by their estimated risk measure in the earlier of each pair of periods; (c) portfolio risk was calculated by averaging the risk of the individual securities. In total 417 observations were used to estimate these matrices (113 shares in periods 1 and 2, 120 in period 2 and 3 and 184 in period 3 and 4). Data sources are described in the Appendix.

	Average Risk Measures within Quintiles in the Estimation Period (Asterisked) and Subsequent Period					
Quintile	Period 1*	Period 2	Period 2*	Period 3	Period 3*	Period 4
	Term structure spread (UTS)					
Low	-2.18	-2.01	-10.16	1.97	-0.58	1.80
2	0.03	-2.26	-6.36	1.88	1.57	2.16
3	1.06	-5.03	-4.24	2.72	2.59	2.43
4	2.11	-2.74	-1.38	3.18	3.78	2.43
High	4.54	-4.94	3.17	2.23	5.91	2.46
	Change in the lira/US$ exchange rate (DLUSD)					
Low	-0.58	0.08	-0.50	0.35	-0.01	-0.01
2	-1.17	-0.04	-0.11	0.50	0.25	-0.09
3	0.06	0.12	0.10	0.36	0.42	-0.01
4	0.33	0.03	0.35	0.53	0.64	-0.07
High	0.81	0.13	0.72	0.56	0.94	-0.04
	Unexpected monthly production (UMP)					
Low	-1.82	-0.05	-1.26	-0.32	-0.81	-0.57
2	-1.12	-0.21	-0.39	-0.21	-0.37	-0.38
3	-0.83	-0.11	0.02	-0.16	-0.08	-0.52
4	-0.59	-0.05	0.41	-0.10	0.17	-0.77
High	-0.15	0.35	1.22	-0.07	0.73	-0.38
	Unexpected change in oil prices (UOILG)					
Low	-0.16	0.22	-0.04	-0.17	-0.29	0.05
2	-0.04	0.14	0.07	-0.15	-0.17	0.06
3	0.06	0.17	0.14	-0.11	-0.10	0.03
4	0.17	0.18	0.24	-0.10	-0.04	-0.03
High	0.38	0.12	0.39	-0.10	0.06	0.04
	Unexpected inflation (UI)					
Low	-5.44	-3.25	-9.15	-2.48	-11.64	4.60
2	-3.15	-2.63	-5.21	-3.08	-5.67	3.83
3	-2.04	-4.51	-2.47	-4.24	-2.71	5.65
4	-0.39	-2.86	-0.02	-2.03	0.09	2.43
High	1.53	-2.12	3.59	-2.44	5.44	3.62

Table 8

Stability of the Risk Measures for the Milan Stock Exchange: Proportion of Sign Reversals

In each subperiod the sensitivity of individual shares to each macroeconomic factor was estimated by running the following multiple regression:

$$R_{i,t} = c_i + b_{i1}UTS_t + b_{i2}DLUSD_t + b_{i3}UMP_t + b_{i4}UOILG_t + b_{i5}UI_t + e_{it}$$

where R_{it} is the return on share i in month t, UMP is the lagged unexpected change in monthly industrial production, UI is unexpected inflation, UTS is the unexpected return on long bonds, UOILG is unexpected change in oil prices (in US dollars), DLUSD is the percentage change in the lira/US dollar spot exchange rate. Panel A shows for each pair of periods (t, t+1) the proportion of securities for which the sign of the sensitivity to each factor changes from the estimation period t to the subsequent control period t+1. Panel B reports the proportion of securities inside each risk quintile for which a sign reversal occurs. In total 417 observations were used to estimate these matrices (113 shares in periods 1 and 2, 120 in period 2 and 3 and 184 in period 3 and 4). Data sources are shown in the Appendix.

	Term Structure Spread (UTS)	Change in lira/US$ exchange rate (DLUSD)	Unexpected Production (UMP)	Unexpected change in oil prices (UOILG)	Unexpected Inflation (UI)
Panel A: Sign Reversals by Period					
Period 1 and 2	67.0	45.6	47.6	41.7	33.0
Period 2 and 3	71.7	38.3	45.0	70.0	38.3
Period 3 and 4	14.1	56.5	48.9	53.8	62.5
Panel B: Sign Reversals by Risk Quintile					
High	13.0	11.5	6.4	11.5	8.8
4	7.4	13.5	7.1	14.0	10.0
3	10.1	7.9	9.3	11.1	9.1
2	8.1	8.4	13.0	9.1	10.3
Low	5.9	7.1	11.5	9.8	9.6

Table 9

Wald Test on the Stability of Individual Securities Risk Measures

In each subperiod the sensitivity of individual shares to each macroeconomic factor was estimated by running the following multiple regression:

$$R_{i,t} = c_i + b_{i1}UTS_t + b_{i2}DLUSD_t + b_{i3}UMP_t + b_{i4}UOILG_t + b_{i5}UI_t + e_{it}$$

where R_{it} is the return on share i in month t, UMP is the lagged unexpected change in monthly industrial production, UI is unexpected inflation, UTS is the unexpected return on long bonds, UOILG is unexpected change in oil prices (in US dollars), DLUSD is the percentage change in the lira/US dollar spot exchange rate. A Wald test was then computed to test the null hypothesis of stability of the risk measures of the single securities in each pair of subsequent periods (t, t+1). The table shows the proportion of rejections of the stability of the risk measures for individual securities at the 1%, 5% and 10% significance levels. In total 417 observations were used (113 shares in periods 1 and 2, 120 in period 2 and 3 and 184 in period 3 and 4). Data sources are described in the Appendix.

	Proportion of Rejections at Different Significance Level		
Variables	1%	5%	10%
	Period 1 and 2: January 1979-December 1986		
Term Structure Spread (UTS)	5.8	12.6	22.3
Lira/US dollar exchange rate (DLUSD)	1.0	1.0	2.0
Unexpected Production (UMP)	1.0	2.9	10.8
Unexpected change in oil price (UOILG)	1.9	3.9	7.8
Unexpected Inflation (UI)	1.0	1.0	5.8
	Period 2 and 3: January 1983 - December 1990		
Term Structure Spread (UTS)	2.5	16.7	33.3
Lira/US dollar exchange rate (DLUSD)	0.8	4.1	7.5
Unexpected Production (UMP)	0.8	0.8	1.7
Unexpected change in oil price (UOILG)	0.4	15.6	30.8
Unexpected Inflation (UI)	0.8	1.7	4.1
	Period 3 and 4: January 1987 - December 1994		
Term Structure Spread (UTS)	0.5	1.1	3.3
Lira/US dollar exchange rate (DLUSD)	0.5	10.3	19.0
Unexpected Production (UMP)	0.5	4.6	5.5
Unexpected change in oil price (UOILG)	1.6	6.0	8.2
Unexpected Inflation (UI)	0.5	4.3	12.5

Table 10

Stability of the Risk Measures: Entire Sample Multivariate Estimates

The figures in the table have been obtained by splitting the sample into two non-overlapping periods of equal length (96 months) and estimating in each sub-period the risk measures for individual securities by running the following multiple regression:

$$R_{i,t} = c_i + b_{i1}UTS_t + b_{i2}DLUSD_t + b_{i3}UMP_t + b_{i4}UOILG_t + b_{i5}UI_t + e_{it}$$

where R_{it} is the return on share i in month t, UMP is the lagged unexpected change in monthly industrial production, UI is unexpected inflation, UTS is the unexpected return on long bonds, UOILG is unexpected change in oil prices (in US dollars), DLUSD is the percentage change in the lira/US dollar spot exchange rate. Panel A shows the average value of the risk measures estimates of individual securities in each period. Panel B shows the correlation coefficients between the risk measures estimates of individual securities in each period. Panel C shows the proportion of individual securities for which the sign of the sensitivity to each factor changes from period 1 to period 2. Panel D reports the proportion of rejections of the stability hypothesis of the risk measures for the individual securities using a Wald Test. Data sources are described in the Appendix.

	Term Structure Spread (UTS)	Change in the lira/US$ exchange rate (DLUSD)	Unexpected Production (UMP)	Unexpected change in oil price (UOILG)	Unexpected Inflation (UI)
			Panel A: Average Value of Estimates		
Period 1	0.46	0.10	-0.63	0.12	-1.70
Period 2	2.23	0.13	-0.27	-0.09	-0.87
			Panel B: Correlation of Estimates		
Period 1 and 2	9.9	24.3	-18.7	7.3	-3.5
			Panel C: Proportion of Sign Reversals		
Period 1 and 2	35.4	40.5	34.1	70.8	56.9
			Panel D: Proportion of Rejection (Wald Test)		
Confidence level:					
1%	5.1	1.3	1.3	3.8	1.3
5%	15.2	1.3	6.3	10.2	6.3
10%	24.1	1.3	11.4	21.6	12.7

Table 11

Stability of the Risk Measures: Entire Sample Univariate Estimates

The figures in the table have been obtained splitting the sample into two non-overlapping periods of equal length (96 months) and estimating in each subperiod risk measures for individual securities by running univariate regressions of securities returns on each macrovariable. Panel A shows the average value of the risk measures estimates of individual securities in each period. Panel B shows the correlation coefficients between the risk measure estimates of individual securities in each period. Panel C shows the proportion of individual securities for which the sign of the sensitivity to each factor changes from period 1 to period 2. Panel D reports the proportion of rejections of the stability hypothesis of the risk measures for the individual securities using a Wald Test. Data sources are described in the Appendix.

	Term Structure Spread (UTS)	Lira/US$ exchange rate (DLUSD)	Unexpected Production (UMP)	Unexpected change in oil price (UOILG)	Unexpected Inflation (UI)
			Panel A: Average Value of Estimates		
Period 1	0.28	0.22	-0.50	0.10	-1.16
Period 2	2.29	0.17	-0.38	-0.10	-3.51
			Panel B: Correlation of Estimates		
Period 1 and 2	7.8	28.3	-25.7	4.0	6.2
			Panel C: Proportion of Sign Reversals		
Period 1 and 2	31.6	29.1	30.4	70.8	37.9
			Panel D: Proportion of Rejection (Wald Test)		
Confidence level:					
1%	7.5	1.3	1.3	2.5	1.3
5%	20.2	1.3	3.8	6.3	2.5
10%	27.8	1.3	8.9	21.5	5.1

Table 12

Stability of the Risk Measures:
Entire Sample Multivariate Estimates on Odd and Even Months

The figures in the table have been obtained by splitting the sample into odd and even months and estimating in each subsample risk measures for individual securities by running the following multiple regression:

$$R_{i,t} = c_i + b_{i1} UTS_t + b_{i2} DLUSD_t + b_{i3} UMP_t + b_{i4} UOILG_t + b_{i5} UI_t + e_{it}$$

where R_{it} is the return on share i in month t, UMP is the lagged unexpected change in monthly industrial production, UI is unexpected inflation, UTS is the unexpected return on long bonds, UOILG is unexpected change in oil prices (in US dollars), DLUSD is the percentage change in the lira/US dollar spot exchange rate. Panel A shows the average value of the risk measure estimates of individual securities in each subsample. Panel B shows the correlation coefficients between the risk measure estimates of individual securities in each subsample. Panel C shows the proportion of individual securities for which the sign of the sensitivity to each factor changes from the odd-month subsample to the even-months subsample. Panel D reports the proportion of rejections of the stability hypothesis of the risk measures for the individual securities using a Wald Test. Data sources are described in the Appendix.

	Term Structure Spread (UTS)	Lira/US$ exchange rate (DLUSD)	Unexpected Production (UMP)	Unexpected change in oil price (UOILG)	Unexpected Inflation (UI)
			Panel A: Average Value of Estimates		
Odd months	1.97	0.17	-0.88	0.13	-1.75
Even months	1.54	0.35	-0.46	-0.06	-3.42
			Panel B: Correlation of Estimates		
Odd and even	45.5	13.7	19.4	14.1	12.6
			Panel C: Proportion of Sign Reversals		
Odd and even	15.1	44.3	25.1	58.3	29.1
			Panel D: Proportion of Rejection (Wald Test)		
Confidence level:					
1%	1.3	1.3	2.5	5.1	1.2
5%	1.3	6.3	3.8	11.4	7.5
10%	3.8	13.9	10.3	19.9	11.3

References

Ball, R. and Kothary, S., (1989), *Nonstationary Expected Returns: Implications for Tests of Market Efficiency and Serial Correlation in Returns*, “Journal of Financial Economics”, 14, pp. 51-74.

Bodo, G., Cividini, A. and Signorini, L.F., (1991), *Forecasting the Italian Industrial Production Index in Real Time*, “Journal of Forecasting”, 10, pp. 285-99.

Blume, M., (1971), *On the Assessment of Risk*, “Journal of Finance”, 26, pp. 1-10.

_________, (1975), *Betas and Their Regression Tendencies*, “Journal of Finance”, 30, pp.785-95.

Breeden, D., (1979), *An Intertemporal Asset Pricing Model with Stochastic Consumption and Investment Opportunities*, “Journal of Financial Economics”, 7, pp. 256-96.

Brooks, R., Faff, R. and Lee, J., (1992), *The Form of Time Variation of Systematic Risk: Some Australian Evidence*, “Applied Financial Economics”, 2, pp. 191-98.

Chan, K. C. and Chen, N., (1988), *An Unconditional Asset-Pricing Test and the Role of Firm Size as an Instrumental Variable for Risk,* “Journal of Finance”, 43, pp. 309-25.

__________________ and Hsieh, D., (1985), *An Exploratory Investigation of the Firm Size Effect,* “Journal of Financial Economics”, 14, pp. 451-71.

Chang, E. and Pinegard, M., (1990), *Stock Market Seasonals and Prespecified Multifactor Pricing Relations*, “Journal of Financial and Quantitative Analysis”, 25, 4, pp. 517-32.

Chen, N., Roll, R. and Ross, S., (1986a), *Economic Forces and The Stock Market: Testing The APT and Alternative Asset Pricing Theories*, Working Paper, 119, Graduate School of Business, University of Chicago.

______________________, (1986b), *Economic Forces and The Stock Market*, “Journal of Business”, 59, pp. 383-403.

Chen, S. and Jordan, B., (1993), *Some Empirical Tests in the Arbitrage Pricing Theory: Macrovariables vs. Derived Factors*, “Journal of Banking and Finance”, 17, pp. 65-89.

Chow, G. and Lin, A. L., (1971), *Best Linear Unbiased Interpolation, Distribution and Extrapolation of Time Series by Related Series*, “Review of Economic and Statistics”, 21, November.

Collins, D., Ledolter, J. and Rayburn, J., (1987), *Some Further Evidence on the Stochastic Properties of Systematic Risk*, “Journal of Business”, 60, 3, pp. 425-48.

Conway, D. and Reinganum, M., (1988), *Stable Factors in Security Returns: Identification Using Cross Validation*, "Journal of Business & Economic Statistics", 6, pp. 1-15.

Cornell, B., (1983), *Money Supply Announcements and Interest Rates: Another View*, “Journal of Business”, 56, pp. 1-23.

Connor, G. and Korajczyk, R., (1991), *The Attributes, Behavior, and Performance of U.S. Mutual Funds*, “Review of Quantitative Finance and Accounting", 1, pp. 5-26.

Dimson, E. and Marsh, P., (1983), *The Stability of UK Risk Measurement and the Problem of Thin Trading*, “Journal of Finance”, 3, pp. 753-83.

Eckbo, E., Masulis, R. and Norli, O., (2000), *Seasoned Public Offerings: Resolution of the ‘New Equity Issues Puzzle”*, Journal of Financial Economics, 56, 2, pp. 251-92.

________ and Smith, D., (1998), *The Conditional Performance of Insider Trades*, "Journal of Finance", 53, 2, pp. 467-98.

Elton, E., Gruber, M. and Blake, C., (1995), *Fundamental Economic Variables, Expected Returns, and Bond Funds Performance*, “The Journal of Finance”, 50, September, pp. 1229-56.

Eun, C. and Resnick, B., (1992), *Forecasting the Correlation Structure of Share Prices: A Test of New Models, "*Journal of Banking and Finance", 16, pp. 643-56.

Fabozzi, F. and Francis, J., (1978), *Beta as a Random Coefficient*, “Journal of Financial and Quantitative Analysis”, 13, 1, pp. 101-16.

Fama, E., (1981), *Stock Returns, Real Activity, Inflation, and Money*, "American Economic Review", 71, September, pp. 545-65.

________, (1991), *Efficient Capital Markets: II*, "The Journal of Finance", 46, December, pp. 1575-617.

________ and Schwert, G., (1977), *Asset Returns and Inflation*, “Journal of Financial Economics”, 5, pp. 115-46.

Ferson, W. and Harvey, C., (1991), *Time Variation of Economic Risk Premiums*, “Journal of Political Economy", 21, pp. 385-414.

Friedman, M., (1988), *Money and The Stock Market*, "Journal of Political Economy", 96, 21, pp. 221-45.

Goldenberg, D. and Robin, A., (1991), *The Arbitrage Pricing Theory and Cost of Capital Estimation: the Case of Electric Utilities*, “The Journal of Financial Research”, XIV, 3, p. 181-96.

Greene, W., (1993), *Econometric Analysis*, Second Edition, New York, Macmillan Publishing Company.

Hamao, Y., (1988), *An Empirical Examination of the Arbitrage Pricing Theory Using Japanese Data*, “Japan and the World Economy”, 1, pp. 45-61.

He, J. and Ng, L., (1994), *Economic Forces, Fundamental Variables, and Equity Returns*, “Journal of Business”, 1994, 67, 4, pp. 599-609.

Honda, Y. and Ohtani, K., (1986), *Modified Wald Tests in Test of Equality between Sets of Coefficients in Two Linear Regressions under Heteroschedasticity*, “The Manchester School”, 2, pp. 208-18.

Jensen, G., Mercer, J. and Johnson, R., (1996), *Business Conditions, Monetary Policy and Expected Security Returns*, Journal of Financial Economics, 40, pp. 213-37.

Keran, M., (1971), *Expectations, Money, and The Stock Market*, "Federal Reserve Bank of St. Louis Review" , 53, January, pp. 16-32.

Lehmann, B. and Modest, D., (1988*), The Empirical Foundations of The Arbitrage Pricing Theory,* "Journal of Financial Economics", 21, pp. 213-54.

Lucas, R., (1978), *Asset Prices in an Exchange Economy*, "Econometrica", 46, pp. 1429-545.

Marchetti, D. J. and Parigi, G., (1998), "*Energy Consumption, Survey Data and the Prediction of Industrial Production in Italy"*, Banca d'Italia, Temi di Discussione, 342, March.

McElroy, M. and Burmeister, E., (1988), *Arbitrage Pricing Theory as a Restricted Non-linear Multivariate Regression Model*, "Journal of Business & Economic Statistics", 6, 1, pp. 29-42.

McQueen, G. and Roley, V., (1990), *Stock Prices, News and Business Conditions,* NBER Working Paper, 3520, November.

Nelson, C., (1976), *Inflation and the Rates of Return on Common Stock*, "The Journal of Finance", 31, May, pp. 471-83.

Panetta, F., (1996), *The Factor Structure of The Italian Stock Market*, Ph.D. Dissertation, London Business School, London.

Roma, A. and Schlitzer, A., (1996), *The Determinants of Italian Stock Market Returns: Some Preliminary Evidence*, "Economic Notes", 25, 3, pp. 515-40.

Pearce, D. K. and Roley, V., (1983), *The Reaction of Stock Prices to Unanticipated Changes in Money: A Note*, "Journal of Finance", 38, pp. 1323-33.

Ross, S., (1976), *The Arbitrage Theory of Capital Asset Pricing*, "Journal of Economic Theory", 13, pp. 341-60.

_______, (1977), *Risk, Return and Arbitrage*, in Friend, I. and Bicksler, J., "Risk and Return in Finance", Cambridge, Massachusetts, Ballinger.

Schlitzer, G., (1993), *Metodi per la stima in tempo reale della produzione industriale: una riconsiderazione e un confronto*, Banca d'Italia, Temi di Discussione, 191, March.

Schwert, W., (1981), *The Adjustment of Stock Prices to Information about Inflation*, "Journal of Finance", 36, pp. 15-29.

Shanken, J. and Weinstein, M., (1987), *Macroeconomic Variables and Asset Pricing: Estimation and Tests*, Working Paper, University of Rochester.

Sharpe, W. and Cooper, G. M., (1972), *Risk-Return Classes of New York Stock Exchange Common Stocks, 1931-67*, "Financial Analyst Journal", 28, pp. 46-54.

Sunder, S., (1980), *Stationarity of Market Risk: Random Coefficient Tests for Individual Stocks,* "The Journal of Finance", 35, pp. 883-96.

Ulrich, T. and Watchel, P., (1981), *Market Response to the Weekly Money Supply Announcements in the 1970s*, "The Journal of Finance", 36, pp. 1063-72.

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