

The Pricing of Italian Equity Returns

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In this paper, we investigate the relationship between common risk factors and average returns for Italian stocks. Our research has identified the Italian stock market's economic variables by using the results from factor analyses and time series regressions.

We study several multi-factor models combining the relevant macro-economic variables with the mimicking equity portfolios SMB (small minus big) and HML (high minus low) proposed by Fama and French (1993). The key question we want to ask ourselves, is whether the influential role of the size and book-to-market equity factors in explaining average stock returns can stand up well when competing with some macroeconomic factors. In other words, do stock returns carry some risk premium that is independent of either the market return or the economic forces that underlie the common variation in returns?

Our empirical work estimates risk premiums using both traditional two-pass procedures and one-pass (full information) methodologies. We show that only the market index and variables linked to interest rate shifts are consistently priced in the Italian stock returns. The role of other factors, and in particular both the size and the price-to book ratio, are crucially dependent on the estimation procedure.

(J.E.L.: G11, G12).

1. Introduction

Asset pricing theories imply that the expected returns of securities are related to their sensitivity to changes in the state of the economy. In the capital asset pricing model (CAPM) of Sharpe (1964), Lintner (1965), Mossin (1966) and Black (1972), this sensitivity is measured by a unique economic state variable: the securities β 's coefficients with a mean-variance efficient market portfolio. Intertemporal models such as those of Merton (1973), Long (1974), Lucas (1978) and Breeden (1979) and the arbitrage pricing theory (APT) of

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Ross (1976) have given us the insight that a small number of economic state variables suffice to describe the relationship between average returns and systematic risks.

These prominent theoretical contributions have led to a voluminous empirical research, first on a single risk premium perspective and successively within a multi-factor framework. Many papers have explored the relationship between stock returns and some *fundamental*, where the selection of the economic state variables has been guided essentially either by intuition (see, for example, Chen *et al.*, 1986) or by their popularity among market participants (the most important early examples are the size effect of Banz (1981) and the p/e effect of Basu (1983)). The empirical asset pricing literature has reached a critical frontier: the identification of systematic risks became an empirical matter not explicitly linked to any theoretical model, and several contributions held that stock returns were very sensitive to variables constructed using the market price of firms and some accounting measures (e.g., earnings, book equity and leverage). The debate on the empirical validity of the CAPM has caught fire again thanks to a study by Fama and French (1992a), who conclude that:

- (i) the market index β shows a very weak relation with average returns from the period 1941–90;
- (ii) two easily measured variables, size and book-to-market equity, provide a simple and powerful characterization of the cross-section of average stock returns for the 1963–90 period.

The study by Fama and French has received a great deal of attention from the press and has been promptly challenged in some academic papers. Amihud *et al.* (1992) and Kothari *et al.* (1992) have found a significant positive relationship between average return and β , either by adopting different methodologies or controlling for some data selection biases. There is still room for β in the CAPM. Next, Fama and French (1993) go on and present, with a time-series approach, strong evidence for the role of their three-factor model for pricing stock returns.¹ The bottom line of the Fama–French contribution (see also their complementary study (Fama and French, 1992b)) is that factors related to size (SMB – small minus big) and to the book-to-market equity ratio (HML – high minus low) are probably proxying some fundamental risk linked to firms' profitability and growth. The disturbing fact is that these risks do not seem to be captured by stock market movements.

¹ He and Ng (1994), using US data, have found that the role of size subsumes stock's risk exposures associated with the Chen–Roll–Ross (CRR) factors and that the CRR multifactor model does not explain the book-to-market effect. More recently, in a follow-up study, Fama and French (1996) have shown that, except for the continuation of short-term returns (momentum strategies), the relationship between average returns and firm characteristics (e.g. market capitalization, earnings/price, cash flow/price, book-to-market equity, past sales growth, etc.) largely disappears when using a three-factor model.

Our study aims to shed some new light on which factors best account for the common movements in Italian equity returns. To this end, we extend the set of state variables used in earlier studies to include both macroeconomic series and equity factor portfolios that could more properly capture risks affecting the Italian stock market.²

We study several multi-factor models combining the relevant macroeconomic variables with the mimicking equity portfolios SMB and HML. The key question we want to ask ourselves, put forward also by Fama and French (1992a, p. 450; 1993, p. 55), is whether the influential role of the size and book-to-market equity factors in explaining average stock returns can stand up well when competing with some macroeconomic factors. In other words, do stock returns carry some risk premium that is independent of either the market return or the economic forces that underlie the common variation in returns?

We find two variables that significantly and systematically explain Italian average stock returns: the market index β and the spread between bank loan rates and long-term government bond rates. This result is robust across different econometric methods. On the contrary, the significance of other variables, such as oil prices and the size and book-to-market equity factors, seem related to the estimation procedure. Our results suggest that both size and book-to-market may have a role for pricing Italian stocks, but they do not subsume the relevance of other risk factors.

The remainder of the paper is structured as follows. In section 2, we outline our research project, describing data, methodologies, and motivating the most important choices of the empirical work. Section 3 presents results for time-series regression analyses, and section 4 illustrates asset-pricing results considering several cross-sectional models and econometric techniques. Section 5 concludes and highlights some implications of our study.

² Recent research on Italian stock returns that pursued similar analyses include Roma and Schlitzler (1996), Panetta (1997), Doria *et al.* (1998) and Mazzariello and Roma (1999). Furthermore, some international studies do not support pricing being determined by the CRR set of macroeconomic variables either in the USA or in other countries. Shanken and Weinstein (1990) reconsidered the CRR evidence for the US market; in this paper, only the pricing of industrial production is supported. Hamao (1988) found a significant pricing relationship in Japan only for expected inflation, a default premium and a term structure factor. Finally, we are aware of two European studies that did not find a significant pricing relationship between stock returns and the CRR set of macroeconomic variables: Martínez-Rubio (1989) for Spain and Poon-Taylor (1991) for the UK. See also Asprem (1989) for the time-series relationship between European stock indices and macroeconomic variables.

2. Data and Research Design

2.1. Security Returns

We use the stocks of firms listed on the Italian Stock Exchange during the thirteen-year period 1981–93. All our tests were conducted on individual security returns rather than on returns of groups of securities, as is often done in the empirical asset pricing literature. We were forced to make this choice by the small number of listed securities which distinguishes the Italian Stock Market. In the sample period, this number ranged from 169 to 308, but the number of stocks with complete time series returns is even smaller. If we had been working with well-diversified portfolios, the resulting number of assets would not have been sufficient to obtain reliable factor risk premium estimates. However, there is a more important consideration regarding this issue. Our experiment is of interest since it produces evidence which is not subject to the likely spurious results which would emerge from an arbitrary selection of securities to be included in a given portfolio: a point first raised a long time ago in Roll's critique (1977), and which was specifically addressed more recently in the interesting analysis by Lo and MacKinlay (1990). On the other hand, our analyses are potentially subject to the well-known risks of measurement errors in the overlapping β estimates (Black *et al.*, 1972; Fama and MacBeth, 1973). We consider this issue, and to control the consequent errors-in-variables (EIV) problem, we adopt econometric methods based on generalized least squares (GLS) and pooled cross-sections to improve the efficiency of the second-pass estimator.³ We adopt also a one-pass nonlinear procedure to deal with efficiency losses caused by factors' unobservability in the estimation of risk premiums.

2.2. Economic State Variables

We construct exogenous macrovariables that may influence future cash flows or the risk-adjusted discount rate of firms. Based on this premise, we parallel the choices in CRR and in other studies that have investigated similar issues for other markets. However, we are careful to consider also macrovariables that are more country specific, since we doubt that the

³ Shanken (1992) provides some support for the use of GLS methods to face the classical EIV problem. Furthermore, the GLS approach constitutes a viable alternative to standard portfolio grouping in asset pricing tests. Our modified two-pass method (see section 2.3) allows for time variation in the covariance matrix and it can also lessen survival effects; see Brown *et al.* (1992). Recently, Amihud *et al.* (1992) have used these methods in the context of either a one-factor or a multi-factor model.

same set of macrovariables may be valid for capital markets with rather different economies. Another distinction of our research is that it considers equity-related variables which mimic two of the most frequently cited stock market risk premia: size and book-to-market equity. Our goal is to examine their role in competition with macroeconomic variables in the pricing of Italian equities. While, in the Appendix, we provide a detailed description and discussion of our data, we shall now briefly present their definitions.

- MPS is the change in industrial production calculated from the monthly Industrial Production Index.
- UI is the unanticipated inflation constructed from a time series analysis of the consumer price index (CPI).
- Foreign exchange risk premia are represented by two variables: DDM is the Italian lira volatility with respect to the cross rate lira/DM in the EMS central parity and DTT are changes in the terms of trade.
- OG is the change in oil prices.
- UPR is a default risk variable approximated by the spread between average bank loan rates and long-term government bond rates.
- Interest rate related variables are the changes in the difference between long- and short-term rates for government securities (the term structure spread UTS), and changes in the official discount rate (DTUS).
- One variable is inserted to capture the role of consumption: CG, which is the change in real consumption.

On the side of equity related variables we analyse market indices and mimicking portfolios for size and book-to-market equity risk factors.

- VW and EW are, respectively, the returns of the value-weighted and the equally weighted Italian stock market indexes.
- The effects of co-movements between the Italian market and foreign stock markets are considered by including the changes in the Morgan Stanley Capital International World Equity index (MSCI), including dividends.
- SMB is the factor portfolio that mimics the spread returns between small and big firms, controlling for the book-to-market equity risk.
- HML is the factor portfolio that mimics the spread returns between high and low book-to-market equity firms, controlling for the size risk.
- SMB and HML are the Italian equivalent, for the 1981–93 time period, of the US market factors constructed by Fama and French (1993, pp. 8–9).

2.3. Econometric Approach to Cross-sectional Analyses

We approach the estimation of the average return- β s relationship in three ways. In the first two, we run usual two-pass procedures and pooled time-series and cross-sectional regressions (CSR); in both cases, we present GLS estimates. With the third approach, we run a one-pass multivariate regression model with across-equations restrictions.

Two-step Procedures

We adopt two strategies to run CSR with monthly data: a standard Fama–MacBeth (1973) type and the feasible estimator procedure developed by Amihud *et al.* (1992). In both cases, only generalized least squares (GLS) estimations are performed. Standard ordinary least squares (OLS) assumptions on the cross-sectional residuals are frequently violated, and, in our study, these biases may be particularly severe, given that we use individual stock returns.

Nonlinear One-pass Procedure

We follow the approach suggested in Burmeister and McElroy (1988) and McElroy and Burmeister (1988).

If we assume that a linear factor model with k factors generates asset returns

(1)

$$R_i(t) = E_t[R_i(t)] + \sum_{j=1}^k b_{i,j} f_j(t) + \varepsilon_i(t) \quad \text{for } i = 1, \dots, n; t = 1, \dots, T$$

and accept the usual assumption of most APT models that the expected return is approximately given by

$$(2) \quad E_t[R_i(t)] = \lambda_0(t) + \sum_{j=1}^k b_{i,j} \lambda_j(t) \quad \text{for } i = 1, \dots, n$$

we can bypass the usual econometric difficulties associated with the standard two-pass procedures estimating (1) as a multivariate nonlinear regression model.

With the substitution of (2) in (1), we obtain a system of nonlinear equations for, say, a group of $N < n$ assets over T time periods

$$(3) \quad R_i(t) = \lambda_0(t) + \sum_{j=1}^k b_{i,j}(t) \lambda_j(t) + \sum_{j=1}^k b_{i,j}(t) f_j(t) + \varepsilon_i(t)$$

for $i = 1, \dots, N; t = 1, \dots, T$

where the term

$$\sum_{j=1}^k b_{i,j} \lambda_j$$

represents the $N - K$ nonlinear across-equation restrictions and $\lambda_0(t)$ has the general form

$$\lambda_0 = \alpha_0 \iota_T + \lambda_0^\bullet$$

with λ_0^\bullet an observed proxy for the risk-free rate. We can rewrite (3) as

$$(4) \quad r_i(t) = \alpha_0 + \sum_{j=1}^k b_{i,j} \lambda_j + \sum_{j=1}^k b_{i,j} f_j(t) + \varepsilon_i(t)$$

for $i = 1, \dots, N; t = 1, \dots, T$

where $r_i(t)$ is the excess return.

The estimation of system (4) is obtained using ITNL3SLS (iterated nonlinear three stages least squares). NL3SLS are, even in the absence of normality of the disturbances, consistent and asymptotically normal (Gallant, 1987), hence they offer robust estimators and tests even if the assumption of normality is incorrect. Under mild conditions, the ITNL3SLS estimators achieve the Cramer–Rao lower bound.

2.4. Basic Statistics of Economic State Variables

Summary statistics for the variables are presented in Tables 1 and 2. Looking at the correlations (Table 1) with the equally weight (EW) stock market index, we note a significantly positive coefficient for three variables: the terms of trade (DTT), the default premium (UPR), and the MSCI index. With a significant negative coefficient, we find instead four variables: oil prices (OG), the official discount rate (DTUS), the term structure spread (UTS) and the size related variable (SMB). Using the VW market index, we reach the same conclusions. Thus, a first look at the aggregate serial-correlations suggests that the Italian equity indices do not seem related to variables such as industrial production, unexpected inflation, consumption growth or the book-to-market equity factor (HML) proposed by Fama and French (1993). Industrial production is not significantly correlated with any variable. For the unexpected inflation variable (UI), we see a weak positive coefficient with oil prices (OG). Consumption growth (CG) is negatively correlated with the default factor (UPR). The correlation between the two factors SMB and HML is nil (0.007); a measure equivalent – from a purely statistical point of view – to that estimated by Fama and French (1993, p. 9) in the period 1963–91 for

Table 1: Correlation of the Economic State Variables

Variable	EW	VW	DTT	OG	UPR	DTUS	UI	UTS	DDM	MPS	CG	MSCI	SMB	HML
EW	1.000													
VW	0.960	1.000												
DTT	0.239	0.274	1.000											
OG	-0.228	-0.240	-0.505	1.000										
UPR	0.290	0.306	-0.014	0.007	1.000									
DTUS	-0.187	-0.199	-0.033	0.067	-0.189	1.000								
UI	-0.044	-0.063	-0.108	0.150	-0.021	-0.026	1.000							
UTS	-0.263	-0.275	-0.126	0.131	-0.663	0.229	0.119	1.000						
DDM	-0.136	-0.098	-0.074	0.112	-0.094	0.133	-0.044	0.155	1.000					
MPS	-0.107	-0.110	0.064	-0.010	0.004	0.042	0.121	-0.013	0.072	1.000				
CG	0.074	0.021	-0.042	0.052	-0.165	0.061	0.027	0.051	0.088	0.062	1.000			
MSCI	0.442	0.432	0.040	-0.073	0.257	-0.102	0.030	-0.142	-0.003	-0.065	-0.099	1.000		
SMB	-0.167	-0.378	-0.176	0.151	-0.121	0.019	0.070	0.064	-0.048	0.063	0.219	-0.171	1.000	
HML	0.100	-0.016	-0.081	0.037	-0.107	-0.029	0.080	0.028	-0.102	0.035	-0.034	0.090	0.007	1.000

Table 2: Summary Statistics for the Economic State Variables (Monthly data: January 1981–December 1993; 156 observations)

State variable	Symbol	Mean ^a	Std. Dev.	ρ_1	ρ_2	ρ_3	ρ_{12}	ρ_{24}
Equally weighted market index	EW	1.27	0.065	0.187	0.043	0.126	0.069	0.004
Value weighted market index	VW	1.16	0.076	0.130	0.006	0.109	0.056	0.005
Change in the terms of trade	DTT	0.17	0.018	-0.048	-0.081	-0.064	0.191	0.175
Change in price of oil	OG	0.29	0.094	0.254	-0.015	-0.013	-0.017	0.014
Spread between the average bank loan rates and the average long term government bond rates	UPR	3.44	0.009	0.897	0.752	0.624	0.316	-0.031
Change in the official discount rate	DTUS	-0.05	0.004	0.009	0.180	-0.035	-0.041	0.003
Unanticipated inflation	UI	-0.01	0.002	0.090	0.049	-0.107	-0.107	0.198
Term structure spread	UTS	0.28	0.008	0.616	0.493	0.382	0.398	-0.020
Change in the lira-DM	DDM	-0.06	0.062	-0.039	-0.011	-0.024	-0.031	-0.138
Change in the monthly industrial production	MPS	0.10	0.028	-0.494	0.003	-0.034	0.039	-0.163
Change in real consumption	CG	0.18	0.002	0.543	0.453	0.238	0.105	0.205
World market return	MSCI	1.36	0.048	0.109	0.064	-0.044	-0.060	0.050
Change in the size equity factor	SMB	-0.34	0.026	0.070	0.097	-0.078	0.035	-0.005
Change in the book-to-market equity factor	HML	0.07	0.034	0.123	0.180	0.046	0.121	-0.144

Notes: ^apercentage per month

the US market. Furthermore, looking at other sample period summary statistics for these two factors, we observe a similarity with US data reported in Fama–French (1992b, table 2) for the period January 1982–December 1991.

Closer looks at the cross-correlations among the variables most linked with the stock market tell us that DTT has a high negative correlation with OG. This is not surprising. For example, a depreciation of the Italian lira against the US dollar would coincide with an oil price increase in domestic terms and, in addition, deterioration in the terms of trade. Another strong correlation is that between UPR and UTS. This is to be expected because of the presence in both of them of the government long-term bond series LGB. Both factors, SMB and HML do not seem strongly correlated with any macroeconomic series, although some weak correlation does indeed show up. A number of other correlations are not negligible, but the variables clearly are far from perfectly correlated and no one variable can be substituted for any other.

Our overall valuation from correlation analysis is that forces that are closely linked to the economic system drive the Italian stock market. Importing and exporting are core activities of the Italian economy; therefore the domestic capital market will reflect this main activity through the variables which will capture its underlying risks for Italian firms.⁴

Table 2 reports the mean, standard deviation, autocorrelations and Pearson correlation coefficients among the economic state variables. Autocorrelation and cross-correlation coefficients outside the interval ± 1.60 (± 0.240) are significantly different from zero at the 5% (1%) level. The variables generally display a moderate autocorrelation; however many of them have seasonal at the first month lag. We note – as expected – strong autocorrelations for UPR and UTS. Given the well-known problems that autocorrelations may induce in the estimation of the loadings of stock returns on these variables, we also took account in our testing of the relationship between stock returns and macroeconomic series with the prewhitened variables. In other words, we constructed the residual series of different ARIMA order combination (p, d, q) models based on full sample data, which represent the unanticipated components of state variables.⁵

⁴ To understand the importance of international trade for the Italian economy, it may be useful to consider the average incidence of exports and imports to GNP in our sample period (1981–93) for Italy and, in comparison, for the US and Japan. The ratio of exports to GNP is 21.7 per cent for Italy and 9.3 per cent and 13.6 per cent respectively for the US and Japan. The ratio of imports to GNP is 24.1 per cent for Italy and 10.2 per cent and 12.7 per cent respectively for the US and Japan.

⁵ The prewhitening process was carried out by fitting a univariate ARIMA model to each series. Looking at the series correlogram and partial correlogram, we identified the process which transforms the residual series into white noise. A first-order MA term was required for EW, OG, MSCI, a second-order autoregressive process was required for HML, an AR(12) was fitted for DTT, the UPR and UTS series producing, respectively, an ARIMA (0,1,1) and an ARIMA (1,1,12). We also tested the null hypothesis that all variables are white noise with a Ljung-Box test, and we could not reject the hypothesis at standard significance level in all cases.

3. Empirical Results

We started our empirical work exploring the time series relationship between mimicking portfolios constructed from factor analysis and a subset of economic state variables chosen on the basis of serial correlation significance with the market index.⁶

In this way, we were able to gain some idea as to which economic variables are more relevant to explain average stock returns. For the whole period of thirteen years, we observed the significance of four variables: the market index, oil prices, and the two portfolios related to the size and book-to-market risk factors. This conclusion seems quite consistent also in the two sub-periods of six years. Only oil prices show a weaker time series relationship in more recent years. Although we have identified a parsimonious set of economic variables that seem to capturing the systematic co-movement of Italian stock returns, those factors, however, may not necessarily coincide with the factors that may help to explain the behaviour of expected returns (i.e. they are not priced).

To perform pricing analyses, we construct a sample of monthly stock returns consisting of listed securities on the Database file for which no returns were missing from January 1981 through December 1993. The resulting number of stocks included in this sample was 83. The market value of the sample as a percentage of the total market capitalization in the sample period was, on average, 64 per cent. Their monthly returns are used to run the pricing model against the identified set of economic state variables.

3.1. Two-pass Procedures

In Table 3, we report the pricing results for two different approaches and set of variables: raw economic state variables, and pre-whitened state variables.⁷ Table 3 reports estimated coefficients (in percentage per month) from monthly cross-sectional regressions of individual stock excess returns on economic state variables. In the first pass, five years of monthly returns are used to estimate β s and in the second-pass the returns of the next month are used to run the following cross-sectional model:

⁶ To perform factor analysis, we followed previous approaches suggested in past studies; see, for example, Lehmann and Modest (1988). We used individual weekly stock returns over eight overlapping five-year periods from 1981 to 1993 and we prespecify at five the number of factors, given the evidence of previous research which looked into this issue; see, for example, Roll and Ross (1980), Brown and Weinstein (1983) and Panetta (1997). Weekly returns of mimicking portfolios are then compounded to produce monthly returns that will be regressed against the observed economic variables.

⁷ To perform timing consistent econometric analyses, raw economic variables that are released with some time lag are aligned with end-of-month stock returns. Specifically, we lagged DTT by 2 months. Furthermore it has to be noticed that pre-whitened variables are proxies for true innovations since we compute in-sample residuals.

Table 3: Pricing Results in the Italian Stock Market: 1986–93 (Monthly cross-sectional regressions)

Estimation method		$\hat{\gamma}_0$	$\hat{\gamma}_1$	$\hat{\gamma}_2$	$\hat{\gamma}_3$	$\hat{\gamma}_4$	$\hat{\gamma}_5$	$\hat{\gamma}_6$
Raw economic state variables								
Fama–MacBeth	GLS	−0.506 (−1.08)	0.283 (0.38)	−0.069 (−0.74)	0.085 (0.44)	0.463 (0.46)	−0.213 (−0.99)	0.153 (0.53)
	OLS	−1.032 (−2.65)	0.861 (2.36)	−0.075 (−0.90)	0.160 (0.99)	−0.196 (−0.33)	−0.050 (−0.27)	0.451 (2.01)
Pooling time-series		[−2.50]	[2.07]	[−0.73]	[0.85]	[−0.26]	[−0.25]	[1.78]
Cross section	GLS	−0.198 (−1.59)	0.821 (4.40)	−0.089 (−2.82)	0.121 (1.63)	−0.042 (−0.16)	−0.190 (−1.33)	0.357 (2.70)
Pre-whitened state variables								
Fama–MacBeth	GLS	−0.499 (−1.18)	0.274 (0.35)	−0.062 (−1.36)	−0.039 (−0.21)	0.443 (0.41)	−0.183 (−0.68)	−0.013 (0.04)
	OLS	−1.279 (−3.35)	1.153 (3.22)	−0.080 (−2.52)	0.098 (0.63)	−0.271 (−0.46)	0.092 (0.48)	−0.117 (−0.49)
Pooling time-series		[−3.19]	[2.72]	[−2.33]	[0.55]	[−0.36]	[0.42]	[−0.42]
Cross section	GLS	−0.127 (−1.04)	1.183 (6.37)	−0.073 (−6.17)	0.095 (1.27)	−0.040 (−1.53)	0.035 (0.39)	−0.112 (−0.85)

Notes: The t-values in parentheses are for the estimated coefficient values from zero, in brackets the t-stats using a Newey-West (1987) correction for heteroscedasticity.

$$r_{it} = \gamma_0 + \gamma_1 \hat{\beta}_{EW} + \gamma_2 \hat{\beta}_{UPR} + \gamma_3 \hat{\beta}_{DTT} + \gamma_4 \hat{\beta}_{OG} + \gamma_5 \hat{\beta}_{SMB} + \gamma_6 \hat{\beta}_{HML} + \varepsilon_{it}$$

$$\text{for } i = 1, \dots, 83; t = 1, \dots, 96$$

where r_{it} is the stock i return in month t in excess of the Italian 6-months T-bill monthly rate.

First, we ran OLS estimations either with the Fama–MacBeth methodology or the pooled method using raw economic state variables, finding very poor results. Using GLS estimations, the results show again no significant results with Fama–MacBeth. On the contrary, GLS estimations with the pooled regressions indicate a significant result with the following β s: the market index, UPR and HML. However, the estimated premium on the terms of trade (DTT) variable is also significant, although only at 10% level. The market index premium is quite high, with a mean of 0.8 per cent per month. Another significant result is the monthly premium for HML: the estimated coefficient is positive and about 0.4 per cent. Taken together, they account for about 1.2 per cent per month, while the other significant coefficients are trivial. The estimated premium for UPR, even if it is large in statistical terms is on average only 0.1 per cent per month. The terms of trade (DTT) enter with a positive coefficient, significant at only the 10% level but again with a small monthly premium.

The results for the stock market index are not consistent with the findings of studies using multifactor models; see, for example, Chen *et al.* (1986) and, more recently, Ferson and Harvey (1991) for the US market and Hamao (1988) for Japan.

The strong significance for pricing of the book-to-market equity factor confirms its explanatory power in time series regressions, but it is puzzling. Fama and French (1992b, 1993) maintain that a book-to-market equity ratio has strong roots in firms' profitability and growth. If their explanation is reasonable, then how come the HML underlying risk is not already contained and captured by the most relevant macroeconomic variables? This result is new in literature, and it may be the outcome of the higher precision with which mimicking portfolio stock returns, such as HML, capture some aspects of firm profitability and growth compared to the macroeconomic variables. Using pre-whitened variables (lower half of Table 3), we confirm most of the previous results. Only for HML, we found conflicting evidence. Summing up, our evidence demonstrates that other common risk factors may not be subsumed in any equity-derived factor, and they have to be identified among the observed macroeconomic variables.

However, these results are based on asymptotic standard errors for the coefficient estimates, which may not be accurate given our small sample. As two-pass procedures may not be very efficient, we turned to nonlinear one-pass procedures to reassess our results. Using the one-pass methodology, the small sample properties of the estimator may be easily assessed by simulation.

3.2. Nonlinear One-pass Procedure

In Table 4, we present the results for the multivariate non-linear regression model (4). Also, with this model, we have been using both raw and pre-whitened state variables. Table 4 reports the ITNL3SLS estimates of the factor premiums of individual monthly stock excess returns on the innovations of economic state variables. The innovations are obtained subtracting the sample mean from the variable time series. The model estimated is the following:

$$r_{it} = \lambda_0 + \lambda_1\beta_{EW} + \lambda_2\beta_{UPR} + \lambda_3\beta_{DTT} + \lambda_4\beta_{OG} + \lambda_5\beta_{SMB} + \lambda_6\beta_{HML} + \beta_1EW \\ + \beta_2UPR + \beta_3DTT + \beta_4OG + \beta_5SMG + \beta_6HML + \varepsilon_{it} \\ \text{for } i = 1, \dots, 83; t = 1, \dots, 156$$

where r_{it} is the stock I return in month t in excess of the Italian 6-month T-bill monthly rate.

Table 4: Pricing Results in the Italian Stock Market ITNL3SLS (January 1981–December 1993)

Parameter	Estimate (percentage per month)	T-stat (p -value)	Bootstrap simulations T-stat (p -value)
Raw economic state variables			
λ_1	0.467	1.83 (0.069)	1.69 (0.092)
λ_2	0.653	6.50 (< 0.001)	4.08 (< 0.001)
λ_3	-0.237	-1.38 (0.169)	-1.14 (0.256)
λ_4	-1.545	-1.91 (0.058)	-1.55 (< 0.123)
λ_5	-0.665	-8.60 (< 0.001)	-5.39 (< 0.001)
λ_6	-0.300	-1.72 (0.087)	-1.49 (0.138)
Pre-whitened economic state variables			
λ_1	0.256	4.07 (< 0.001)	3.07 (0.002)
λ_2	0.203	3.48 (0.001)	2.62 (0.009)
λ_3	-0.642	-2.83 (0.005)	-1.91 (0.058)
λ_4	-1.147	-4.89 (< 0.001)	-3.80 (< 0.001)
λ_5	-0.619	-5.16 (< 0.001)	-3.65 (< 0.001)
λ_6	-0.749	-3.78 (< 0.001)	-2.96 (0.003)

Notes: Based on $R = 1000$ simulations

The joint time-series cross-sectional estimates are computed using ITNL3SLS. Five economic factors are significantly priced and they account for 40–50 per cent of the variability of individual assets. There are 505 parameters to estimate: of the 498 β s, 186 are significant at the 10% level. The number of degrees of freedom for each equation, available to estimate the six λ s are $(NT - NK - K)/N = 149.9$. Within the raw economic state variables, table 4 reports two positive risk premium estimates, both highly significant: the premium for the stock index is of the order of 0.47 per cent per month; analogously the premium for UPR is on average 0.65 per cent per month. The remaining coefficients are negative; the discount associated with oil prices (OG) is -1.55 per cent per month, even if this risk premium is marginally significant at the 5% level as in the case of the stock index. The estimated λ for SMB is highly significant and negative with an average discount of -0.67 per cent per month. The discount for the book-to-market equity factor is about -0.3 per cent per month. Implications from the reported results are that both economy-wide factors and equity related variables are jointly necessary for the pricing of Italian stocks. The size and book-to-market factors retain their importance when used in combination with other macroeconomic variables but, at the same time, factors such as default risk (UPR) and oil prices are significant, implying that their effect is not subsumed by SMB and HML. Last but not least, we find that even accounting for all these effects, the market portfolio is an essential constituent for a complete story of asset returns.

In the case of pre-whitened variables, risk premium estimates retain their sign, but now all of them are highly significant. The differences between the two sets of results – raw versus pre-whitened state variables – may be attributed to the spurious effects of time series autocorrelation.

To analyse the sensitivity of the results to the variables included in our model, we also ran a different model with the term structure (UTS) instead of UPR. Given the high negative correlation between the two variables we expect the substitution to be almost neutral for our findings. Indeed, the estimated risk premium for UTS is -0.647 per cent congruent in size and sign with the results given in Table 4 for UPR. Analogously, given the high contemporaneous correlation between DTT and OG, we ran two five-factor models including only one of the above but, even in this case, we find that the estimates are consistent with the results of the complete six-factors model with the risk premium for OG retaining its significance at the 10% level.

As a further check for robustness, we ran a non-parametric bootstrap procedure on the estimated risk premia to ascertain whether our results are in any event linked to small sample problems. Using 1000 bootstrap samples, we estimated for each one the λ s; see Mazzariello–Roma (1999) for a previous application of this technique to Italian data. *T*-statistics and *p*-value resulting from bootstrap procedure are reported in the last column of Table 4. Indeed, as it can be seen, pricing results are similar, and we only observe a decrease in the level of significance.

To summarize, we believe that results from one-pass procedure are more reliable than GLS two-pass procedure, providing that the higher efficiency and power of the first can overcome the problems related to the choice of sample size, period and state variables which has to be faced anyway. Future work may be directed toward this direction, given the relevant practical implications of precise estimations of risk-premiums.

As a last exercise, we show, in Table 5, the results of the cross-sectional regressions between the mean observed excess return for the 83 stocks for the overall period 1981–93 and the expected return obtained as a by-product of the multivariate nonlinear estimates.

The model estimated is the following:

$$r_i = a + bER + e_i \quad \text{for } i = 1, \dots, 83$$

where r_i is the stock i mean monthly return in excess of the Italian 6-month T-bill monthly rate, and ER is the expected return obtained from the ITNL3SLS estimates either for each factor or for the complete six factor models. The expected return used as regressor is given by

$$E(r_i) = \alpha_0 + \sum_{j=1}^k b_{i,j} \lambda_j$$

In the first line, we present the results using the complete model with six factors; the following lines show what is the contribution of each individual factor when we use as a regressor the expected return calculated from a one-factor-at-the-time ITNL3SLS multivariate regression. The in-sample explanatory power of the six-factors model is quite high. Together, they account for

Table 5: The Relationship between Average and Expected Returns (January 1981–December 1993)

Factor used to estimate expected returns	\hat{a}	\hat{b}	\bar{R}^2
All six factors	-0.000 (-0.63)	0.95 (13.43)	0.69
EW	-0.000 (-0.11)	1.10 (2.82)	0.08
UPR	-0.002 (-2.62)	0.92 (9.46)	0.52
DTT	0.001 (1.04)	1.05 (5.79)	0.28
OG	0.001 (0.60)	1.07 (4.92)	0.22
SMB	0.000 (0.60)	1.00 (6.05)	0.30
HML	0.003 (3.46)	1.05 (3.38)	0.11

Notes: The t -values in parentheses are for the estimated coefficient values from zero.

about 70 per cent of the average returns cross-sectional variability. However, the evidence of the next lines is even more interesting as we can, in this way, disentangle the individual contribution of each risk factor to the above result. What we find is that the spread between bank loans rates and government bond rates is, by and large, the single most important explanatory factor accounting for about 52 per cent of the variability in the observed returns. SMB, DTT and OG have individually an high explanatory power ranging from 22 to 30 per cent as can be seen from the adjusted R^2 . The results for EW and HML are much less strong, the adjusted R^2 in these cases is 8 per cent and 11 per cent, respectively.

4. Conclusions and Implications

This paper searched economic state variables and investigated pricing effects as systematic risks using Italian data. We extended the set of economic variables analysed in previous national and international studies to include both economy-wide factors and equity derived factors, similar to those studied in Fama and French (1993) for the US market, which were intended to capture a size (SMB) and a book-to-market equity risk (HML).

We found that changes in the market index, changes in oil prices, a default premium, changes in interest rates and the SMB and HML factors represent a good summary of the risks captured by the cross-section of average Italian returns. Our findings revealed that the economic risk premium associated with the size and book-to-market equity variables are priced even in combination with macroeconomic factors.

Results for the Italian market reveal that both macroeconomic variables and equity risk factors are relevant for pricing stock returns. This is not surprising for an open economy such as the Italian one, but our paper provides evidence that differs substantially from the well-known results for the US stock market.

Our research has some practical implications in the area of management of Italian stock portfolios. First, our results may be used to implement either a passive strategy, which essentially sterilizes portfolios from excessive exposure to the type of systematic risks we uncover in our research. Second, it can be also used to pursue an active strategy in which a money manager makes bets based on forecasts of risk factor realizations.

However, several caveats are worth mentioning. First, it could be advisable to extend our research to a longer period of time so as to obtain more reliable results, particularly the stability of factor risk premia. Second, although we could not find any relationship between the local market index and the world global index, a future extension of our work should also consider the increasing integration of the Italian market in the world stock markets. We feel these extensions may improve the estimated factor model, and can help to have a better performance, both for predicting returns and for controlling portfolio risk.

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APPENDIX

This appendix provides the details on the data we used.

Macroeconomic Variables

Industrial Production

Monthly growth rate of industrial production calculated from the monthly Industrial Production Index is used as basic series. The index, compiled by the Italian Central Statistics Office (ISTAT), is adjusted for the number of working days in each month and seasonally adjusted. The index is made known from the ISTAT with a lag of two months. If IP_t denotes the index of industrial production in month t , then the monthly growth rate is

$$MPS_t = \frac{IP_t - IP_{t-1}}{IP_{t-1}}$$

Inflation

Only the unanticipated component of inflation is used in pricing. Unanticipated inflation is defined as

$$UI_t = I_t - E[I_t | t - 1]$$

where I_t is the realized percentage variation of the consumer price index (CPI) for month t and $E[I_t | t - 1]$ is the series for expected inflation. The variation in the CPI for month t is published in month $(t + 1)$.

We used two different methodologies to derive the expected inflation variable. First, we followed the procedure found in Fama and Gibbons (1984) as well as in Chen *et al.* (1986). Then, we constructed $E[I_t|t-1]$ from a time-series analysis of the I_t variable. Both the resulting series for UI_t were examined for their influence on pricing, but only the second series resulted significantly different from zero in this study. A second variable that is unanticipated and might have an influence which is distinct from UI_t , is the change in expected inflation

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

Also, the change in expected inflation was derived following the two different procedures mentioned above to calculate the expected inflation variable.

Oil Prices

Italy is heavily dependent on foreign oil. The oil price is included as a systematic factor, since we believe it may proxy as one of the most important production costs. The oil price growth variable, OG_t , is constructed as the monthly percentage variation in the Producer Price Index/Crude Petroleum series obtained from the Bureau of Labor Statistics (US Department of Labor). If $POIL_t$ is the producer price crude petroleum index in month t , then the variable employed is

$$OG_t = \frac{POIL_t - POIL_{t-1}}{POIL_{t-1}}$$

Consumption

To our set of macroeconomic variables, we add the time series of percentage changes in real consumption, CG_t . However, due to the lack in Italy of such a monthly series, we derived CG_t by a simple linear interpolation of the Quarterly National Accounts seasonally adjusted series (at constant prices) produced by ISTAT.

Financial Market Variables

Interest Rates

The effects of term structure variation on the stock market are proxied by the following interest rates difference:

$$UTS_t = LGB_t - TB_{t-1}$$

where LGB_t is the long-term interest rate approximated by the yield on a Government long term bond (Italian BTP) and TB_t is the short-term (3-months) Government interest rate (Italian BOT).

To capture the effects of a change in monetary policy, another variable drawn from the money markets is considered. The variable $DTUS_t$ is defined as

$$DTUS_t = TUS_t - TUS_{t-1}$$

where TUS_t is the Italian Official Discount Rate.

A measure for investors' required return for accepting risk was determined in Chen *et al.* (1986) by calculating the difference between high-grade and low-grade bond returns. Unfortunately, there is no reliable time series data in Italy on corporate bond ratings and returns. One variable we might look at instead, is the spread between interest rates on bank loans and long-term government bonds. The variable, UPR_t , is defined as

$$UPR_t = TIP_t - LGB_t$$

where TIP_t is the average interest rate on bank loans published monthly by the Bank of Italy.

Exchange Rates

The Italian economy is heavily oriented toward international trade and one might suspect that volatility in foreign exchange rates would have a substantial systematic effect upon Italian equities. First, we employ a variable that might capture the Italian lira's volatility in the EMS foreign exchange market. If $LITDM_t$ is the average spot exchange rate (lira/DM) in month t and $PARLITDM_t$ is the EMS central parity for the cross rate lira/DM, then the variable used is defined as

$$DDM_t = \frac{DM_t - DM_{t-1}}{DM_{t-1}}$$

where

$$DM_t = \frac{LITDM_t}{PARLITDM_t}$$

Second, we use the changes in the terms of trade as a macroeconomic series which may capture more directly the role of real exchange rate shocks to the Italian stock market. Terms of trade data, TT_t , is obtained from ISTAT monthly statistics and is defined as the ratio between the price index of export goods and the price index of import goods. Even in this case, the relevant indexes are

published with a two months lag. The variable employed is the change in the terms of trade expressed as

$$DTT_t = \frac{TT_t - TT_{t-1}}{TT_{t-1}}$$

International Stock Market Index

Even casual observations of the changes in the stock market indexes reveal that stock markets across the world are correlated. Therefore, we thought to examine this relationship by employing the monthly percentage variations in the Morgan Stanley Capital International World Equity Index (including dividends). We then converted the US dollar returns series into the Italian lire returns series equivalent.

Stock Returns

Italian stock prices are retrieved from tapes (DRG) of the Computer Center (CEDBORSA) of the Italian Stock Exchange (ISE) and from yearly publications of the Statistical Division of the ISE.

We used both weekly and monthly returns in our empirical analyses. Weekly returns are calculated as the percentage change from the last day-week price with respect to the first day-week price; monthly returns are calculated as the percentage change from the price of the last day of month t with respect to the last day of month $t - 1$. Stock returns are adjusted for splits and right issues, and include cash dividends.

We employ two broad market indexes in our work. The first is a value-weighted (VW) index calculated by the Statistical Division of the ISE, known as the MIB Storico. The second is the MISE – Milan Stock Exchange Equally – an equally weighted (EW) index calculated by the authors. Besides the different weights used by the two indexes, EW is an index that includes cash dividends. However, we conducted tests like those proposed by Gibbons *et al.* (1989) to verify the mean-variance efficiency of the two market indexes during the same time period of this work. We strongly reject the efficiency of the VW, while the efficiency of the EW is accepted most of the time at standard significance levels (details and results are available from the authors).

Other Economic Fundamentals

The research database also includes accounting and other corporate data, which we use to construct the variables SIZE, BE/ME, SMB, and HML. Raw data for these variables come from the yearly publications of MEDIOBANCA,

Le principali società italiane and *Indici e Dati*. ME is the market value of equity (stock price times shares outstanding), and BE is the book value of a firm's stock. Most of the listed firms release their accounting data before June.

We follow the procedure of Fama and French (1993, pp. 8–9) in constructing the SMB and HML mimicking portfolios. However, further details are worth noting, given the rather different institutional structures in Italian markets and US.

In June of each year t from 1980 to 1993, all ISE stocks are ranked on ME and equally divided into two groups (small and big). We used only stocks which had prices for December of year $t - 1$ and June of year t and book common equity for year $t - 1$. Simultaneously, we also rank stocks in the three BE/ME groups (low 30%, medium 40%, high 30%). We constructed six portfolios (SL, SM, SH, BL, BM, BH) intersecting the two ME groups and the three BE/ME groups. Monthly returns on the six portfolios from July of year t to June of year $t + 1$ were calculated as value-weighted returns. Each portfolio was reconstructed starting from June of year $t + 1$. In Italy, most of the listed firms adopt a December-to-December fiscal year. Some firms adopt a different period (for example, from June to June). We thought about a different strategy to form portfolios, deciding to move the starting date either from September of each year t or from December of each year $t - 1$.

Results and portfolio characteristics remained essentially unchanged, and therefore, we decided to use the two mimicking portfolios which parallel the Fama–French methodology. The final time series of monthly returns of the SMB portfolio is then calculated as the difference between the average of the three small-stock portfolios (SL, SM, SH) and the average of the three big-stock portfolios (BL, BM, BH). Monthly returns of the HML portfolio are the difference between the average on the two high BE/ME portfolios (SH, BH) and the average on the two low BE/ME portfolios (SL, BL).

Non-technical Summary

Past research has identified many patterns in average equity returns. Within the framework of the arbitrage pricing theory (APT), some studies suggested that macroeconomic variables such as the rate of inflation, the rate of change in industrial production, the difference in the rate of return between long and short term interest rates, the spread between the rate of return of high grade and low grade bonds, the rate of change in the price of oil, are a good list of priced systematic risks. Another stream of empirical asset pricing papers focused on exploiting patterns in average stock returns (frequently known as stock pricing anomalies), which were usually linked to firm's characteristics and financial ratios. For example, several studies show that a firm's average

stock return is related to its market capitalization (i.e., the size effect), the ratio of the book-to-market equity ratio (i.e., the book-to-market effect), and many other yardsticks which combine both accounting data and market price.

In this paper, we investigate the relationship between common risk factors and average returns for Italian stocks. We first follow some suggestions from APT empirical research and we identify significant macroeconomic risks using the results from factor analyses and time series regressions. Then, we study some multi-factor models combining the relevant macroeconomic variables with the mimicking equity portfolios SMB (small minus big) and HML (high minus low) proposed by Fama and French (1993). SMB and HML are stock portfolios constructed to replicate the two most common equity risks. SMB is the return spread between small capitalization and large capitalization firms controlling for other source of risks as the book-to-market effect. HML is the return spread between high book-to-market stocks and low book-to-market stocks, controlling for other source of risks as the size effect.

Our paper's main motivation is to analyse whether the influential role of the market capitalization and book-to-market equity factors in explaining average stock returns can stand up well when competing with macroeconomic variables. In other words, do equity returns carry some risk premium that is independent of either the market return or the economic forces that underlie the common variation in returns?

Our empirical work estimates risk premiums using both traditional two-pass procedures and one-pass (full information) methodologies. Using traditional two-pass procedures our evidence demonstrates that equity-derived factors such as SMB and HML cannot overcome macroeconomic variables in pricing Italian stocks. Using the one-pass methodology and robust non-parametric procedures, we confirm that both economic variables and mimicking factor risks seem jointly priced in the Italian market. However, across the different methodologies and estimation techniques, we find that only the market β and the spread between average bank loan rates and long-term government bonds are the variables consistently and significantly priced on Italian equity returns. The role of other factors, and in particular both the size and the book-to-market effects, are crucially dependent on the estimation procedure. Our research has some practical implications in the area of management of Italian stock portfolios. First, our results may be used to implement either a passive strategy, which essentially sterilizes portfolios from excessive exposure to the type of systematic risks we uncover in our research. Second, it can be also used to pursue an active strategy in which a money manager makes bets based on forecasts of risk factor realizations.