# HOLDING PERIOD AND CROSS-SECTIONAL STOCK RETURNS: EVIDENCE FROM TAIWAN

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**ABSTRACT**

*This paper employs a hybrid approach that combines an adapted version of Fama-MacBeth two-pass regression with Engle-Granger cointegration test to characterize the relationship between expected stock returns and systematic risks with diverse investment horizons. We find no evidence supporting a positive relationship between the market beta and return for various investment horizons. The book-to-market effect is sensitive to the investment horizon. We find a size effect for diverse investment horizons in period from 1986 to 1993. However, the size effect disappears in the subsequent period.*

**JEL**: G11; G12

**KEYWORDS**: Asset pricing model, cointegration, holding period

## INTRODUCTION

The relationship between expected return and systematic risk is still an important issue in the academic field and to practitioners, because consensus does not exist on how risk factors affect an asset's expected return. There is a long history of exploration on the issue. Among others, Black, Jensen, and Scholes (1972), and Fama and Macbeth (1973) found that market beta is the only factor in explaining an asset's return. Fama and French (1992, 1993) and Carhart (1997) provided evidences that factors other than market beta play an important role in explaining an asset's return. However, Kothari, Shanken, and Sloan (1995) claimed that market beta is still alive, and the relation between expected return and book-to-market is seriously exaggerated by survivor bias. The studies are ample and provide insights on the relationship between expected return and factors' risk, but they fail to consider the effect of investment horizon.

Levy and Spector (1996) stated that the investment horizon plays a crucial role in determining the optimum composition of an investment portfolio. Among others, Gunthorpe and Levy (1994) found that portfolio composition changes dramatically and systematically with changes in investment horizon. This is because changes in investment horizon can affect both the risk and return of the portfolio. Hawawini (1983) and Levy and Cohen (1998) provided evidence that stock risks change as the return interval is lengthened. The results imply that investors need to construct and evaluate their portfolios in a manner consistent with their planned investment horizons. However, they only considered the choice of return interval, for example, daily versus weekly, or monthly versus annually, and then examined the possible impact of the return interval on beta. Although Levy and Samuelson (1992) proved that the Capital Asset Pricing Model (CAPM) will hold in some cases under a diverse investment horizons, few studies have provided empirical evidence on how the investment horizon influences the relationship between systematic risks and expected return.

In this study, we modify the Fama-MacBeth approach to examine the effect of investment horizon on the return-risk relationship in the Taiwan Stock Exchange (TSE). To explore the investment horizon effect, we assume that an investor constructs a portfolio based on estimated betas, holds the portfolio for a number of months, and then sells it. By varying the holding period, we can explore the effect of investment horizon on the risk-return relationship. Specifically, we first estimate betas for each month, and then use the average monthly returns for various investment horizons as a substitute for the next one-month return in the second-pass Fama-MacBeth cross-sectional regression. One question is whether use of the average return is justified, given that there could be unit roots in the average return, which is calculated by an overlapping technique. To tackle this spurious regression problem, we apply Engle and Granger's (1987) cointegration technique to determine whether there is a long memory relationship between the risk premium and average excess return on each expected factor. If a factor's risk can consistently explain the expected return, the risk premium, which is the slope coefficient obtained from running the Fama-MacBeth cross-sectional regression, should be equal to the expected excess return.

Our results showed that the relationship between the market beta and expected return is weak, no matter how long we hold the portfolio. This result contradicts the work of Levy and Samuelson (1992), which states that the CAPM holds for diverse investment horizons. After analyzing data for 12-, 24-, and 36-month holding periods, we found that the book-to-market value could explain the cross-sectional variation in expected stock returns only for a 24-month holding period. On the other hand, the size effect was significant no matter how long we held the portfolio. The results imply that the ability of the beta for a specific risk's factor to explain the cross-sectional variation in expected return varies with the investment period. The results are same as the work of Fama and French (1992, 1993) for American equity market, the work of Chui and Wei (1997) for Asian equity market, and the work of Ma and Shaw (1990) for the Taiwan equity market. However, these results have different implications, because their results can only be used for investors with short-run investment horizon. These results can help investors understand whether estimated betas are useful for predicting stock returns given different investment horizons. Meanwhile, the results also imply that investors with different investment horizons should have different equity allocations.

The next section contains a literature review. Methodology and data are outlined in Section 3 and 4. Section 5 provides an analysis of empirical results. Finally, some concluding remarks are provided.

## LITERATURE REVIEW

The CAPM was developed by Sharpe (1964) and Treynor (1961) to determine an asset's price for risk. The only risk is market beta, which is covariance function of asset's return with market portfolio. Other factors play no role in determining an asset's price. The CAPM was tested by many researchers, including Friend and Blume (1970), Black, Jensen, and Scholes (1972), Miller and Scholes (1972), Fama and Macbeth (1973), Basu (1977), Banz (1981), and Stambaugh (1982), among others. Most empirical studies apply Fama and MacBeth (1973) two-pass regression approach to determine whether a specific risk's factor explains the cross-sectional variations in expected stock returns. For example, Fama and French (1992) provided evidences that the ratio of a firm's book-to-market value and firm size explain American stock returns far better than market beta. Chui and Wei (1998) found the same evidence for this phenomenon in five Pacific-Basin emerging markets. In particular, they applied monthly returns on listed common stocks and estimated market beta using a 60-month holding period. All of the stocks were then ranked by beta and placed into a certain number of portfolios. Next, the betas and returns of these portfolios are estimated for every rolling 60-month holding period. They projected the next one-month realized portfolio returns on the estimated portfolio betas. The design of using a portfolio instead of individual stocks can avoid measurement errors. Finally, the researchers use ordinary t-test to examine the relationship between returns and betas.

Most early studies found evidence in supporting the CAPM, including Friend and Blume (1970), Black, Jensen, and Scholes (1972), Miller and Scholes (1972), Fama and Macbeth (1973). However, other researchers found that factors other than market beta can explain a large portion of asset returns. Basu (1977) stated that low price/earnings portfolios have higher return than expected by CAPM. Basz (1981) found that firm size can play an important role in explaining firm's returns. Keim (1983) identified seasonality in stock returns. Fama and French (1992, 1993) provided evidence that market, size, and book-to-market, are the only three equity's risk's factors. Moreover, controlling the size factor, the market beta lost its role in explaining an equity's expected return. Davis, Fama, and French (2000) also presented evidence that the positive relation between average return and book-to-market is strong. However, Kothari, Shanken, and Sloan (1995) challenged the results of Fama and French (1992, 1993) by claiming that the relation between expected return and book-to-market is seriously exaggerated by survivor bias. He et al. (1996) also found that size and book-to-market can only explain a small portion of the stock expected return. Downs and Ingram (2000) showed that average returns relate positively with market beta but not with size. In the bond market, Gebhardt, Hvidkjaer, and Swaminathan (2005) found that default beta is significantly related to the cross-section of average return.

Studies on the return-risk relationship are extensive, but most studies didn't consider the effect of investment horizons. Samuelson (1969, 1990, and 1994) proved that a optimal asset allocation is independent of the investment horizon. However, the work of Lloyd and Haney (1980) doesn't coincide with the argument of Samuelson. Lloyd and Haney pointed out that the volatility of a portfolio's value can be reduced by lengthening the holding period. This is the concept of time diversification. Gunthorpe and Levy (1994) and Levy and Spector (1996) found that optimal weights of risky asset are changed as investment horizon lengthens. Merrill and Thorley (1996) and Levy and Cohen (1998) also proved that lengthening the holding period could reduce risk. Moreover, Hawawini (1983) and Levy and Cohen (1998) found that stock's market betas change as the holding period is lengthened. The results imply that the relationship between an asset's returns and risk need further exploration for different holding periods.

Many studies have explored the return-risk relationship in the TSE. Among others, Ma and Shaw (1990) provided evidence that firm size and earnings to price can explaining expected returns. Huang (1997) found an inverse relationship between return and market beta. However, Chen (2003) indicated that the relationship between return and market beta was statistically significant. Sheu, Wu, and Ku (1998) and Ku (2005) suggested that market beta, volume, sales to price and momentum are better risk's factors in explaining asset's returns. Jan and Shiu (2008) also provided that an investor can be better off by holding risky assets over longer investment horizon. However, none of these studies considered the influence of investment horizons on the return-risk relationship in the TSE.

## METHODOLODY

The motivation of our analysis is to determine whether the risk's factors are priced with a diverse investment horizon. The empirical models we study include the CAPM, 3-factor asset pricing model of Fama and French (1993), and 4-factor asset pricing model of Carhart (1997). All of these models are of the form

$$E(R_{it+1}) = \sum_{j=1}^{p} \alpha_{j,t+1} \beta_i^j , \qquad (1)$$

where $E(\bullet)$ is the expectation operator; $R_{it+1}$ denotes the excess return, the difference between the return on portfolio $i$ and the risk-free return $R_f$ at time $t+1$; $\beta_i^j = \text{cov}(R_i, R_j)/\text{var}(R_j)$ is the beta for factor portfolio $j$ and $\alpha_{jt+1}$ is the corresponding expected risk premium at time $t+1$. The value of $p$ determines the model represented in Equation (1), as follows. If $p=1$, then we have the CAPM, and the risk's factor is the market portfolio. If $p=3$, then we have the 3-factor asset pricing model, and the risk's factors are the market portfolio, size portfolio, and book-to-market ratio portfolio. If $p=4$, then we have the 4-factor asset pricing model, where the first three factors are identical to those in the 3-factor model, and the fourth factor is a one-year momentum portfolio. In all of the above models, the factors are systematic risks since idiosyncratic risks can be diversified away.

We modify the Fama and Macbeth's (1973) two-pass regression approach to investigate the relationship between the betas and expected returns on the factors. The original version of this approach involves three steps. First, a time series regression is used to estimate the betas $\beta_i^j$ based on the previous 60 monthly returns for every portfolio. Second, a cross-sectional regression is performed on the expected returns $E(R_{it+1})$ for each month based on the betas $\beta_{it}^j$ obtained at time $t$. Using the technique introduced by Fama and MacBeth and employed by later researchers, the next one-month realized excess return is used as a substitute for the expected excess return. That is,

$$R_{it+1} = \alpha_{0,t+1} + \sum_{j=1}^{p} \alpha_{j,t+1} \beta_{i,t}^j + \varepsilon_{t+1} \ , \tag{2}$$

where $\varepsilon_{t+1}$ is an error term and $\alpha_{0,t+1}$ is a parameter associated with an intercept. Finally, the conventional $t$-statistics is used to test whether the time-series slope coefficients $\alpha_{j,t+1}(\alpha_{0,t+1})$ from the cross-sectional regression are different form zero at a conventional significance level.

A limitation of the original Fama-MacBeth approach is that it is not applicable to long holding periods. Each investor has her or his own planned investment horizon. For instance, some investors may prefer to invest for 12 months, while others may want to invest for 24 months. If we construct a portfolio based on the betas of the estimated factors and hold it more than one month, the next one-month return may not shed light on the relationship between the betas and the expected return. To more accurately characterize the effect of a long investment horizon, we substitute the average monthly return for the next one-month return in the second pass regression of Fama-MacBeth. This approach is similar to the work of Black, *et al.* (1972). However, they only regressed the average returns to the betas once. That is, in place of $R_{it+1}$ in equation (2),

$$\overline{R}_{it+k} = \alpha_{0,t+k} + \sum_{j=1}^{p} \alpha_{j,t+k} \beta_{i,t}^j + \varepsilon_{t+k} \ , \tag{3}$$

where $\overline{R}_{it+k}$ is the $k$-month average excess return on portfolio $i$ from month $t$ to month $t+k$. In this study, $k$ is either 12, 24, or 36 months. To increase the robustness of our sample, we employ overlapping holding periods. Specifically, if $k=12$, the first monthly average excess return is calculated for the period from January 1986 to December 1986, the second is for the period from February 1986 to January 1987, and so on. Therefore, the data for monthly average return represents overlapping returns. The null hypotheses of the Fama-MacBeth approach are that the average slope coefficients are zero. If the null hypothesis is rejected for the beta of a given factor, then this factor has power to explain the cross-sectional variation in expected returns. However, the above results may produce spurious problems, because we use the overlapping approach to compute the monthly average returns, and the time-series average returns could have a unit root. In this case, the time-series slope coefficients estimated from Equation (3) would not be independent, and furthermore, could also have unit roots. The coefficient of correlation between the time-series slope coefficients (see Appendix) could approach unity. If the regressors which come from estimated time-series slope coefficients were not independent or had unit roots, according to the work of Fuller (1996), the conventional $t$-statistics could be biased due to the misestimated variance of estimator.

When we use long-run average returns, most time-series slope coefficients and moving average returns on factor portfolios cannot reject the hypothesis of a unit root. Nevertheless, if Equation (3) correctly describes the cross-section of average returns, then the slope coefficient is the corresponding factor's risk premium, (See for example, the work of Fama and French, 1993) which is equal to the expected excess return on the corresponding factor portfolio. That is,

$$\alpha_{jt} = E(R_{jt}) \ , \tag{4}$$

where $R_{jt}$ denotes the excess return on factor portfolio $j$ at time $t$. Therefore, the time-series slope coefficients and expected returns on the factor portfolio must have a long memory relationship. To determine whether such a long memory relationship holds, we use the average monthly return to estimate the expected return and apply a test by Engle and Granger (1987) to examine whether they are cointegrated. The test is performed as follows. First, we run the following regression:

$$\alpha_{jt+k} = \gamma_0 + \gamma_1 \overline{R}_{jt+k} + \xi_{jt+k}, \quad (5)$$

where $\overline{R}_{jt+k}$ is the $k$-month average excess return on factor portfolio $j$ from month $t$ to month $t+k$, $\alpha_{jt+k}$ is the corresponding cross-sectional slope coefficient obtained from Equation (3), $\gamma_0$ and $\gamma_1$ are the cointegration vector parameters, and $\xi_{jt+k}$ is an error term. If both the time-series slope coefficients and the moving average excess returns have unit roots, rejection by the error term of the null hypothesis (unit root existence) implies that they are cointegrated. We also run the reversal regression, which regresses the average excess return on the time-series slope coefficient. The results are very similar to those of Equation (5). To save the space, we do not present the results here. However, they are available upon request to the author.

Data

To avoid the survival bias critique described by Brown et al. (1995), we include all stocks contained in the Taiwan Economic Journal (TEJ) from July 1982 to December 2005. The sample contains 602 stocks and 282 monthly observations. All the stocks in the TSE are included until they are delisted, even if a firm's book values are negative. This method is designed to avoid the problem of survival bias. We group the stocks into portfolios based on two criteria. First, we classify the stocks into 5 size categories based on their market values. Second, we classify each size category into 5 portfolios based on the book-to-market ratio of each firm. The result is 25 portfolios $P_{i,j}$, where $1 \le i \le 5$ represents size and $1 \le j \le 5$ represents book-to-market-ratio. Many studies have used this system of five size and book-to-market categories to examine the accuracy of a specific asset pricing model, because the resulting 25 portfolios are very hard to price correctly. See, for example, Fama and French (1993), and Daniel and Titman (1997).

In addition to the 25 portfolios, we include four special portfolios, each of which mimics a given risk's factor. The first is the Taiwan Weighted Index (TWI), which we use as the market portfolio. The risk-free rate is the one-month deposit rate at the First Commercial Bank. In an approach introduced by Fama and French (1993) and later used by Carhart (1997), we create three zero-investment portfolios, which measure the effect on simple average returns of three risk's factors. The first is size portfolio (SMB), which consists of the difference between the simple average returns of the smallest 40 percent market value portfolios and the simple average returns of the highest 40 percent market value portfolios. The second is value-growth portfolio (HML), which is the spread between the simple average returns of the highest 40 percent book-to–market ratio portfolios and the simple average returns of the lowest 40 percent book-to-market ratio portfolios. The last is one-year momentum portfolio (PR1YR), which is the difference between the simple average returns of the firms with the highest 40 percent lagged one month return and the lowest 40 percent lagged one month return.

Table 1 presents summary statistics for the 25 portfolios, sorted by size and book-to-market ratio. The statistics include the average monthly returns, the standard deviation of the monthly returns (shown in parentheses), the market value, and the book-to-market ratio for each portfolio. Panel A shows the market value weighted average monthly returns. Small market value portfolios have relatively high average returns. Though a high book-to-market ratio does not guarantee higher average returns, on average the portfolios $P_{\bullet,1}$ and $P_{\bullet,5}$ with the lowest and highest book-to-market ratios have higher returns. On the other hand, the price of a high average return is generally a high standard deviation, as shown by the standard deviation figures in parentheses. Panels B and C show the market value and book-to-market ratio sample averages for each portfolio. In Panel B, the sample average for portfolio $P_{1,1}$ in the upper left hand corner is negative because we included stocks with a negative book value until they were delisted from the TSE. The other 24 portfolios show book-to-market ratios of less than one, with the exception of the high book-to-market ratio portfolios $P_{1,5}$, $P_{2,5}$ and $P_{4,5}$. The phenomena are very similar to other markets, such as the U.S. stock market documented by Fama and French (1992), the Japanese stock market documented by Jagannathan et al. (1998), the Pacific-Basin stock markets documented by Chui and Wei (1998), and international stock markets documented by Bauman et al. (1998) and Fama and French (1998).

Table 2 shows summary statistics for the four risk factor-mimicking portfolios. The average monthly returns vary from 0.556 percent for the book-to-market ratio portfolio to 1.858 percent for the one-year momentum portfolio. Interestingly, the one-year momentum portfolio has the lowest variation and the book-to-market ratio portfolio has the highest variation. Table 2 also presents the coefficients for risk factor cross correlations. The market factor has a low correlation with the other factors, suggesting that inclusion of factors other than the market factor may increase a model's power to explain the cross-section expected returns. By contrast, size factor displays a 0.451 correlation with book-to-market portfolio.

## EMPIRICAL RESULTS

Table 3 presents the results of the Augmented Dickey-Fuller (ADF) Unit Root Tests. Panel A shows the results for the time-series slope coefficients estimated from Equation (3). The hypothesis of unit root existence was not rejected in most cases. There were rejections for three of the 12-month investment horizons: the market factor in the 3-, and 4-factor asset pricing models and the one-year momentum factor in the 4-factor asset pricing model. However, when the investment horizon rose above 12 months, the estimates of the time-series slope coefficients all can't reject the existences of unit root. Panel B displays the results for the moving average monthly returns of the factor portfolios. None of the tests rejected the unit root hypothesis.

Table 4 displays the results of the Engle-Granger cointegration tests. We only report the results when the unit root hypothesis has not been rejected for both the time-series slope coefficients and the moving average excess returns. When one variable is stationary and the other is not, the linear combination is nonstationary. Therefore, use of the cointegration test would be meaningless.

The results for the full sample are shown in Panel A, while those for the sub-sample are shown in Panel B and C. The first portion of Panel A presents the results for the CAPM, which show that we cannot reject the null hypotheses of unit root existence for the error terms from running Equation (5), implying that the time-series slope coefficient and the moving average excess returns of the market portfolio are not cointegrated. As a result, the market factor doesn't have the power to predict the expected return, implying that the CAPM cannot describe the return-risk relationship cross-sectional for a long investment horizon.

Table 1: Summary Statistics on the 25 Size and Book-to-Market Sorted Portfolios

| Panel A: Average Monthly Return | | | | | |
|---|---|---|---|---|---|
| Portfolio | Book-to-Market Ratio | | | | |
| Size | Lowest | 2 | 3 | 4 | Highest |
| Smallest | 4.120 | 4.122 | 3.086 | 3.902 | 4.185 |
| | (17.474) | (16.768) | (15.871) | (17.102) | (17.743) |
| 2 | 3.401 | 2.824 | 2.911 | 3.093 | 2.811 |
| | (14.737) | (14.020) | (14.011) | (14.892) | (15.262) |
| 3 | 3.254 | 2.697 | 2.202 | 2.564 | 2.661 |
| | (14.054) | (13.759) | (12.516) | (13.745) | (15.100) |
| 4 | 2.748 | 2.421 | 1.612 | 1.914 | 2.756 |
| | (13.113) | (12.484) | (12.238) | (11.920) | (13.995) |
| Largest | 2.797 | 2.095 | 2.279 | 2.313 | 2.498 |
| | (14.838) | (13.196) | (11.653) | (12.776) | (13.900) |
| **Panel B: Average Market Value (Millions of Dollars)** | | | | | |
| Smallest | 2256.1 | 1920.6 | 1772.5 | 1770.6 | 1539.8 |
| 2 | 4553.3 | 3977.4 | 3388.8 | 3496.6 | 3355.5 |
| 3 | 10518.6 | 6443.5 | 7066.6 | 6113.5 | 5925.7 |
| 4 | 11093.8 | 10160.5 | 10673.2 | 10404.3 | 10179.7 |
| Largest | 78249.9 | 46192.1 | 38861.0 | 35818.5 | 43316.5 |
| **Panel C: Average Book-to-Market Ratio** | | | | | |
| Smallest | -1.099 | 0.524 | 0.648 | 0.848 | 1.425 |
| 2 | 0.310 | 0.475 | 0.617 | 0.805 | 1.179 |
| 3 | 0.270 | 0.414 | 0.520 | 0.670 | 0.960 |
| 4 | 0.248 | 0.389 | 0.512 | 0.661 | 1.022 |
| Largest | 0.185 | 0.303 | 0.420 | 0.557 | 0.856 |

*The data span from July 1982 to December 2005. The standard deviations appear in parentheses. The average returns and standard deviations are represented as percentages.*

Table 2: Summary Statistics for the Risk Factor Portfolios

| Factor Portfolio | Average Return | Standard Deviation | Cross Correlation | | | |
|---|---|---|---|---|---|---|
| | | | TWI | SMB | HML | PR1YR |
| TWI | 1.618 | 11.867 | 1.00 | | | |
| SMB | 1.620 | 10.409 | -0.085 | 1.00 | | |
| HML | 0.556 | 12.569 | -0.127 | 0.451 | 1.00 | |
| PR1YR | 1.858 | 6.207 | 0.066 | 0.023 | 0.164 | 1.00 |

*TWI is return on the Taiwan value-weighted index. SMB, HML, and PR1YR are monthly returns on the value-weighted, zero-investment, factor-mimicking portfolios for size, book-to-market ratio, and one-year momentum in stock returns. The average returns and standard deviations are given as percentages.*

The full sample results for the 3-factor asset-pricing model are shown in the middle portion of Panel A. The time-series slope coefficients and the moving average returns of the market portfolio are not cointegrated in either the 24- or 36-month investment horizon. The cointegration vector estimates in the 24-month investment horizon are negative, indicating an inverse relationship between return and market beta. When we increase the investment horizon to 36 months, the cointegration vector estimates become positive. On the other hand, the time-series SMB slope coefficients and moving average excess returns are cointegrated at the 5% significance level. Moreover, the cointegration vector estimates are positive and very close to one. The time-series HML slope coefficient is only cointegrated with the moving average excess return in the 24-month investment horizon. The cointegration vector estimates are also positive and close to one.

Table 3: Augmented Dickey-Fuller Unit Root Tests

| Panel A: Cross-Sectional Slope Coefficients for Fama-Macbeth Regression | | | | | |
|---|---|---|---|---|---|
| Model | Horizon | $\alpha_{TWI}$ | $\alpha_{SMB}$ | $\alpha_{HML}$ | $\alpha_{PR1YR}$ |
| CAPM | 12 | -3.061 | --- | --- | --- |
| | 24 | -1.282 | --- | --- | --- |
| | 36 | -1.202 | --- | --- | --- |
| 3-factor asset pricing model | 12 | -4.660* | -1.962 | -2.619 | --- |
| | 24 | -3.353 | -1.719 | -2.189 | --- |
| | 36 | -2.857 | -1.710 | -2.108 | --- |
| 4-factor asset pricing model | 12 | -4.550* | -1.9042 | -2.582 | -3.523* |
| | 24 | -3.342 | -1.581 | -1.569 | -2.907 |
| | 36 | -2.140 | -1.534 | -2.142 | -2.218 |
| **Panel B: Time-Series Excess Returns for the Factor Portfolios** | | | | | |
| Horizon | | TWI | SMB | HML | PR1YR. |
| 12 | | -3.110 | -2.448 | -1.982 | -1.322 |
| 24 | | -3.392 | -1.516 | -1.924 | -1.507 |
| 36 | | -3.111 | -2.435 | -1.978 | -1.184 |

*The unit root tests are conducted with both constant and trend. We don't report the lags, which are based on BIC criteria, and are available upon request. * denotes the 5% significance level. The horizon is the number of months the portfolio is held.*

The bottom portion of Panel A displays the full sample results for the 4-factor asset pricing model. Again, the time-series slope coefficient for the market portfolio does not cointegrate with the moving average excess return. On the contrary, the time-series slope coefficients and moving average returns of the SMB are cointegrated for diverse investment horizons. Furthermore, the cointegration vector estimates are also close to one. The slope coefficient and moving average return for the HML portfolio are also cointegrated for the 12- and 24-month investment horizons. The moving average return and the time-series slope coefficient for the PR1PR portfolio are not cointegrated. The cointegration vector estimates are negative, indicating that the beta of the one-year momentum portfolio cannot explain the cross-section of the average returns in the TSE.

Panel B and C present the results for the two sub-samples. The first sample is from July 1986 to December 1993 and the second is from January 1994 to December 2005. The sub-samples are designed to test whether our results are robust in different periods. The results of first sample are same as the full sample, implying a strong size effect for the long investment horizon in the period from July 1986 to December 1993. However, the strong size effect disappears in the second sub-sample, implying that we can't use the size premium to predict the long horizon expected returns. These results are similar to the works of Ferson and Harvey (1999), which showed that the size effect couldn't pass the robust diagnosis. In summary, our results imply that size has power in explaining the return-risk relationship in period from July 1986 to December 1993. The effect of book-to-market ratio on a stock's expected return varies with the investment horizon. The market factor and the one-year return momentum do not provide explanatory power for the expected return cross-section in the TSE.

Table 4: Engle-Granger Cointegration Tests

**Panel A: Full Sample (1982/7~2005/12)**

| Model | | TWI | SMB | HML | PR1YR |
|---|---|---|---|---|---|
| CAPM | 12 | -2.407<br>(-0.027 –0.528) | --- | --- | --- |
| | 24 | -1.249<br>(-0.023 -0.467) | --- | --- | --- |
| | 36 | -2.254<br>(-0.015 1.393) | --- | --- | --- |
| 3-factor asset pricing model | 12 | NA. | -4.250*<br>(-0.001 0.856) | -5.242<br>(-0.006 1.085) | --- |
| | 24 | -3.442<br>(-0.001 -0.255) | -3.595*<br>(-0.005 0.976) | -5.008*<br>(-0.007 1.010) | --- |
| | 36 | -2.958<br>(-0.000 0.122) | -3.991*<br>(0.002 0.725) | -3.208<br>(-0.006 0.408) | --- |
| 4-factor asset pricing model | 12 | NA. | -4.215*<br>(-0.002 0.915) | -3.728*<br>(-0.006 0.938) | NA. |
| | 24 | -3.353<br>(-0.002 –0.236) | -4.097*<br>(-0.005 1.011) | -3.908*<br>(-0.006 0.787) | -3.352<br>(0.002 -0.276) |
| | 36 | -2.481<br>(-0.002 0.381) | -3.465*<br>(0.002 0.722) | -2.320<br>(-0.004 0.291) | -2.353<br>(-0.002 -0.085) |

**Panel B: Sub-sample (1982/7~1993/12)**

| Model | | TWI | SMB | HML | PR1YR |
|---|---|---|---|---|---|
| CAPM | 12 | -3.389 | --- | --- | --- |
| | 24 | -1.987 | --- | --- | --- |
| | 36 | -3.554* | --- | --- | --- |
| 3-factor asset pricing model | 12 | NA | -4.413* | -4.422* | --- |
| | 24 | -2.635 | -3.814* | -5.421* | --- |
| | 36 | -2.459 | -5.025* | -2.451 | --- |
| 4-factor asset pricing model | 12 | NA | -4.683* | -2.859 | NA. |
| | 24 | -2.813 | -4.243* | -4.038* | -2.762 |
| | 36 | -2.188 | -4.763* | -2.352 | -2.208 |

**Panel C: Sub-sample (1994/1~2005/12)**

| Model | | TWI | SMB | HML | PR1YR |
|---|---|---|---|---|---|
| CAPM | 12 | -0.841 | --- | --- | --- |
| | 24 | -0.062 | --- | --- | --- |
| | 36 | -0.901 | --- | --- | --- |
| 3-factor asset pricing model | 12 | NA | -2.967 | -3.164 | --- |
| | 24 | -2.430 | -1.547 | -2.104 | --- |
| | 36 | -2.056 | -1.026 | -1.152 | --- |
| 4-factor asset pricing model | 12 | NA | -2.857 | -2.775 | NA. |
| | 24 | -2.688 | -1.486 | -1.839 | -2.378 |
| | 36 | -1.800 | -1.071 | -1.392 | -1.693 |

*The cointegration vectors are shown in parentheses, where the former value represents the constant. The conintegration vectors are not shown in panel B and C to save the space. NA signifies that one variable is stationary and the other is nonstationary. Therefore, use of the Enger-Granger test would be meaningless. * denotes the 5% significance level.*

## CONCLUSION

This paper modifies Fama-MacBeth cross-sectional approach to examine the relationship between expected stock returns and systematic factor's risks with diverse investment horizons in the Taiwan Stock Exchange. If a specific risk's factor is important to describe the cross-sectional variation of expected returns, the estimated risk premium should equal its expected risk premium. That is, these two time-series data should cointegrate with each other. Therefore, we can examine whether a specific risk's factor has power in explaining return-risk relationship by use of the Engle-Granger test. We find no evidence in supporting a positive relationship between market beta and return. The size effect with various investment horizons just exists in the period from July 1986 to December 1993. The book-to-market effect is sensitive to the investment horizon. Meanwhile, including the return on one-year momentum does not improve the ability to explain the cross-section of average returns in the TSE. Our results show that the effect of a specific risk's factor on expected return depends on the investment horizon, which implies that investors with different investment horizons should have different equity allocations. Our modified method links the expected stock returns to candidated factors' risks in the long investment horizon. However, only four factors are examined. Moreover, the four factors are tested separately. Some comovements between factors may be ignored. Therefore, including more candidate factors and using a multivariate cointegrated method would improve the empirical evidences. We leave these issues for further research.

## APPENDIX

Equations (2) and (3) can be rewritten as:

$$R_{t+1} = \beta_t \alpha_{t+1} + \varepsilon_{t+1}, \quad \text{(A1)}$$

$$\overline{R}_{t+k} = \beta_t \delta_{t+k} + u_{t+k}, \quad \text{(A2)}$$

where $\beta_t$ is the factor's beta vector, $\alpha_{t+1}$ and $\delta_{t+k}$ are the estimated slope coefficient's vectors, and $\varepsilon_{t+1}$ is the unexpected return with mean $E\{\varepsilon_{t+1}) = 0$, variance $\sigma^2(\varepsilon_{t+1}) = \sigma^2$, and covariance $Cov(\varepsilon_t, \varepsilon_{t+1}) = 0$. From (A1) and (A2), we can show that the surprise return $u_{t+k}$ is the average of the error terms; i.e.,

$$u_{t+k} = \frac{1}{k}\sum_{i=1}^{k} \varepsilon_{t+i}. \quad \text{(A3)}$$

It follows that time-series surprise return $u_{t+k}$ is autocorrelated. In particular, the covariance between $u_{t+k}$ and $u_{t+k+1}$ equals

$$Cov(u_{t+k}u_{t+k+1}) = \frac{k-1}{k}\sigma^2. \quad \text{(A4)}$$

Under the assumption that the regressor $\beta_t$ is uncorrelated with $u_{t+k}$, we can use an ordinary least square method to estimate the parameters $\delta_{t+k}$ and their variance

$$\hat{\delta}_{t+k} = (\beta_t'\beta_t)^{-1}\beta_t'\overline{R}_{t+k}, \quad \text{(A5)}$$

$$Var(\hat{\delta}_{t+k}) = \frac{\sigma^2}{k}(\beta_t'\beta_t)^{-1}. \quad \text{(A6)}$$

It can also be shown that the covariance and the coefficient of correlation between $\hat{\delta}_{t+k}$ and $\hat{\delta}_{t+k+1}$ are

$$Cov(\hat{\delta}_{t+k}\hat{\delta}_{t+k+1}) = \frac{(k-1)\sigma^2}{k^2}(\beta_t'\beta_t)^{-1}\beta_t'\beta_{t+1}(\beta_{t+1}'\beta_{t+1})^{-1}, \quad \text{(A7)}$$

$$\rho(\hat{\delta}_{t+k}\hat{\delta}_{t+k+1}) = \frac{k-1}{k}(\beta_t'\beta_t)^{-1/2}\beta_t'\beta_{t+1}(\beta_{t+1}'\beta_{t+1})^{-1/2}. \quad \text{(A8)}$$

The coefficient of correlation between $\hat{\delta}_{t+k}$ and $\hat{\delta}_{t+k+1}$ may approach unity when the investment horizon $k$ is large.

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