Intertemporal Market Risks and the Cross-Section of Greek Average Returns[1]

Michail Koubouros[2]

University of Peloponnese & University of Piraeus

Ekaterini Panopoulou[3]

National University of Ireland, Maynooth & University of Piraeus

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[2]Corresponding author: University of Peloponnese, Department of Economics, Terma Karaiskaki, 221 00 Tripolis, Greece. Phone: (+30) 2710 230129, fax: (+30) 2710 230139, e-mail: m.koubouros@uop.gr.

[3]Department of Banking and Financial Management, University of Piraeus, Greece and Department of Economics, National University of Ireland Maynooth, email: apano@nuim.ie.

Abstract

This paper examines whether the overall market risk along with risks reflecting uncertainty related to the long run dynamics of market cash flows (dividends) and discount rates (returns) price average returns on single-sorted portfolios of the Greek stock market. Our results suggest that a two-beta intertemporal pricing model explains half of the cross-sectional variation in average returns and delivers an economically and statistically acceptable estimate of the coefficient of relative risk aversion. Despite the relative importance of market discount-rate risk, it is market dividend-growth risk that turns out to be far more significant in determining average returns on Greek portfolios.

JEL: G11, G12, G14

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1 Introduction

Numerous studies have shown that the single beta CAPM, at least in its unconditional form, performs poorly since the cross-sectional variation in unconditional market betas cannot match the observed spread in average excess returns.[1] Recently, Campbell and Vuolteenaho (2004) and Campbell, Polk and Vuolteenaho (2005) show that the market beta can be decomposed into a relatively bad cash-flow beta, reflecting news about the market's future cash flows (dividend growth rates), and a relatively good discount-rate beta, reflecting news about the market's future discount rates (returns). According to their model the two parts of total market risk have different implications in asset pricing. Specifically, since market cash-flow shocks and discount-rate shocks represent permanent and temporary shocks to overall wealth respectively, rational conservative investors are particularly averse to the former and require a higher premium. More importantly, this cash-flow risk premium should be a multiple of their attitude toward risk. Empirically, Campbell and Vuolteenaho (2004) find that their decomposition could solve the small-value puzzle found in US data.

In this paper we study the cross-sectional behavior of cash-flow and discount-rate risks along with their ability to price returns for a set of 25 single sorted portfolios of the Greek stock market (Athens Stock Exchange, A.S.E.) for the period from 1991 to 2003. Using the empirical methodology of Campbell (1991), Campbell and Mei (1993), Campbell and Vuolteenaho (2004) and Campbell, Polk and Vuolteenaho (2005), we first estimate market cash-flow and discount-rate news and betas and then check whether the sensitivities of portfolio returns to these total market risk components can serve as sufficient risk measures which are priced in A.S.E. returns. Although some recent studies examine the properties of the two components of aggregate market return in several emerging markets (e.g. Phylaktis and Ravazzolo, 2002), there is no other study, to the best of our knowledge, which examines the asset pricing implications of this decomposition using A.S.E. data. In this respect, our study comes as a direct complement to these empirical findings since it provides some

[1] For a recent review on the CAPM literature see, among others, Fama and French (2004).

new insights, in terms of a small and emerging market, on the independent role of economic fundamentals in pricing the cross-section of average stock returns.

Our results indicate that the two-beta decomposition of the total market risk increases the ability of the static, single factor, CAPM to price Greek stock returns. More in detail, all portfolios exhibit considerable spread in risk exposure to market cash-flow and discount-rate risk and both types of risk are cross-sectionally priced. Furthermore, by employing a discrete-time intertemporal asset pricing model, we find that cash-flow risk is more important for the cross-section of average A.S.E. returns since it embodies a beta-risk premium that is much higher than the one embodied in discount-rate risk. Specifically, the two-beta model captures almost half of the variation in portfolio mean returns, performs slightly better than the popular Fama-French (1993) model and delivers meaningful and highly significant values of risk aversion. Overall, and in line with the US findings of Campbell and Vuolteenaho (2004) and Campbell, Polk and Vuolteenaho (2005), the two-beta model explains the spread in returns found across value and size portfolios and thus provides valuable insights for the small-over-large and value-over-growth puzzles.

The remainder of the paper is organized as follows: Section 2 provides the theoretical decomposition of total market risk into two parts; return risks associated with market's cash-flow and discount-rate dynamics. It also develops the intertemporal asset pricing framework that will be used for the asset pricing estimation. The dataset and the econometric methodology employed to extract the news components of market unexpected returns are given in Section 3. Section 4 presents the empirical results and, finally, Section 5 offers some concluding remarks.

2 The Model

Agents are assumed to choose their optimal consumption and portfolio positions using the recursive utility framework provided by Epstein and Zin (1989, 1991) and Weil (1989).

The lifetime utility function of the investor is given by the recursive utility function U_t, defined over current real consumption, C_t, and expected utility of future real consumption, $E_t[U_{t+1}]$:

$$U_t(C_t, E_t[U_{t+1}]) = \left\{(1-\delta)C_t^{\frac{1-\gamma}{\theta}} + \delta E_t[U_{t+1}^{1-\gamma}]^{\frac{1}{\theta}}\right\}^{\frac{\theta}{1-\gamma}} \tag{1}$$

where $0 < \delta < 1$ is the subjective discount factor, $\gamma > 0$ is the constant, under this specification, coefficient of relative risk aversion (CRRA), θ is a parameter defined as $\theta = (1-\gamma)/(1-\sigma^{-1})$, and $\sigma > 0$ is the elasticity of intertemporal substitution (EIS) between current and expected future consumption. Equation (1) has the advantage of breaking the tight link between CRRA and EIS given by power utility ($\gamma = \frac{1}{\sigma}$), thus, disconnecting investors' risk attitude across states of nature (described by γ) and across time (described by σ).

The consumer is assumed to finance all her consumption plan entirely from her total real wealth W_t, given the following dynamic budget constraint:

$$W_{t+1} = (1 + R_{W,t+1})(W_t - C_t) \tag{2}$$

where $R_{W,t+1}$ is the net real return on total wealth. Epstein and Zin (1989) solve for the optimal portfolio and consumption policies and show that the following set of conditional moment restrictions hold for each asset i and the total-wealth portfolio W:

$$E_t[\beta^\theta G^{-\frac{\theta}{\sigma}} R_{W,t+1}^{\theta-1} R_{i,t+1}] = 1; \text{ for } i = 1, ..., N \tag{3}$$

where $G_{t+1} \overset{def}{=} \frac{C_{t+1}}{C_t}$ is the optimally chosen gross growth rate of real consumption between t and $t+1$. The above set of non-linear moment restrictions can be linearized using the assumption of joint conditional log-normality of asset returns and consumption in the spirit of Hansen and Singleton (1983). Using these strong assumptions along with the dynamic budget constraint in (2), Campbell (1993, 1996) derives the following cross-sectional linear restrictions on assets' risk premia that places no role in consumption as a priced risk factor:

$$E_t[R^e_{i,t+1}] = \gamma \operatorname{cov}_t(r_{i,t+1}, r_{W,t+1} - E_t[r_{W,t+1}]) + (1-\gamma) \operatorname{cov}_t(r_{i,t+1}, -N^{DR}_{W,t+1}), \quad (4)$$

where $E_t[R^e_{i,t+1}] \overset{def}{=} E_t[R_{i,t+1}] - R_{f,t+1}$, and $R_{f,t+1}$ is the simple return on the risk-free asset. The above equation can be viewed as a discrete-time version of Merton's (1973) I-CAPM where changes in the future investment opportunity sets (captured by news about future total wealth portfolio returns, $N^{DR}_{W,t+1}$) are also priced in addition to the contemporaneous market risk (the first covariance term).

Campbell and Vuolteenaho (2004) go one step further and, using the unexpected return decomposition developed by Campbell and Shiller (1988a) and further extended by Campbell (1991), break the first factor (market return innovation) into news about future dividend (cash-flows) growth rates and news about future total returns (discount-rates). Formally, Campbell (1991) has derived the following approximate log linear decomposition of returns into time $t+1$ revision in expectations (news) about the present value of all future total-wealth dividend growth rates (cash-flow news, N^{CF}_W) and the time $t+1$ revision in expectations about the present value of all future total-wealth returns (discount-rate news, N^{DR}_W):

$$r_{W,t+1} - E_{t+1}[r_{W,t+1}] = N^{CF}_{W,t+1} - N^{DR}_{W,t+1}, \quad (5)$$

where $N^{CF}_{W,t+1} = E_{t+1}[\sum_{j=0}^{\infty} \rho^j \Delta d_{W,t+1+j}] - E_t[\sum_{j=0}^{\infty} \rho^j \Delta d_{W,t+1+j}]$ and $N^{DR}_{m,t+1} = E_{t+1}[\sum_{j=1}^{\infty} \rho^j r_{W,t+1+j}] - E_t[\sum_{j=1}^{\infty} \rho^j r_{W,t+1+j}]$, $P_{W,t+1}$ is the real aggregate (market) stock price measured at the end of period $t+1$ (ex-dividend), $d_{W,t+1} = \log(D_{W,t+1})$ is the log of the real dividend payment to total wealth during this period, $r_{W,t+1} = \log(\frac{P_{W,t+1}+D_{W,t+1}}{P_{W,t}})$ is the one-period holding log real gross return on the total wealth portfolio, $\rho_W = 1/[1 + \exp(W)]$ and $\overline{\delta_W} = E[\log(d_{W,t} - p_{W,t})]$ is the unconditional mean of the log aggregate dividend-price ratio. The first term in (5) is the time $t+1$ revision in dividend growth expectations and represents a permanent positive effect on total wealth since it is never reversed subsequently, whereas the second one is the time $t+1$ revision in expectations about future returns on total wealth and thus can be viewed as a temporary shock to the total wealth since the unexpected capital gain today $\left(r_{W,t+1} - E_{t+1}[r_{W,t+1}] > 0\right)$ is at a cost of lower future investment opportunities, i.e. $E_{t+1}[\sum_{j=1}^{\infty} \rho^j r_{W,t+1+j}] - E_t[\sum_{j=1}^{\infty} \rho^j r_{W,t+1+j}] < 0$.

Using the above decomposition of the total wealth unexpected return and the two factor asset pricing restriction in (4) we get the following asset pricing model that assigns different roles for aggregate dividend growth rates' news and returns' news in determining asset risk premia:

$$E_t[R^e_{i,t+1}] = \gamma \operatorname{cov}_t(r_{i,t+1}, N^{CF}_{W,t+1}) + \operatorname{cov}_t(r_{i,t+1}, -N^{DR}_{W,t+1}), \tag{6}$$

The covariance risk premium representation in (6) can have an equivalent beta-like premium representation (Cochrane, 2001). Multiplying and dividing by the variance of total-wealth return innovations, $\operatorname{var}_t(r_{W,t+1} - E_t\left[r_{W,t+1}\right])$, we get:

$$E_t[R^e_{i,t+1}] = \lambda_{CF,t}\beta_{i,CF,t} + \lambda_{DR,t}\beta_{i,DR,t} \tag{7}$$

with $\lambda_{CF,t} = \gamma \operatorname{var}_t(r_{W,t+1} - E_t[r_{W,t+1}])$, $\lambda_{DR,t} = \operatorname{var}_t(r_{W,t+1} - E_t[r_{W,t+1}])$ and:

$$\begin{aligned} \beta_{i,W,t} &= \frac{\operatorname{cov}_t(r_{i,t+1}, N^{CF}_{W,t+1})}{\operatorname{var}_t(r_{W,t+1} - E_t[r_{W,t+1}])} + \frac{\operatorname{cov}_t(r_{i,t+1}, -N^{DR}_{W,t+1})}{\operatorname{var}_t(r_{W,t+1} - E_t[r_{W,t+1}])} \\ &= \beta_{i,CF,t} + \beta_{i,DR,t} \end{aligned} \tag{8}$$

Equation (8) states that the required risk premium on asset i is jointly determined by the betas of its return with the corresponding decomposed components of the total market risk; cash-flow and discount-rate beta that add to the full total wealth, CAPM, beta. A conservative risk-averse investor ($\gamma > 1$) demands a higher risk price for risks associated with total-wealth cash-flow (dividend growth) uncertainty ($\beta_{i,CF}$) rather than for risks linked to shocks to total wealth portfolio returns ($\beta_{i,DR}$), since any positive (negative) shock to wealth discount rates is at a benefit (cost) of worse future investment opportunities, whereas the investor is never compensated later for every positive (negative) shock to dividends. Hence, the beta price of market cash-flow risk λ_{CF} is a γ multiple of the beta risk prices of market discount-rate risk λ_{DR}. Thus, for a conservative investor it must be $\lambda_{CF} > \lambda_{DR} > 0$.

In order to get comparable results to the empirical literature of the unconditional CAPM and, more importantly, to the empirical findings of the two-beta model of Campbell and Vuolteenaho (2004) that places a relatively more important role in cash-flow risk, we condition down equation (7) and proceed with its unconditional version.

3 Data and Empirical Methodology

Our study is based on monthly Greek asset and macroeconomic data for the period from June 1991 to May 2003 (133 monthly observations) obtained from the Datastream International database. Specifically, our data consist of (a) different sets of common stock test portfolios sorted on various firm specific characteristics such as book-to-market, dividend yield, market capitalization, price-earnings ratio and 3-month momentum, and (b) a set of economy-wide variables that serve as instruments. The sorting characteristics where chosen in order to generate clear spreads in average returns that will challenge the empirical validity of the two-beta asset pricing model. On the other hand, and following the common practice, the state variables have been selected under the assumption that they exhibit some forecasting ability over future portfolio returns. Lastly, we assume that the total market value-weighted portfolio is a good proxy for the total-wealth portfolio in the Greek economy, so that $R_W = R_M$.

We employ a variant of the Fama and French (1993) methodology to construct value-weighted returns on 25 firm-characteristic portfolios sorted on the above characteristics, and returns on the two Fama and French (1993) aggregate size and book-to-market factor mimicking portfolios, Small-Minus-Big (SMB) and High-Minus-Low (HML), respectively. The latter factor-mimicking portfolios will be used as benchmarks in our asset pricing tests.

The portfolio construction procedure has as follows. In June, every year, we break the full menu of A.S.E. common stocks available into 5 groups based (once at a time) on last-month book-to-market, dividend yield, market capitalization, price-earnings ratio and 3-month momentum, so that each group contains an equal number of stocks. We first collect monthly closing prices for each stock and since the theoretical decomposition in (5) requires continuous data on dividends we divide the annual dividend payment by 12 and add it to the monthly closing price.[2] Then, we compute the value-weighted monthly holding period simple portfolio return by weighting each stock by its relative contribution to the portfolio's total capitalization. The procedure is repeated every year and we end up with time-series data of simple returns on each characteristics-sorted portfolio. Finally, and although the model in (7) is written in real log returns, we assume that for the monthly test interval we employ, inflation rates are almost fully forecastable, and thus we proxy real log returns with nominal log returns.

For the construction of the returns on the aggregate value factor-mimicking zero-cost portfolio (High Minus Low, HML) we used the 30-40-30 rule employed by Fama and French (1993). However, for the aggregate size factor-mimicking zero-cost portfolio (Small Minus Big, SMB) we adjust the formation procedure to account for the characteristics of the Greek data. We use the 70th quantile of the total market value instead of the median that was used by Fama and French. Given, that few large stocks dominate the Greek stock market, a 50% sorting would generate a small-cap portfolio that would represent only a very small proportion of the total market value. In this respect, using a larger breakpoint we can create a distribution of aggregate market value across portfolios that is relatively similar to the distribution in Fama and French(1993), while the small capitalization portfolio represents on average the 8% of the total A.S.E. market capitalization.[3] At the end of June

[2]Although this technique of spreading the dividends over the year on the closing price assumes strong form efficiency of the market, it is common practice when constructing total market indexes that assume reinvestment of dividends over the next period. We thank an anonymous referee for pointing this out.

[3]For a similar construction procedure using data from the UK market, see Dimson, Nagel and Quigley (2003).

of each year, we create the size and book-to-market double-sorted portfolios of Fama and French (1993) (SL, SM, SH, BL, BM and BH) and calculate the value-weighted monthly returns for the next 12 months. Then, the returns on the zero-cost aggregate book-to-market and size portfolios are defined as $HML = (SH + BH)/2 - (SL + BL)/2$ and $SMB = (SL + SM + SH)/3 - (BL + BM + BH)/3$, respectively.

The second set used in our analysis consists of variables that have proven successful in predicting the future state of the economy and asset returns. The innovations of these variables are used to generate cash-flow and discount-rate news through a VAR(1) specification. More in detail, we use: (a) the monthly log difference of the OECD leading indicator, $\Delta \log (LI)$, (b) the market log price-earnings ratio, $p - e$, and (c) the small-stock value spread, VS, defined as the difference between the log(B/M) of the small high-B/M portfolio and the log(B/M) of the small low-BE/ME portfolio.[4]

The asset pricing model in (7) uses cash-flow and discount-rate news as priced factors. We follow Campbell (1991) and we estimate them using a first-order vector autoregressive, VAR(1), model. We first estimate expected returns and the revisions in expectations about future returns ($E_t[r_{M,t+1}]$ and $(E_{t+1} - E_t) \sum_{j=1}^{\infty} \rho_M^j r_{M,t+1+j}$, respectively) and then we use $r_{M,t+1}$ and equation (5) to back out the market cash-flow news. This practice has an important advantage as it relies only on the dynamics of expected returns and there is no need for modelling the dynamics of dividends since the latter are derived by the VAR estimates and the realizations of returns and state variables.

We assume that the data are generated by the following VAR(1) model:

$$y_{t+1} = \Gamma + A y_t + u_{t+1}, \tag{9}$$

where $y_{t+1} = (r_{m,t+1}, y_{1,t+1}, ..., y_{m,t+1})$ is a $m \times 1$ vector of variables containing returns

[4]Recently, the value spread VS variable has been found to be a good forecaster of US returns. See, among others, Campbell and Vuolteenaho (2004), Campbell, Polk and Vuolteenaho (2005) and Koubouros, Malliaropulos and Panopoulou (2005). For a general discussion on its foecasting ability see Liu and Zhangy (2005). Following this evidence we use the value spread as a predictor of A.S.E. returns.

as its first element and $(m-1)$ variables which have predictive power for returns, Γ is a $m \times 1$ vector of constants and A is a $m \times m$ matrix of constants. We estimate (9) for the market return and then compute cash-flow and discount-rate news as linear functions of the $t+1$ vector of innovations, u_{t+1}:

$$N_{M,t+1}^{DR} = e1'\lambda u_{t+1}, \qquad N_{M,t+1}^{CF} = (e1' + e1'\lambda)u_{t+1}, \tag{10}$$

where $e1$ is a $m \times 1$ vector with the first element equal to unity and the remaining elements equal to zero. The mapping of the shock vector to the news vectors is given by $\lambda \equiv \rho A(I_m - \rho A)^{-1}$. $e1'\lambda$ captures the long-run significance of each individual VAR(1) shock to discount-rate expectations. The greater the absolute value of a variables coefficient in the return prediction equation (the top row of A), the greater the weight the variable receives in the discount-rate-news formula (10). Also, more persistent variables should also receive more weight, which is captured by the term $(I_m - \rho A)^{-1}$.

4 Empirical Evidence

4.1 Estimation of Cash-Flow and Discount-Rate News and Betas

Table 1 (Panel A) reports parameter estimates for the market VAR(1) model. Our estimates suggest that the state variables have some predictive power for stock market excess returns (adj.-R^2 of 9.5%). Specifically, monthly market returns display some degree of mean reversion as depicted in the statistically significant autoregressive coefficient of 0.147. The effect of the log change of the OECD Leading Indicator, $\Delta \log(LI)$, on market returns is positive, a finding consistent with the positive relationship of output growth and stock market returns. The remaining state variables, namely the log price-to-earnings ratio $(p-e)$ and the small-stock value spread (VS), positively predict the market return. Our findings are in contrast with findings in previous research (see, e.g. Campbell and Shiller, 1988a, 1988b, 1998, Rozeff, 1984, Fama and French, 1988, 1989, Eleswarapu and Reinganum, 2004 and Brennan, Wang and Xia, 2004). The remaining columns of Table 1 summarize the dynamics of the state variables. The growth of the OECD Leading Indicator process is positively autocorrelated with a coefficient of 0.528, while both the $p-e$ and the VS display an increased degree of persistence as suggested by a coefficient estimate of 0.97. This persistence does not induce any estimation problems as no instability is apparent at the VAR(1) residuals.

To be on the safe side, though, we also tested our variables for stationarity prior to estimating the VAR(1) model. In the present case, the results from a variety of unit-root tests (reported in Panel B of Table 1) are, as usual, mixed especially for the $p-e$ and the VS.[5] When the null hypothesis of stationarity is tested, the KPSS test fails to reject the null for all the variables at hand. When the null hypothesis of a unit root is tested, the standard Dickey-Fuller (DF) or Phillips-Perron (PP) tests typically fail to reject the null for the $p-e$ and the VS. The GLS versions of the DF tests, however, being more powerful than the standard DF tests, reject the unit root null in all the cases under consideration. The general picture emerging from the empirical literature and our own tests suggests treating our variables and especially $p-e$ and VS as having a highly persistent but ultimately stationary univariate representation. Moreover, Panel C of Table 1 reports the ARCH-LM tests for heteroskedasticity in the VAR(1) residuals, which do not suggest any second-order dependence in the error terms.

Table 2 summarizes the behavior of implied (from the VAR(1) specification) cash-flow news ($N_{M,t+1}^{CF}$) and discount-rate news ($N_{M,t+1}^{DR}$) components of market returns. The top panel shows that the standard deviation of discount-rate news is more than twice the standard deviation of cash-flow news. This finding is consistent with Campbell (1991) and Campbell and Vuolteenaho (2004). However, in contrast to Campbell and Vuolteenaho (2004), but in line with Campbell (1991 and 1996), the two components of return exhibit some degree of

[5] The unit root null is tested by means of the following tests: the standard Dickey-Fuller test (Dickey and Fuller, 1979), the Dickey-Fuller test with GLS detrending (Elliott, Rothenberg and Stock, 1996) and the Phillips-Perron test (Phillips and Perron, 1988). The stationary null hypothesis is tested by means of the KPSS test (Kwiatkowski, Phillips, Schmidt and Shin, 1992).

correlation, 0.563. In what follows, we use discount-rate neutral cash-flow news resulting from regressing market cash-flow news on discount-rate news and keeping the estimated constant plus the residuals, in order to examine the independent ability of the two in pricing average returns.

The bottom panel of Table 2 reports correlations of cash-flow and discount-rate news with innovations in market excess returns and state variables. Discount-rate and cash-flow news are negatively correlated with innovations in the market return and the price-earnings, respectively. In contrast, innovations to the value spread are strongly positively correlated with discount-rate and cash-flow news.

Empirical measures of the cash-flow and discount-rate betas in (8) are derived using a methodology similar to this employed in Campbell and Vuolteenaho (2004) to ensure that our sample estimates are not affected by non-synchronous trading (see, for example, Scholes and Williams, 1977 and Dimson, 1979) and under-reaction of stock prices to changes in the market index, especially for large stocks (see, for example, McQueen, Pinegar and Thorly, 1996 and Peterson and Sanger, 1995). Our two sample betas, which will be used in the cross-sectional regression analysis, are defined as the sum of contemporaneous, one lag and two lag full-sample covariances of portfolio returns at $t+1$ with market news, divided by the full-sample variance of the market return innovations, $\widehat{\text{var}}(r_{M,t+1} - E_t[r_{M,t+1}])$. As a result, the beta components of the full market beta (cash-flow news' beta $\widehat{\beta}_{i,CF}$ and discount-rate news' beta $\widehat{\beta}_{i,DR}$) are estimated as follows:

$$\widehat{\beta}_{i,CF\ (\text{or}\ -DR)} = \sum_{k=0}^{2} \frac{\widehat{\text{cov}}(r_{i,t+1}, N_{M,t+1-k}^{CF\ (\text{or}\ -DR)})}{\widehat{\text{var}}(r_{M,t+1} - E_t[r_{M,t+1}])} \tag{11}$$

The popular three-factor Fama-French (1993) asset pricing model is used as a benchmark. In order to keep the comparison of the results of this asset pricing model in line with the two-beta asset pricing model, we estimate betas with the aggregate market, size and value factor mimicking portfolios ($\widehat{\beta}_{i,M}$, $\widehat{\beta}_{i,SMB}$ and $\widehat{\beta}_{i,HML}$, respectively) with two lags in the covariance term as follows (11):

$$\widehat{\beta}_{i,p} = \sum_{k=0}^{2} \frac{\widehat{\text{cov}}(r_{i,t+1}, r_{p,t+1-k})}{\widehat{\text{var}}(r_{p,t+1} - E_t[r_{p,t+1}])}, \text{ for } p = r_M, r_{SMB}, r_{HML} \tag{12}$$

Table 3 reports the summary statistics of the annualized mean and standard deviation as well as the 1, 2, 3, 6, and 12 month autocorrelations of the returns on the value-weighted market portfolio and the book-to-market, dividend yield, size, price-to-earnings and 3-month momentum sorted portfolios, respectively. The average annualized return on the market portfolio is 9.85% with a standard deviation of 3.1%. The autocorrelation of the return is diminishing with the lag length, even turning negative for horizons of 9-12 months. Our data set reveals an average annual value premium of 7.48% and an average annual size premium of 25.19%. Similarly, high dividend-yield, low price-earnings and 3-month momentum portfolios yield an average annual premium of 7.49%, 14.27% and 14.77%, over the low dividend-yield, high price-earnings and 3-month losers' portfolios, respectively. The considerable spread in average returns provides a challenge to traditional asset pricing theory since it should be matched with an equivalent spread in aggregate risk exposure.

Table 4 reports the estimated betas given by our definition in (11) along with their respective standard errors.[6] The main characteristic of our results is that our methodology generates considerable spread in the overall market risk $\beta_{i,M}$ (the sum of individual cash-flow and discount-rate betas defined in (8)) especially for the value and size portfolios. This fact may be consistent with the static CAPM that states that overall market risk (beta) can be sufficient to capture differences in the cross-section of expected returns. The observed spread in the two aggregate bad (cash-flow) and good (discount-rate) betas confirm the story argued by Campbell and Vuolteenaho (2004) and Campbell, Polk and Vuolteenaho (2005) that value stocks (high B/M) have relatively high cash-flow betas while growth stocks have

[6] Those beta coefficients and their related standard errors were obtained by regressing the relevant components and adjusting for the disparity caused by the modified variance. For example, if we want to estimate $\beta_{i,CF}$ as given by (11), we run the regression of $r_{i,t+1}$ on $N_{M,t+1}^{CF}[\text{var}(r_{M,t+1} - E_t[r_{M,t+1}])/\text{var}(N_{M,t+1}^{CF})]$ as well as the two lag terms.

relatively high discount-rates betas. More importantly, both components of total market risk increase with value and decrease with size indicating that both are important for the relative riskiness of value-growth and small-large portfolios, respectively. Lastly, all portfolios exhibit considerable spread in their return exposure to aggregate size and distress risk as captured by SMB and HML betas, indicating that the three-factor Fama-French (1993) model could be an alternative to the CAPM. However, and since the two factor portfolios do not mimic clear fundamental (economy-wide) sources of risk, we will use this model as a practical tool for comparison purposes.

Given that the Greek stock market has undergone a considerable amount of changes in the period under examination, such as the abolition of capital controls in 1994, the peak of the aggregate market index in 1999 and its upgrade from emerging to developed market in 2001, we engaged in examining the structural stability of our estimated betas prior to including them in the asset pricing models.[7] In this respect, we conducted a series of Chow's breakpoint tests (1960) using several dates as possible breakpoints. Our results for a possible breakpoint in January 2001 are reported in Table 5. Interestingly, we fail to reject the null of no structural change at the 5% level for both the cash-flow and discount-rate betas, as well as the overall market betas. On the other hand, evidence with respect to the HML and SMB betas is not robust. Specifically, we reject the null of stability for 14 and 4 out of 25 HML and SMB betas respectively.[8]

4.2 Are Cash-Flow and Discount-Rate Risks Individually Priced?

Having estimated the full-sample cash-flow and discount-rate betas given our specification of the return generating processes in (9) we proceed with cross-sectional asset pricing tests to evaluate the ability of our two-beta model to capture cross-sectional variation in A.S.E. average portfolio returns. We follow Campbell and Vuolteenaho (2004) and study

[7] We thank an anonymous referee for pointing this out.

[8] Our results with respect to other breakpoints are qualitatively similar and are not reported for brevity, but are available from the authors.

the unconditional version of the asset pricing model in (7). However, and given the low quality of risk-free rate data for our sample period we proceed with the zero-beta versions of our asset pricing tests. So, the constant term λ_0 in the linear specifications below is no longer the average pricing error as it would be in (4), (6) and (7), and thus, it can (or better should, under the hypothesis of the existence of a zero-beta asset in A.S.E.) be different from zero. The model is tested against the static CAPM and the Fama-French (1993) three-factor.

More specifically, we consider the following cross-sectional specification of the two beta (cash-flow and discount-rate) model:

$$E_T[R_i] = \lambda_0 + \lambda_{CF}\widehat{\beta}_{i,CF} + \lambda_{DR}\widehat{\beta}_{i,DR}, \tag{13}$$

and we test this two-beta specification against the popular static single-beta CAPM that imposes the same risk prices in cash-flow and discount-rate risk and thus prices aggregate market risk, $\beta_{i,M}$:

$$E_T[R_i] = \lambda_0 + \lambda_M\widehat{\beta}_{i,M}, \tag{14}$$

and the popular three-factor Fama-French (1993) model that adds aggregate value (HML) and size (SMB) factor mimicking portfolios as competing factors to the aggregate market return:

$$E_T[R_i] = \lambda_0 + \lambda_M\widehat{\beta}_{i,M} + \lambda_{HML}\widehat{\beta}_{i,HML} + \lambda_{SMB}\widehat{\beta}_{i,SMB}, \tag{15}$$

In all equations $E_T[R_i]$ denotes average (sample mean) portfolio returns and $\widehat{\beta}_{i,k}$ denote the estimated betas on the k^{th} factor as defined in (11) and (12). We estimate the unconditional unrestricted prices of beta risks ($\widehat{\lambda}$s) for the aforementioned models as well as the following restricted version of the two-beta model in (16):

$$E_T[R_i] = \lambda_0 + \gamma\lambda\widehat{\beta}_{i,CF} + \lambda\widehat{\beta}_{i,CF} \tag{16}$$

This last version enables to estimate the coefficient of relative risk aversion γ and the risk premium on the discount-rate factor λ. The model predicts that the premium associated with market cash-flow risk must be a γ multiple of the premium associated with discount-rate risk. For a conservative risk-averse investor ($\gamma > 1$ in (1)), λ_{CF} must be greater than λ_{DR}, i.e. $\lambda_{CF} > \lambda_{DR}$.

Table 6 presents the empirical findings of the cross-sectional asset pricing tests. The table reports the mean and standard error for each estimate, as well as the average adj.-R^2 of the regression. Figure 1 gives a visual illustration of the empirical ability of the alternative models by plotting the realized and fitted average returns. The better the model performs the closest to the 45-degree line the points fall. A perfect match ($R^2 = 100\%$) is achieved when all points fall on the 45-degree line.

Contrary to many US studies (e.g. Fama and French, 1992, Campbell, Polk and Vuolteenaho, 2004), the traditional static CAPM performs quite well and explains almost half of the cross-sectional variation in average returns. However, it fails to produce a significant estimate for the zero-beta coefficient ($\widehat{\lambda}_0 = -0.005$ with $s.e. = 0.0048$).

Next, we check whether the two-beta decomposition in (13) with unrestricted prices of beta risk can improve the empirical validity of the standard static CAPM and whether there are different roles in market cash-flow and discount-rate risks. The model performs quite well and generates statistically significant premia and explains 46.6% of the observed cross-sectional variation in A.S.E. portfolio returns. More importantly, it generates significant risk premia for both types of risk with the premium associated with market cash-flow risk being much higher than that associated with market's discount-rate risk ($\widehat{\lambda}_{CF} = 0.0274$ and $\widehat{\lambda}_{DR} = 0.0096$). These results are in line with Campbell and Vuolteenaho (2004) and in favor of the total market risk decomposition in (7) and (8). Further, when we estimate the restricted version of the model in (16) the factor of proportionality, which is restricted to be equal to the coefficient of relative risk aversion, γ, is both economically and statistically significant. Specifically, the estimate of $\widehat{\gamma} = 2.8572$ ($s.e. = 0.1612$) is in the range hypothesized by Mehra and Prescott (1985) that could solve the well known equity premium puzzle. We also tested for unconstrained risk premia using cash-flow and discount-rate risk once at a time (see, columns labelled CF and DR). Again, our results indicate that, although both types of intertemporal market risks are needed to describe the cross-section of returns, cash-flow risk is much more important with a risk premium three times higher than the one of discount-rate risk. The respective estimates are $\widehat{\lambda}_{CF} = 0.0320$ and $\widehat{\lambda}_{DR} = 0.0107$. The two Fama-French (1993) factors, HML and SMB, perform relatively well by explaining the same proportion of cross-sectional volatility as the two-beta model, but fail to deliver positive and statistically significant premium for the overall market risk ($\widehat{\lambda}_M = -0.0015, s.e. = 0.0064$). What is more, none of the aggregate value and size premia (λ_{HML} and λ_{SMB}) are significant at the 1% level. Lastly, and for experimental purposes we use all factors in an extended model. Our results suggest that the relative importance of cash-flow risk is clear but we are inconclusive on the one of the discount-rate risk, especially when the significant size risk is included.

5 Conclusions

This paper builds on the decomposition of the overall market, or CAPM, risk into parts reflecting time variation related to the dynamics of aggregate market cash flows and discount rates using data from the small and emerging Greek stock market (Athens Stock Exchange). Employing the methodology of Campbell (1991), Campbell and Mei (1993) and Campbell and Vuolteenaho (2004) we decompose market betas into two sub-betas, associated with revisions in expectations about future market dividend growth rates and future returns. Using a VAR(1) approach and a discrete time version of Merton's I-CAPM, we test whether these components of overall market risk are rationally priced and thus explain the value, size and momentum premia observed in our monthly 1991-2003 sample. The theoretical model predicts that although both types of risk are important for the cross-section, market cash-flow risk (captured by the sensitivity of returns to market cash-flow news) should earn a higher beta-risk premium than market discount-rate risk.

The two-beta model performs quite well in pricing average returns on single-sorted portfolios according to book-to-market, dividend-yield, market capitalization, price-earnings and 3-month momentum. Consistent with theory, the model delivers an economically and statistically significant estimate of the coefficient of relative risk aversion (close to 3), explains almost half of the cross-sectional variation in A.S.E. portfolio returns and generally performs at least as good as the popular three-factor Fama-French (1993) model. We find that the exposure of Greek stock portfolios to risks associated with permanent shocks to aggregate market value (captured by market cash-flow risk) is compensated with higher unconditional risk prices than the exposure to risks associated with future market returns. Our results are in favor of a rational risk I-CAPM-type story where economic agents have a long-term optimizing behavior, do not behave myopically and value stocks according to their long-run riskiness.

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Table 1.

VAR estimates for market portfolio and diagnostic tests

Panel A: VAR Estimates				
	$r_{M,t+1}$	$\Delta \log (LI_{t+1})$	$p_{t+1} - e_{t+1}$	VS_{t+1}
constant	−0.033	0.023	0.055	−0.032
	(0.021)	(0.009)	(0.029)	(0.041)
$r_{M,t}$	0.147	−0.056	0.043	−0.185
	(0.086)	(0.039)	(0.119)	(0.165)
$\Delta \log (LI_t)$	0.194	0.528	−0.077	0.405
	(0.160)	(0.073)	(0.223)	(0.308)
$p_t - e_t$	0.007	−0.011	0.969	0.044
	(0.001)	(0.007)	(0.021)	(0.028)
VS_t	0.036	−0.003	−0.034	0.967
	(0.006)	(0.006)	(0.018)	(0.025)
R^2	9.5%	32.9%	94.7%	91.9%
F-stat.	3.530	16.499	598.4	384.70

Panel B: Unit Root Tests				
	r_M	$\Delta \log (LI)$	$p - e$	VS
ADF	−10.199***	−6.391***	−1.722	−1.746
ADF-GLS	−10.200***	−6.088***	−1.740*	−1.759*
PP	−10.259***	−6.388***	−1.944	−2.046
KPSS	0.229	0.095	0.180	0.137

Panel C: LM Test for Heteroscedasticity (ARCH Test: lag = 4)				
	$\hat{u}_{r_M}$	$\hat{u}_{\Delta \log(LI)}$	$\hat{u}_{p-e}$	$\hat{u}_{VS}$
F-stat.	0.327	1.081	2.134	0.709
p-value	[0.859]	[0.369]	[0.080]	[0.593]

Note: Panel A. presents estimates of the VAR(1) system in (9). $r_{M,}$ is the value-weighted market return, $\Delta \log (LI)$ is the change in the logarithm of the OECD leading indicator, $p - e$ is the market log price-earnings ratio and VS is the value-spread defined as the difference between the log(B/M) of the small high-B/M portfolio and the log(B/M) of the small low-BE/ME portfolio. Standard errors of the estimates are in parentheses. Panel B. presents the unit-root tests for the state variables used in the VAR(1). ADF, ADF-GLS, PP and KPSS stand for the values of the Dickey-Fuller, Dickey-Fuller with GLS detrending, Phillips-Perron and Kwiatkowski-Phillips-Schmidt-Shin tests, respectively. *, ** and *** denote significance at 10%, 5% and 1%, respectively. Panel C. reports the values of ARCH heteroscedasticity tests on the estimated VAR(1) residuals. The sample period spans from June 1991 to May 2003.

Table 2.

Market portfolio cash-flow and discount-rate news

	Covariance matrix of news			News corr/std.dev.	
	N_M^{CF}	N_M^{DR}		N_M^{CF}	N_M^{DR}
N_M^{CF}	0.0081	0.0034	N_M^{C}	0.091	0.563
N_M^{DR}	0.0034	0.0046	N_M^{D}	0.563	0.215
	Correlations of innovations with news			Functions	
Innovations/News	N_m^{CF}	N_m^{DR}		N_M^{CF}	N_M^{DR}
r_M	0.679	-0.225	r_M shock	1.052	0.052
$\Delta \log (LI)$	0.277	0.295	$\Delta \log (LI)$ shock	0.683	0.683
$p - e$	-0.224	0.398	$p - e$ shock	0.273	0.273
VS	0.603	0.855	VS shock	0.399	0.399

Note: The table reports the estimated covariance matrix (upper-left) and the correlation matrix with standard deviations (upper-right) of the estimated market portfolio cash-flow and discount rate news using equations (9) to (10), the correlations of innovations of state variables with market news (lower-left) and the mapping functions defined in (10). The sample period spans from June 1991 to May 2003.

Table 3.

Summary statistics

Portfolio	Mean	Std. Dev	ρ_1	ρ_2	ρ_3	ρ_6	ρ_9	ρ_{12}
Market portfolio								
	9.85	3.06	0.13	0.08	0.06	0.05	-0.11	-0.05
Panel A. Book-to-market Portfolios								
High	16.07	3.67	0.12	0.01	0.16	0.12	-0.10	0.02
2	8.16	3.42	0.19	0.09	0.07	0.03	-0.13	-0.06
3	9.99	3.44	0.16	0.14	-0.02	0.07	-0.15	0.00
4	5.63	3.30	0.13	0.16	0.09	0.00	-0.03	-0.07
Low	8.59	3.25	0.15	0.06	0.04	0.05	-0.01	-0.10
Panel B. Dividend-Yield Portfolios								
High	13.62	2.96	0.07	0.02	0.06	0.05	-0.17	-0.02
2	6.05	3.47	0.11	0.15	0.00	0.06	-0.13	0.06
3	9.40	3.39	0.18	0.11	0.01	0.00	-0.06	-0.09
4	8.25	3.44	0.15	0.09	0.15	0.05	-0.06	-0.08
Low	6.13	3.50	0.16	0.22	0.05	0.06	-0.11	-0.08
Panel C. Size Portfolios								
Large	9.08	2.97	0.10	0.02	0.02	0.04	-0.15	-0.02
2	9.05	3.69	0.19	0.19	0.16	0.09	0.03	-0.03
3	14.68	4.09	0.26	0.17	0.19	0.07	0.03	-0.07
4	16.28	4.33	0.25	0.23	0.25	0.13	0.05	-0.05
Small	34.27	5.02	0.27	0.26	0.27	0.15	0.08	-0.15
Panel D. Price-to-Earnings Portfolios								
High	5.53	3.89	0.24	0.14	0.08	-0.03	-0.02	-0.12
2	6.60	3.53	0.17	0.11	0.04	0.06	0.00	-0.06
3	10.19	3.37	0.18	0.24	0.08	0.11	-0.10	-0.04
4	9.32	2.99	0.13	0.10	-0.01	0.11	-0.11	0.02
Low	19.80	3.05	0.13	0.07	0.03	0.14	-0.10	0.03
Panel E. 3-Month Momentum Portfolios								
Winners	17.78	3.37	0.28	0.22	0.19	0.11	-0.11	-0.06
2	13.84	3.52	0.09	0.12	0.10	0.10	-0.15	-0.10
3	13.70	3.56	0.21	0.22	0.09	0.06	-0.01	-0.03
4	5.26	3.71	0.10	0.08	0.01	0.03	0.06	-0.09
Losers	3.01	4.03	0.04	0.04	0.01	0.08	-0.06	0.01

Note: The table presents the annualized mean and standard deviation as well as the 1, 2, 3, 6, and 12 month autocorrelations of the returns on the value-weighted market portfolio and the book-to-market, dividend yield, size, price-to-earnings and 3-month momentum sorted portfolios, respectively. The sample period spans from June 1991 to May 2003.

Table 4.

CAPM, cash-flow, discount-rates, HML and SMB betas

Panel A. Book-to-market portfolios

	$\widehat{\beta}_{i,M}$	s.e.	$\widehat{\beta}_{i,CF}$	s.e.	$\widehat{\beta}_{i,DR}$	s.e.	$\widehat{\beta}_{i,HML}$	s.e.	$\widehat{\beta}_{i,SMB}$	s.e.
High	1.445	0.222	0.760	0.131	0.685	0.235	-0.185	0.246	0.364	0.243
2	1.300	0.216	0.714	0.149	0.586	0.253	-0.189	0.218	0.224	0.205
3	1.170	0.196	0.706	0.162	0.464	0.226	0.038	0.233	0.140	0.224
4	1.183	0.218	0.621	0.122	0.562	0.235	0.244	0.245	0.027	0.190
Low	0.893	0.212	0.479	0.124	0.414	0.232	0.374	0.210	0.077	0.195

Panel B. Dividend-yield portfolios

	$\widehat{\beta}_{i,M}$	s.e.	$\widehat{\beta}_{i,CF}$	s.e.	$\widehat{\beta}_{i,DR}$	s.e.	$\widehat{\beta}_{i,HML}$	s.e.	$\widehat{\beta}_{i,SMB}$	s.e.
High	1.025	0.156	0.684	0.132	0.341	0.186	-0.177	0.154	-0.030	0.163
2	1.046	0.217	0.633	0.185	0.412	0.240	-0.115	0.232	0.022	0.209
3	1.270	0.240	0.681	0.135	0.589	0.257	0.274	0.239	0.005	0.199
4	1.102	0.262	0.526	0.122	0.576	0.258	0.351	0.232	0.379	0.221
Low	1.170	0.272	0.577	0.151	0.593	0.241	0.346	0.228	0.251	0.227

Panel C. Size portfolios

	$\widehat{\beta}_{i,M}$	s.e.	$\widehat{\beta}_{i,CF}$	s.e.	$\widehat{\beta}_{i,DR}$	s.e.	$\widehat{\beta}_{i,HML}$	s.e.	$\widehat{\beta}_{i,SMB}$	s.e.
Large	0.850	0.141	0.608	0.107	0.242	0.190	0.037	0.155	-0.198	0.188
2	1.474	0.312	0.641	0.171	0.834	0.279	0.151	0.288	0.572	0.203
3	1.713	0.372	0.632	0.183	1.081	0.301	0.103	0.342	1.025	0.220
4	1.833	0.398	0.648	0.171	1.185	0.307	0.233	0.352	1.434	0.238
Small	2.275	0.517	0.714	0.234	1.561	0.363	0.154	0.430	1.764	0.308

Panel D. Price-to-earnings portfolios

	$\widehat{\beta}_{i,M}$	s.e.	$\widehat{\beta}_{i,CF}$	s.e.	$\widehat{\beta}_{i,DR}$	s.e.	$\widehat{\beta}_{i,HML}$	s.e.	$\widehat{\beta}_{i,SMB}$	s.e.
High	1.364	0.287	0.658	0.163	0.706	0.275	0.518	0.296	0.402	0.220
2	1.178	0.194	0.624	0.128	0.554	0.227	0.287	0.237	0.115	0.225
3	1.150	0.245	0.625	0.156	0.525	0.265	0.025	0.218	0.230	0.217
4	1.024	0.176	0.610	0.138	0.414	0.188	-0.062	0.177	-0.018	0.170
Low	1.025	0.194	0.730	0.173	0.296	0.215	-0.029	0.167	0.070	0.150

Panel E. 3-month momentum

	$\widehat{\beta}_{i,M}$	s.e.	$\widehat{\beta}_{i,CF}$	s.e.	$\widehat{\beta}_{i,DR}$	s.e.	$\widehat{\beta}_{i,HML}$	s.e.	$\widehat{\beta}_{i,SMB}$	s.e.
Winners	1.338	0.348	0.699	0.151	0.639	0.340	0.299	0.224	0.539	0.261
2	1.081	0.288	0.526	0.111	0.555	0.294	0.214	0.224	0.385	0.245
3	1.324	0.328	0.678	0.152	0.646	0.315	0.251	0.247	0.432	0.200
4	1.213	0.236	0.518	0.166	0.696	0.253	0.040	0.301	0.241	0.193
Losers	1.101	0.248	0.424	0.163	0.677	0.249	-0.159	0.341	0.077	0.176

Note: The table presents the estimated market (CAPM) betas ($\widehat{\beta}_{i,M}$), cash-flow betas ($\widehat{\beta}_{i,CF}$), discount–rate betas ($\widehat{\beta}_{i,DR}$), HML betas ($\widehat{\beta}_{i,HML}$) and SMB betas ($\widehat{\beta}_{i,SMB}$) of the book-to-market, dividend yield, size, price-to-earnings and 3-month momentum sorted portfolios, respectively, along with their standard errors. The betas were estimated using (10), (11) and (12). The estimation period spans from June 1991 to May 2003.

Table 5.

Structural stability tests: CAPM, cash-flow, discount-rates, HML and SMB betas

Panel A. Book-to-market portfolios										
	$F_{\widehat{\beta}_{i,M}}$	p-val	$F_{\widehat{\beta}_{i,CF}}$	p-val	$F_{\widehat{\beta}_{i,DR}}$	p-val	$F_{\widehat{\beta}_{i,HML}}$	p-val	$F_{\widehat{\beta}_{i,SMB}}$	p-val
High	2.099	0.150	0.181	0.671	3.620	0.057	5.888	0.015	5.779	0.018
2	3.682	0.057	0.176	0.676	2.248	0.134	6.597	0.010	2.029	0.157
3	0.218	0.642	0.124	0.726	2.162	0.144	2.781	0.095	1.654	0.201
4	1.955	0.164	0.559	0.456	3.206	0.076	3.753	0.053	1.344	0.248
Low	0.035	0.851	0.012	0.912	0.516	0.474	7.906	0.005	2.103	0.149
Panel B. Dividend-yield portfolios										
	$F_{\widehat{\beta}_{i,M}}$	p-val	$F_{\widehat{\beta}_{i,CF}}$	p-val	$F_{\widehat{\beta}_{i,DR}}$	p-val	$F_{\widehat{\beta}_{i,HML}}$	p-val	$F_{\widehat{\beta}_{i,SMB}}$	p-val
High	1.255	0.265	0.029	0.866	1.378	0.240	1.915	0.166	1.161	0.283
2	3.660	0.058	0.569	0.452	1.505	0.220	6.995	0.008	2.514	0.115
3	0.014	0.907	0.260	0.611	0.580	0.448	7.698	0.006	2.639	0.107
4	0.085	0.771	0.041	0.840	0.886	0.348	5.921	0.015	3.750	0.055
Low	0.310	0.579	0.015	0.903	1.159	0.284	3.427	0.064	2.486	0.117
Panel C. Size portfolios										
	$F_{\widehat{\beta}_{i,M}}$	p-val	$F_{\widehat{\beta}_{i,CF}}$	p-val	$F_{\widehat{\beta}_{i,DR}}$	p-val	$F_{\widehat{\beta}_{i,HML}}$	p-val	$F_{\widehat{\beta}_{i,SMB}}$	p-val
Large	1.459	0.229	0.029	0.865	3.496	0.062	9.381	0.002	2.373	0.126
2	0.869	0.353	0.040	0.842	2.373	0.126	1.565	0.211	2.981	0.087
3	0.353	0.553	0.063	0.802	1.971	0.163	8.467	0.004	3.177	0.077
4	0.145	0.704	0.066	0.797	1.171	0.281	6.508	0.011	3.736	0.055
Small	0.014	0.907	0.588	0.444	0.645	0.423	1.885	0.170	3.503	0.063
Panel D. Price-to-earnings portfolios										
	$F_{\widehat{\beta}_{i,M}}$	p-val	$F_{\widehat{\beta}_{i,CF}}$	p-val	$F_{\widehat{\beta}_{i,DR}}$	p-val	$F_{\widehat{\beta}_{i,HML}}$	p-val	$F_{\widehat{\beta}_{i,SMB}}$	p-val
High	0.712	0.400	0.028	0.867	2.111	0.149	6.921	0.009	1.657	0.200
2	1.345	0.248	0.038	0.845	3.657	0.058	6.884	0.009	5.085	0.026
3	0.224	0.637	0.018	0.894	1.526	0.219	5.685	0.017	0.985	0.323
4	1.251	0.265	0.704	0.403	1.916	0.169	2.042	0.153	1.393	0.240
Low	0.808	0.370	0.070	0.792	2.437	0.118	1.014	0.314	0.775	0.380
Panel E. 3-month momentum										
	$F_{\widehat{\beta}_{i,M}}$	p-val	$F_{\widehat{\beta}_{i,CF}}$	p-val	$F_{\widehat{\beta}_{i,DR}}$	p-val	$F_{\widehat{\beta}_{i,HML}}$	p-val	$F_{\widehat{\beta}_{i,SMB}}$	p-val
Winners	0.148	0.701	0.487	0.486	0.310	0.579	5.713	0.017	0.319	0.573
2	0.116	0.734	0.022	0.881	0.915	0.340	6.413	0.011	1.457	0.229
3	1.121	0.292	0.230	0.632	2.366	0.126	4.375	0.036	1.295	0.257
4	0.607	0.437	0.001	0.973	2.114	0.148	2.489	0.115	5.663	0.019
Losers	3.541	0.062	0.211	0.646	2.511	0.113	2.546	0.111	6.951	0.009

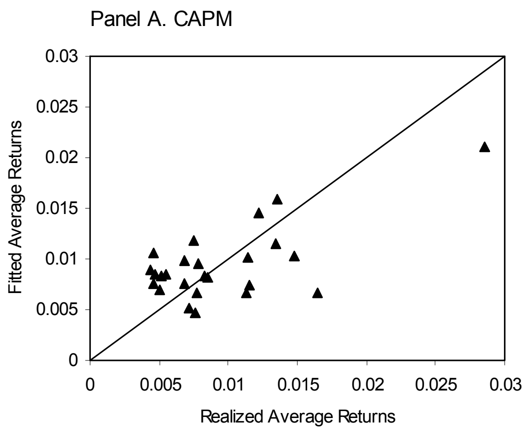
Note. The table presents the values and the p-values of the Chow's breakpoint tests for market (CAPM) betas ($\widehat{\beta}_{i,M}$), cash-flow betas ($\widehat{\beta}_{i,CF}$), discount–rate betas ($\widehat{\beta}_{i,DR}$), HML betas ($\widehat{\beta}_{i,HML}$) and SMB betas ($\widehat{\beta}_{i,SMB}$) of the book-to-market, dividend yield, size, price-to-earnings and 3-month momentum sorted portfolios, respectively. The estimation period spans from June 1991 to May 2003 and the breakpoint was set to January 2001.

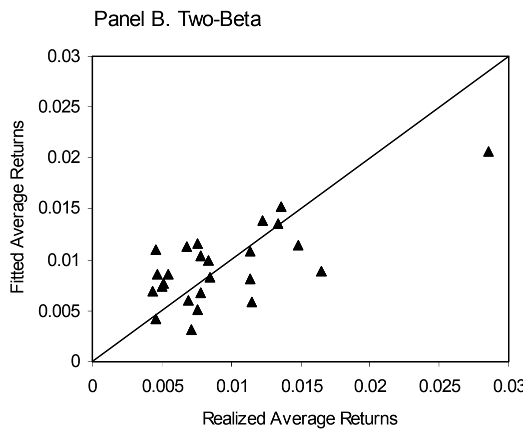
Table 6.

Cross-sectional asset pricing tests

	CAPM	Two-Beta	CF	DR	Fama-French	All
λ_0	-0.005	-0.0139^{**}	-0.0107	0.0026	0.0088	0.0025
	(0.0048)	(0.0062)	(0.0065)	(0.0031)	(0.0072)	(0.0070)
λ_M	0.0115^{***}				-0.0015	
	(0.0038)				(0.0064)	
λ_{CF}		0.0274^{***}	0.0320^{***}			0.0166^{**}
		(0.0093)	(0.0113)			(0.0076)
λ_{DR}		0.0096^{**}		0.0107^{**}		-0.0132
		(0.0043)		(0.0051)		(0.0082)
λ_{HML}					-0.0066^{*}	-0.0051
					(0.0036)	(0.0032)
λ_{SMB}					0.0098^{**}	0.0159^{***}
					(0.0044)	(0.0054)
adj.-R^2	42.2%	46.6%	21.5%	30.4%	48.5%	62.6%
γ		2.8572^{***}				
		(0.1612)				
λ		0.0096				
		(0.0043)				

Note. The table presents the results of the cross-sectional asset pricing regressions using the book-to-market, dividend yield, size, price-to-earnings and 3-month momentum sorted portfolios, respectively. It reports the estimates of the beta-prices of risk (λs), their standard errors in parentheses and the adj.-R^2 of the regression. The specification of the CAPM, the Two-Beta and the Fama-French (FF) model are given (14), (13) and (15), respectively. The CF and DR corresponds to (13) where only $\widehat{\beta}_{i,CF}$ and $\widehat{\beta}_{i,DR}$ were used in the estimation. γ and λ are the estimates of the coefficient of relative risk aversion and the cash-flow beta price of risk when the restricted two-beta model is estimated (equation (16)). *, ** and *** denote significance at 10%, 5% and 1%, respectively.





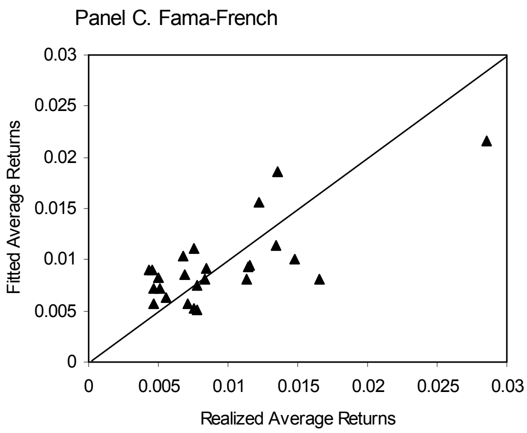


Figure 1.

Realized versus fitted average returns

Realized versus fitted monthly average returns on the 25 book-to-market, dividend yield, size, price-to-earnings and 3-month momentum sorted portfolios. Panel A: unconditional CAPM (equation (14)), Panel B: two-beta model (equation (13)), Panel C: Fama and French (1993) model (equation (15)). The estimation period spans from June 1991 to May 2003.