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Fama-French Model Explanations of the Stock Market Anomaly

Ruzita Abdul-Rahim[†]
Universiti Kebangsaan Malaysia

Abstract: The main purpose of this study is to test the ability of the Fama-French model as opposed to CAPM in explaining seasonality effect. The study employed multiple time-series regressions on monthly returns data of nine double-sorted ME/BM portfolios constructed from 220 to 500 stocks listed on Bursa Malaysia over the period of 1985:01 to 2005:12. The results indicate that when CAPM is used to explain portfolio returns, evidence of February effect persist in portfolios that are composed of stocks of small and distressed firms. Nonetheless, the seasonality effect disappears entirely when the Fama-French model is used. Overall, this finding lends strong support for the hypotheses that seasonality effect that has seemingly been persistent in Bursa Malaysia is merely due to model inadequacy.

Keywords: Seasonality effect, February effect, CAPM, Fama-French Model, Malaysian equity market

1. Introduction

Regardless of how elusive the goal is, the ability to predict security returns is undebatably an ever alluring quest among academicians and practitioners in finance. There has not been proof of declining interest even more than a decade after Fama (1991) concluded that the stock market is informationally efficient in the semi-strong form. In fact, January effect is probably still the most closely examined anomaly of efficient market hypothesis since the last two decades implying deep interest even in weak-form market efficiency. This phenomenon is particularly true for major capital markets like the New York Stock Exchange (NYSE) where studies on this issue are both voluminous and lenient towards supporting the January anomaly (cf. Rozeff and Kinney 1976; Keim 1983; Lakanishok and Smidt 1988; Haugen and Jorion 1996). Less rigorous studies in other countries too (Gultekin and Gultekin 1983; Kato and Schallheim 1985; Yong 1991; Pandey 2002; Yakob *et al*. 2005; Abdul-Rahim and Harjito 2006; Abdul-Rahim *et al*. 2006) also indicate evidence of January effect. Such urge to understand stock price behaviour is justified given the vital role a stock market plays in any economy. Major economic indicators are related to stock market indicators and Bursa Malaysia (formerly known as Kuala Lumpur Stock Exchange, KLSE) is not an exception. Despite being relatively new and less sophisticated compared to those in developed countries, Bursa Malaysia deals with security transactions worth on average RM0.8 billion per day or RM200 billion per year (Bank Negara Malaysia 2005). This is equivalent to about 50 per cent of the country's Gross National Product. On that grounds, understanding the anomalous behaviour of prices of stocks traded in the exchange is rational and contributes towards the efficiency of the equity market and economy.

Corresponding author: E-mail: *ruzitaar@pkrisc.cc.ukm.my*

A study on seasonality effect in an economy that does not impose tax on capital gains has its own merit because tax motivation is the most compelling and tested explanation for abnormal return behaviour in the month following the tax month, which is January in most countries. This hypothesis implies that January effect should be absent in economies that do not impose tax on capital gains. On the contrary, previous studies have found evidence of January effect in such systems (cf. Kato and Schallheim 1985; Jones *et al.*1987) while others found evidence against the January effect where the tax motivation applies (cf. Cox and Johnston 1998; Mehdian and Perry 2002). Motivated by the evidence of January effect regardless of tax-motivation, the compelling evidence of other explanations for January effect (cf. Abdul-Hadi 2005; Keim 1983; Lau *et al.* 2002) and the fact that evidence on seasonality effect in Malaysia is so far still mixed (Yong 1991; Abdul-Karim 2002; Pandey 2002; Yakob *et al.* 2005; Abdul-Rahim and Harjito 2006; Abdul-Rahim *et al.* 2006), the present study attempts to determine whether seasonality effect (if any) in the Malaysian equity market (i) persists in a longer study period, and (ii) can be explained from the perspective of the Fama-French model.

The choice of the Fama-French model in this study is timely because it has been accepted as the " workhorse for risk adjustment ..." (Hodrick and Zhang 2001). This is notwithstanding the fact that the model has been a subject of controversy particularly because of the lack of theoretical explanation underlying the roles of its additional risk factors. Despite the criticisms, the empirical multifactor model has been proven effective in various settings. More importantly, the model passed the acid test for an asset pricing model when it has been proven effective to explain several major anomalies (Fama and French 1996) including IPO under pricing (Chen and Pan 1998). On that regard, besides building more evidence in the literature on seasonality effect in Asian emerging stock markets, this study contributes by adding seasonality effect into the list of the applications of the Fama-French model. The study used a sample of 9 portfolios constructed from 220 to 500 stocks listed on the Main Board of Bursa Malaysia over the period of 21 years from January 1985 to December 2005, a period most probably among the longest ever tested on the Fama-French model based on the stock market in Malaysia (cf. Allen and Cleary 1998; Drew and Veeraraghavan 2002; 2003; Abdul-Hadi and Mohd. Nor 2006). The rest of the article is organised as follows. Section 2 presents the background studies on seasonality effect and the Fama-French model. Section 3 describes the data and methodology. Section 4 presents the findings and discussion on the results while, Section 5 concludes and discusses the implications.

2. Background Studies

In the United States where the issue on seasonality effect apparently originated, the exceptionally high but unexploitable returns have almost always been associated with January. Of the countless studies, one that is widely accepted as the earliest and most important study on this so-called January effect was done by Rozeff and Kinney (1976). Looking at the average monthly returns on stocks traded on the New York Stock Exchange (NYSE) over a 70-year period from 1904 to 1974, they found the average return in January to be higher than any other month except for the period of 1929 to 1940. Lakonishok and Smidt (1988), in a more comprehensive study which extended the period to 90 years (1897-1986) and using a different set of data, that is, Dow Jones Industrial Average (DJIA) data, found that returns are persistently anomalous in January. Similar evidence on January effect was also detected in a shorter study period (1963-1979) by Keim (1983) from a sample of securities traded in the NYSE. Haugen and Jorion (1996) found that the January effect remains elegant for the period of 1926 until 1993 with no significant sign of disappearance even after the reintroduction of the issue in 1976. In the other parts of the world, studies on January effect are relatively less rigorous but the market anomaly remains supported. The study by Gultekin and Gultekin (1983) could be the most comprehensive with respect to January effect as an international phenomenon. They found significantly unusual market activity in January in the US as well as several other European countries, Australia, Japan and Singapore. Pandey (2002), Yakob *et al.* (2005), Abdul-Rahim and Harjito (2006) and Abdul-Rahim *et al.* (2006) are among those who found evidence of January effect in ASEAN countries. With evidence in support of January effect sufficiently established, the interest of the more recent studies has shifted toward the explanations of the stock market anomaly.

Particularly in the US where the seasonality effect is associated with January, the most frequently cited and tested explanation is tax-loss selling hypothesis because the abnormal returns occur in the month following the tax-month. Proponents of tax-loss selling hypothesis like Dyl (1977), Givoly and Ovadia (1983), Reinganum (1983), Keim (1983), Badrinath and Lewellen (1991), Dyl and Maberly (1992), Eakins and Sewell (1993) and Fant and Peterson (1995) argue that at the end of the year, investors holding poor performing stocks take short positions to reduce the taxable capital gains. At the turn of the year as investors re-enter the market, stock prices rally creating upward price pressure and therefore, abnormal returns during the month. However, because the tax-loss selling hypothesis implies that January effect should not be the phenomenon in the absence of tax on capital gains, it can be challenged easily. As is the case in Malaysia where capital gains are not subject to tax, the evidence on January effect is mixed. Using 6-sector indexes of the Bursa Malaysia, Yong (1991) found evidence consistent with tax-loss selling hypothesis when January effect failed to be detected. However, using several sector indexes as well as portfolios of stocks in the same market in more recent years, Abdul-Karim (2002) found strong evidence in favour of January/February effect. The evidence in Malaysia is not an isolated case because January effect was also found in several emerging ASEAN markets despite the tax exemptions on capital gains (Gultekin and Gultekin 1983; Abdul-Rahim and Harjito 2006; Abdul-Rahim *et al.* 2006). In a much earlier study, Kato and Schallheim (1985) also found the January effect present in a sample of Japanese firms despite the absence of capital gains tax system in the country. Similarly, extending their search back to 1871, Jones *et al.* (1987) found that the January effect in the US market had already existed since the pre-tax period.

Arguments against tax-loss selling are exacerbated when January effect is absent in countries that impose capital gain taxes. In a sample of firms listed in NYSE and American Stock Exchange (AMEX) over the period of 1888 to 1992, Cox and Johnston (1998) found that stocks with high potential for tax loss selling do not exhibit abnormal return in January. Similarly, using market indexes (DJ Composite, NYSE Composite, and S&P 500), Mehdian and Perry (2002) also found that after the 1987 market crash, the January return is no longer significantly different from returns of other months. While the paradox surrounding tax-loss hypothesis is far from being solved, research efforts were shifted toward other explanations for the January effect.

When Keim (1983) found support for January effect, he also found a stable nega relation between abnormal return and firm size with the relationship being more pronoun in January. Similarly, Haugen and Jorion (1996) discovered that January effect magnifie the smallest firm categories. In fact, in almost all empirical studies of the tax-loss sell hypothesis, abnormal returns in January is mainly contributed by returns of small firms. common argument linking January effect and small firm is that because the flavour of month (that is, when investors re-enter the market) is small thinly traded stocks, the lar impact of price pressure on such stocks exaggerate abnormal return in January. Others l at the behaviour of fund managers during the turning point of the year to explain the mar anomaly. For the purpose of 'window dressing' or 'performance hedging' these manag re-balance their portfolios to comprise conservative, low risk stocks (normally of la companies) for performance evaluation at the end of the year. With respect to the Malays market, Chui and Wei (1998) found a significant relationship between stock returns and f size for the period of 1981 to 1993. Covering a study period from 1988 to 1996, Lau *et* (2002) also found evidence consistent with the size effect. Nonetheless, for a more rec period of 1990 to 2003, Abdul-Hadi (2005) found contradicting evidence on the size eff

Besides size which normally is measured by the firm market value of equity (M seasonality effect is also linked with other characteristics of the firm such as beta, book- market ratio (B/M), earning yield (E/P), dividend yield (D/P), sales growth and so forth. focus of the present study is the effect of size (ME) and book-to-market ratio (B/M) gi the growing attention these factors have received after the critical findings of Fama French (1992; 1993; 1996), who argued and proved that in addition to market risk fac stock returns are significant explained by two firm-specific factors, namely size (ME) and M. Davis (1994) and Kothari and Shanken (1997) among others also found reliable evide that B/M tracks cross-sectional and time-series variations in expected returns. The stu by Chui and Wei (1998) provide initial evidence on the potential role of B/M in explaini seasonality effect in the Malaysian stock market.

In this study, we evaluated the ability of ME and B/M simultaneously in the context a three-factor model introduced by the Fama-French model as potential explanations seasonality effect in Bursa Malaysia. The reason for selecting this model over countl others is because after the very much debated empirical failure of the standard one-fac capital asset pricing model (CAPM), this model gained so much attention that it is curren the workhorse for risk-adjustment in the academic circles (Hodrick and Zhang 2001). Desp the allegation, the results of 33 studies reported in Table 1 (excluding the studies in Pane by Fama and French including Davis *et al.* (2000) which naturally provide evidence support of the model) suggest that the performance of this model is rather mixed. instance, Panel B shows that of 19 other studies that were conducted on the same sam market, only 7 (36.8 per cent) lend support to this model. Similarly, the summary in Pane indicates that only 4 (40.0 per cent) of 10 studies conducted in other stock markets supp the Fama-French model. With respect to the testing of the Fama-French model on Malaysian equity market, as summarised in Panel D of Table 1, all of the four studies (All and Cleary 1998; Drew and Veeraraghavan 2002, 2003; Abdul-Hadi and Mohd. Nor 200 provide evidence in favour of the Fama-French model. Overall, these statistics indicate th the performance of the Fama-French model is much more convincing compared to that CAPM. The present study attempts to investigate to what extent this allegation holds

Table 1. Summary of empirical studies on the Fama-French model

No	Studies	Sample markets	Study period	Statistical tests	Support model?
Panel A: Empirical studies by Fama and French in the United States					
1	Fama and French (1993)	NYSE/AMEX/ Nasdaq	1963-1991	t(a), F-GRS	Yes
2	Fama and French (1995)	NYSE/AMEX/ Nasdaq	1963-1992	t(a), R^2	Yes
3	Fama and French (1996)	NYSE/AMEX/ Nasdaq	1963-1993	t(a), F-GRS	Yes
4	Davis *et al*. (2000)	NYSE/AMEX/ Nasdaq	1925-1996	t(a), F-GRS	Yes
Panel B: Empirical studies by other researchers in the United States					
1	Brennan and Subrahmanyam (1996)	NYSE	1984-1991	t(a),F-GRS	No
2	Jaganathan and Wang (1996)	NYSE/AMEX	1962-1990	t(a), HJ distance	No
3	Haugen and Baker (1996)	Russell Index	1979-1993	t(a), R^2	No
4	Daniel and Titman (1997)	NYSE/AMEX/ Nasdaq	1963-1993	t(a), F-GRS	No
5	Kim (1997)	NYSE/AMEX/ Nasdaq	1958-1993	t(a), R^2	Yes
6	Porras (1998)	NYSE/AMEX/ Nasdaq	1982-1995	t(a), F-GRS	No
7	Lewellan (1999)	NYSE/AMEX/ Nasdaq	1964-1994	t(a), F-GRS	Yes
8	Velu and Zhou (1999)	NYSE/AMEX/ Nasdaq	1964-1992	GMM	No
9	Brennan *et al*. (2001)	NYSE/AMEX/ Nasdaq	1950-1999	t(a), GARCH	Yes
10	Hodrick and Zhang (2001)	NYSE/AMEX/ Nasdaq	1952-1997	H-J Distance	No
11	Mun *et al*. (2001)	S&P500 Index	1986-1996	t(m)	Yes
12	Wu (2002)	NYSE/AMEX	1958-1995	t(a), GMM	Yes
13	Faff (2003)	Russell Index	1995-1999	GMM	Yes
14	Tai (2003)	NYSE/AMEX/ Nasdaq	1953-2000	MGARCH	Yes
15	Bali and Cakici (2004)	NYSE/AMEX/ Nasdaq	1958-2001	t(a), R^2	No
16	Chollete (2004)	NYSE/AMEX	1962-2001	F-GRS, GMM	No
17	Da and Gao (2004)	NYSE/AMEX/ Nasdaq	1962-2002	F-GRS, GMM	No
18	Liu (2004)	NYSE/AMEX/ Nasdaq	1960-2003	t(a), R^2	No
19	Barthold and Peare (2005)	NYSE	1970-1996	t(a), R^2	No

Continued

Table 1 *(Continued)*

Panel C. Empirical studies in other markets					
1	Griffin (2002)	US/UK/Japan/ Canada	1981-1995	t(a), R^2	Yes
2	Fletcher (2001)	UK	1982-1996	t(a), F-GRS	No
3	Miralles andMiralles (2005)	Spain	1994-2002	GMM	No
4	Daniel *et a*l. (2001)	Japan	1975-1997	t(a), F-GRS	No
5	Faff (2004)	Australia	1996-1999	GMM	No
6	Gaunt (2004)	Australia	1981-2000	t(a), R^2	Yes
7	Chan and Faff (2005)	Australia	1989-1998	GMM	No
8	Wang (2004)	China	1994-2000	t(a)	Yes
9	Cao *et al.* (2005)	China	1999-2002	t(m), ANN	No
10	Chen and Pan (1998)	Taiwan	1992-1994	CAR	Yes
Panel D. Empirical studies in Malaysia					
1	Allen and Cleary (1998)	Malaysia (MAS)	1977-1992	t(b), R^2	Yes
2	Drew and Veeraraghavan (2002)	Malaysia	1991-1999	t(a), R^2	Yes
3	Drew and Veeraraghavan (2003)	MAS/HK/ Korea/Phil	1991-1999	t(a), R^2	Yes
4	Abd-Hadi and Mohd. Nor (2006)	Malaysia	1991 - 2004	R^2 & Errors	Yes

Note: Abbreviations F-GRS = multivariate method of Gibbons *et al.* (1989); (M) GARCH = multivariate generalised autoregressive conditional heteroscedasticity; GMM = generalised method of moments; H-J Distance = error metric of Hansen and Jaganathan (1997); and *t*(.) = *t*-statistics for intercept (α), β, and mean (μ); CAR = cumulative abnormal returns; IPO = initial public offering; and ANN = artificial neural networks.

examining if the Fama-French model can explain a market anomaly that CAPM has failed to explain in earlier studies. By testing both models simultaneously, this study is expected to shed light on the question whether seasonality effect (if any) in this market is real or merely due to the model's inefficiency. Thus, this study contributes to the existing literature in several ways. First it provides additional evidence on the Fama-French model which is still limited in Malaysian stock market. Second, it is expected to produce evidence on seasonality effect using a more comprehensive sample. This study covers a much longer study period and uses a wider range of sample stocks than those in previous studies on the Malaysian equity market. More importantly, it extends the applications of the Fama-French model as explanation of CAPM anomalies (Fama and French 1996) to include seasonality effect.

3. Data and Methodology

3.1 The Data

The study employed data for 220 to 500 companies listed on the Main Board of Bursa Malaysia and covered a 21-year period from January 1985 to December 2005. Two sets of data were used: (i) monthly data on stock closing prices, interest rate on 3-month Treasury Bills and Exchange Main Board All Shares (EMAS) closing price index, and (ii) year-end data on market value of equity (ME) and book-to-market ratio (B/M). The data was sourced from *Thompson's Data Stream* and *Investors' Digest*. The large sample is necessary for this study to ensure that the resulting portfolios are well-diversified. Copeland and Weston (1988) and Womack and Zhang (2003) suggest that optimal diversification effect is achieved in portfolios that are composed of at least 30 stocks. As is the practice in previous studies on this issue (Fama and French 1993, 1996; Drew and Veeraraghavan 2002; 2003), the construction of the portfolios is repeated at the end of each year such that the number of stocks that falls under each category or portfolios varies from one year to another. Details of the portfolio composition are provided in Table A.1 in the Appendix.

3.2 The Dependent Variables

The dependent variables in this study are the monthly value weighted-average rate of returns on the test portfolio net of the risk-free rate of returns (R_i-R_F). To construct the test portfolios, at the end of December of year *t-1*, the sample stocks wereranked and sorted into: (i) three ME categories that is, 30 per cent smallest (S), 40 per cent medium (M) and 30 per cent biggest (B); and (ii) three B/M categories that is, 30 per cent highest B/M (H), 40 percent medium (M) and 30 percent lowest B/M (L). Then, following the procedure illustrated in Figure 1, we constructed 9 test portfolios double-sorted on ME and BM. The choice of 9 portfolios is based on the need to produce well-diversified portfolios (Copeland and Weston 1988; Womack and Zhang 2003) given the limited number of sample firms during the earlier study periods.

3.3 The Independent or Explanatory Variables

In general, the study used three variables to explain the portfolio returns, namely the market risk premium plus premiums on two additional risks related to size and distress. The market risk premium (R_M-R_F) is the return on the market portfolio net of the risk-free rate of return (R_F). This study has chosen EMAS over the KLSE Composite Index (KLCI) to proxy for market portfolio because the former is more representative of the sample population, that is, Main Board companies. Unlike KLCI which is based on 100 component stocks, EMAS is composed of all Main Board stocks and as such is more consistent with the market portfolio formed by Fama and French (1993; 1996) which includes all stocks listed on NYSE, NASDAQ and AMEX. Following Fama and French (1993), this study uses ME and B/M to proxy for

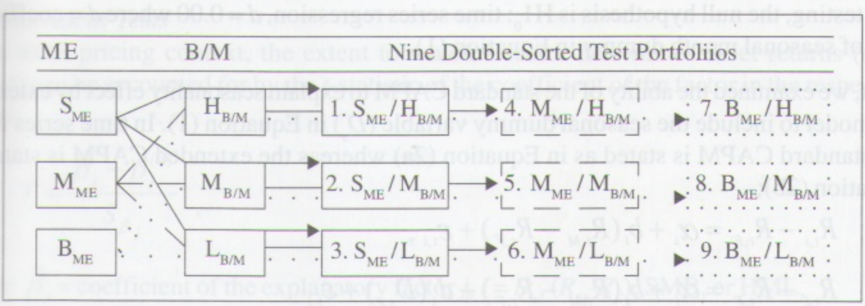

Figure 1. Procedure for constructing the ME/BM double-sorted test portfolios

Note: abbreviation S = small , M = medium, B = big, H = high, L = low, ME = market value of equity and B/M = book-to-market ratio. The resulting test portfolios are in numbered (1 to 9) boxes.

size and distress, respectively. The proxy for risk-free rate of return (R_F) is the monthly-adjusted-rate of return on the 3-month Treasury-Bills.

In the framework of the Fama-French model, premiums on the additional risks related to size and distress are denoted as SMB and HML, respectively. Using the same procedure illustrated in Figure 1 (except for ME which are only divided into S and B categories), we formed zero-investment portfolios to mimic risk related to size (SMB) and distress (HML). SMB is the difference between the simple average of returns on Small and Big ME portfolios (that is, [SH+SM+SL]/3 – [BH+BM+BL]/3). This procedure ensures that the premium on size risk is relatively free from the influence of distress risk because the Small and Big portfolios have about the same weighted-average B/M. Similarly, HML is the simple average of returns on High minus Low B/M portfolios (that is, [HS + HB]/2 – [LS+LB]/2).

3.4 Specification of the Models and the Hypotheses

The development of Fama-French model is based on the time series regression proposed by Black *et al.* (1972). Accordingly, similar to earlier studies (cf. Fama and French 1993, 1996a; Davis *et al.* 2000; Drew and Veraraghavan 2002; 2003; Bali and Cakici 2004) the factor loadings in this study were also estimated using time-series multiple regressions. To set the stage, we determined the presence of seasonality effect in the studied market using the equation previously used by Pietranico and Riepe (2004):

$$R_{t,i} - R_{t,F} = \alpha_i + d_i(D_{s,t}) + \varepsilon_{t,i} \tag{1}$$

where $R_{t,i}$ = realised monthly returns on the *i*th portfolio, i = 1, …, 9, at the end of month t
α_i = intercept term for the *i*th portfolio
d_i = loading on the seasonal dummy variable for the *i*th portfolio
R_F = return on the risk-free security
D_s = dummy variable that takes a value 1 if the month is the identified seasonal month or 0 otherwise
ε_i = disturbance term

In the form of a null hypothesis:

$H1_0$: There is no evidence of seasonality effect detected in the studied market. For statistical testing, the null hypothesis is $H1_0$: time series regression, $d = 0.00$ where d = coefficient of seasonal month dummy in Equation (1).

Next, we examined the ability of the standard CAPM to explain seasonality effect by extending the model to include the seasonal dummy variable (D_s) in Equation (1). In time series form, the standard CAPM is stated as in Equation (2a) whereas the extended CAPM is stated in Equation (2b);

$$R_{t,i} - R_{t,F} = \alpha_i + b_i(R_{t,M} - R_{t,F}) + \varepsilon_{t,i}\,, \tag{2a}$$

$$R_{t,i} - R_{t,F} = \alpha_i + b_i(R_{t,M} - R_{t,F}) + d_i(D_{s,t}) + \varepsilon_{t,i} \tag{2b}$$

where $R_{t,I}$, R_F, α_i, d_i, D_s, and ε_i = as in Equation (1)
b_i = loading on the market risk premium
R_M = return on the market portfolio

CAPM successfully explains the seasonality effect if loading on D_S becomes insignificant at the same time that loading on market risk premium (R_M-R_F) is significant. In other words, market risk premium captures the variations in returns due to seasonality effect. In the form of null hypothesis:

$H2_a$: There is no evidence of seasonality effect when CAPM is used to explain realised excess returns on portfolios. For statistical testing the null hypothesis is $H2_0$: time series regression, $d = 0.00$ while simultaneously $b = 0$ where d = coefficient of the seasonal dummy variable (D_S) and b = coefficient of market risk premium in Equation (2b).

Finally we have the specifications of the Fama-French model which, in the time-series regression form, is as represented in Equation (3a);

$$R_{t,i} - R_{t,F} = \alpha_i + b_i(R_{t,M} - R_{t,F}) + s_i(SML)_t + h_i(HML)_t + \varepsilon_{t,i} \qquad (3a)$$

$$R_{t,i} - R_{t,F} = \alpha_i + b_i(R_{t,M} - R_{t,F}) + s_i(SML)_t + h_i(HML)_t + d_{s,i}(D_{s,t}) + \varepsilon_{t,i} \qquad (3b)$$

where $R_{t,i}$, R_F, α_i, d_i, D_s, and ε_i = as in Equation (1)
$R_{M,t}$, and b = as in Equation (2)
s_i and h_i = estimated loadings on SMB and HML, respectively for the ith portfolio
SMB = premium on risk related to size
HML = premium on risk related to distress

While Equation (2b) tries to explain seasonality effect using CAPM, Equation (3b) uses the Fama-French model instead. Specifically, if the coefficient of the seasonal month dummy (D_s) becomes insignificant at the same time that coefficients of R_M-R_F, SML and HML are significant, then the Fama-French model is said to effectively explain the seasonality effect. In the form of null hypothesis:

$H3_a$: There is no evidence of seasonality effect when Fama-French model is used to explain realised excess returns on portfolios. For statistical testing, the null hypothesis is $H3_0$: time series regression, $d_s = 0.00$ while simultaneously $\gamma = 0$ where d = coefficient of the seasonal dummy variable (D_S) and γ = coefficients other than d_s in Equation (3b).

3.5 *Statistical Tests*

In an asset pricing context, the extent to which a factor determines asset returns (or is priced) can be accounted for by the t-statistic of the coefficient of the factor in the respective model:

$$T = \frac{\hat{\beta}_j - \beta_j^{(0)}}{S_{\hat{\beta}_j}}, \qquad (4)$$

where $\hat{\beta}_j$ = coefficient of the explanatory factor j, $i = D_s$, (R_M-R_F), SMB, or HML,

$$S_{\hat{\beta}_j} = \frac{S_{y|x}}{S_x\sqrt{n-1}} = \text{standard error of } \hat{\beta}_j,$$

$S_{y/x}$ = estimated standard deviation in data y and $\hat{y}$ with variable x, and
n = number of observations in the time series.

In brief, an explanatory factor in a particular model is significantly in existence if the null hypothesis (H_0: $l_j^p = 0.0$) is rejected, that is, $|T| \geq t_{N-2,1-\alpha/2}$.

4. Results and Discussion

4.1 Descriptive Statistics

Fama and French's (1992; 1993; 1996) argument that returns are associated with risks related to size and distress implies that when the two characteristics are combined to form a portfolio, the risk-return trade-off rule would require that the portfolio will generate returns high enough to compensate for the high risks. With respect to our portfolios, this implies that Portfolio SH (which proxies for relatively small sized and distressed companies) produces the highest while Portfolio BL (which proxies for relatively big size and healthy companies) produces the lowest standard deviation and monthly returns. The results in Table 2 are obviously quite consistent with this prediction. First, the standard deviation reported for Portfolio SH (15.6 per cent) is the second highest while the Portfolio BL (8.2 per cent) is the lowest relative to those of the other portfolios. Second, Portfolio SH also reports the second highest (3.6 per cent) while Portfolio BL reports the lowest (1.4 per cent) average monthly returns. Despite a slight deviation from our prediction, the risk-return trade-off theory remains supported given the fact that Portfolio SL, which shows the highest measure of risk (15.8 per cent) is also the one that reports the highest average monthly returns (3.8 per cent). Note also that all four statistics reported in Table 2 show values that are almost consistently declining monotonically from Portfolio SH to BL. These findings uphold the traditional risk-return trade-off and to this extent Fama-French's arguments in relating size-distress with risk seem justified. The significantly high correlations (0.634 – 0.942) among size-distress portfolio returns could however, lead toward different results regarding the effect of size and distress on seasonality. In other words, seasonality effect, if it were to be traced in this market, might not be easily explained from the ME/BM framework given the high degree of interdependence among the returns on the test portfolios.

Next, reported in Table 3 are the characteristics of the time series variables, both the dependent and explanatory. They suggest that in general, the time-series regression models are appropriate to test the hypotheses in this study. First, regarding the normality distribution of the data, the Jarque-Bera (JB) statistics suggest that normally distributed data have a skewness value of zero ($S = 0.00$) and kurtosis of ($K = 3.00$) such that the JB statistic for normally distributed data should be zero. The null hypothesis (H_0: JB = 0.00) is rejected if JB $> \chi^2_{d.f.=2}$. As reported in Table 3, the JB statistics indicate that the normality distribution assumption is violated (JB = 13.83 to 6434.71). Nonetheless, such violation is normal for stock return series. Fortunately, the assumption that is of greater concern in time series analysis is the stationarity of the series, which in this study is determined by computing the Augmented Dickey-Fuller (ADF). The test specifies that the null hypothesis (H_0: $\gamma = 0$) that the series have a unit root is rejected if the ADF statistic is greater than the MacKinnon

Table 2. Descriptive statistics of and correlations coefficients

	Min	Max	Mean	StDev	SH	SM	SL	MH	MM	ML	BH	BM	BL
SH	-0.323	0.804	0.036	0.156	1								
SM	-0.303	0.882	0.032	0.142	.824	1							
SL	-0.335	0.904	0.038	0.158	.754	.733	1						
MH	-0.i321	0.759	0.025	0.128	.885	.877	.819	1					
MM	-0.245	0.635	0.020	0.114	.870	.855	.820	.942	1				
ML	-0.296	0.641	0.019	0.119	.833	.862	.789	.899	.911	1			
BH	-0.329	1.101	0.022	0.141	.814	.817	.714	.908	.886	.879	1		
BM	-0.344	0.545	0.019	0.096	.809	.742	.696	.867	.893	.857	.868	1	
BL	-0.280	0.322	0.014	0.082	.696	.667	.634	.753	.779	.779	.740	.865	1

Note: In all cases, N = 218 monthly observations. All correlations are significant at 1 per cent ($\alpha \leq 0.01$) level.

Table 3. Moments of the time series data and correlations among explanatory variables

	Skewness	Kurtosis	J-Bera	ADF_{12}	ADF_{6}	ADF_{1}	R_M-R_F	SMB	HML
Panel A: Test portfolios									
SH	1.796	9.495	523.390	-4.0857	-5.7008	-9.8683			
SM	1.830	10.138	611.297	-4.4682	-5.1006	-8.8786			
SL	1.705	8.371	384.547	-4.4314	-4.8518	-9.1125			
MH	1.419	8.287	342.081	-3.8979	-5.4154	-9.3465			
MM	1.523	8.886	417.241	-4.0071	-5.3349	-8.9682			
ML	1.451	8.668	385.157	-4.3441	-5.0705	-8.3604			
BH	2.650	19.181	2754.125	-4.2944	-5.9913	-9.8847			
BM	1.132	9.874	497.496	-3.9823	-5.9722	-9.0249			
BL	0.166	4.868	34.195	-3.8230	-6.1769	-9.2571			
Panel B: Explanatory variables									
R_M-R_F	0.163	5.760	73.393	-3.8595	-5.5303	-8.6580	1		
SMB	1.871	10.367	648.548	-4.2209	-4.7905	-9.6362	0.344	1	
HML	1.831	18.549	2424.224	-4.1446	-6.2552	-13.202	0.358	0.240	1

Note: In all cases, N = 218 monthly observations. All correlations are significant at 1 per cent ($\alpha \leq 0.01$) level. The unit root tests are done on three lags. The McKinnon critical values for lag 12 are -3.4621, -2.8750 and -2.5739, for lag 6 they are -3.4613, -2.8747, and -2.5737 and for lag 1 they are -3.4607, -2.8744 and -2.5736 at 1 per cent, 5 per cent and 10 per cent significant levels respectively.

critical value. As shown in Table 3, the unit root hypothesis is consistently rejected (p-value < 1%) at lags 1, 6 and 12. The ADF values for all series are always greater than the McKinnon critical value, indicating that the time-series data is suitable for time-series regression.

Another important characteristic, especially in the context of an asset pricing model, is the extent of independence between explanatory variables which ensures that each variable explains different common risk in stocks (Fama and French 1993; 1996). The extreme right columns in Panel B of Table 3 address this concern. The correlation between SMB and the market premium is only 0.344 whereas that between HML and market premium is only 0.358. These levels of independence allow us to conclude that SMB and HML proxy risk factors in stocks that are unique from market risk (R_M-R_F). Also note that even though in Table 2 the correlations among returns on the portfolios appear to be very high, the resulting correlation between SMB and HML is only 0.240. The correlations between Fama-French factors in this study are consistent with those found in Fama and French (1996).

4.2 *Preliminary Findings on Seasonality Effect*

Before we statistically test for the existence of seasonality effect in this market, we shall first identify which among the twelve months (Jan – Dec) is most appropriate to be designated as the seasonal month (s) in the case of the Malaysian stock market. As shown in Figure 2, with regard to Malaysian stock market, the seasonality effect seems to be most prevalent in February since except for Portfolio BL, the value-weighted return of February is always highest. This evidence is consistent with earlier finding (Abdul-Karim 2002; Pandey 2002; Abdul-Rahim and Harjito 2006; Abdul-Rahim *et al.* 2006) who found that in most cases the highest average monthly returns takes place in February. Portfolio BL, the portfolio that is supposed to represent the biggest and most healthy companies, reports the highest returns in December. In fact, December consistently reports the second highest returns for all Big portfolios as well as portfolio ML. Figure 2 also reveals a particular pattern that might help to explain why earlier studies have been contemplating between January and February when returns are highest in this market. Note that January returns are consistently second highest in the "Small" portfolio category such that the January effect is an attribute of small companies. Details of average monthly returns are provided in Table A.2 in the Appendix.

Regardless of the month when seasonality actually takes place, to a great extent, this study agrees with the widely-accepted explanation that seasonality effect in Malaysian equity market is due to the trading activities before and after Chinese New Year (CNY) celebration. In other words, January/February effects are the results of the behaviour of Chinese investors, the dominant traders in Bursa Malaysia, around these months. This argument is compelling because for the last 25 years from 1980 to 2005, CNY has been celebrated either in January or February. Thus, for the purposes of the present study, the seasonality effect is comfortably associated with February and with that we proceed with the regression tests by establishing the month of February to take the value of 1 in the dummy variable (D_S or now specifically D_{FEB}) in Equations (1), (2b) and (3b).

Another pattern that is also worth noting is that, as shown in Figure 2, all portfolios except portfolio SM consistently report August as the month with the lowest (and negative) average monthly returns. Even then, portfolio SM reports the second lowest return in August. This feature is not unique to this study or even to the Malaysian stock market because the "quiet month" of August is also common in the ASEAN region (Abdul-Rahim *et al.* 2006) as well as the US, the UK, and Japan (Abdul-Karim 2002). Overall, the preliminary results on seasonality patterns that we gather so far are similar with those in Abdul-Karim (2002), Pandey (2002), Abdul-Rahim and Harjito (2006) and Abdul-Rahim *et al.* (2006).

On identifying the seasonality in this market, we ran the time series regression as specified in Equation (1) to test the first null hypothesis of no significant evidence of seasonality. Specifically, in Equation (1) $H1_0$: *$d = 0.00$ where d = coefficient of seasonal month dummy*. The results, as reported in Table 4, show that the coefficient of seasonal

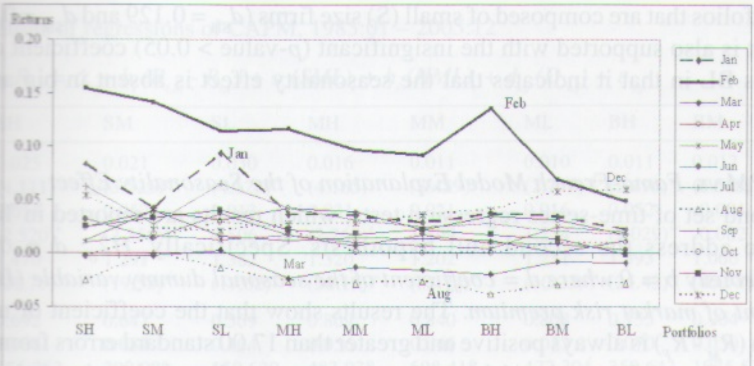

Figure 2. Pattern of the average monthly returns on the test portfolios, 1985:01 – 2005:12

Notes: Portfolios in this figure are formed double-sorted on ME (proxies size) and B/M (proxies distress). The first letter of the portfolio label always represents the ME categories (i.e., S = small, M = medium, and B = big ME or size companies) whereas the second letter always represents the B/M ratio (i.e., H = high, M = medium, and L = low B/M or distress companies).

Table 4. Regression results on seasonality effect, 1985:01 – 2005:12

$$R_{t,i} - R_{t,F} = \alpha_i + d_i (D_{s,t}) + \varepsilon_{t,i} \quad \text{.............(1)}$$

Variable	SH	SM	SL	MH	MM	ML	BH	BM	BL
Constant	0.014	0.011	0.019	0.006	0.002	0.001	0.000	0.003	0.000
	(1.286)	(1.134)	(1.776)	(0.711)	(0.228)	(0.079)	(0.013)	(0.521)	(-0.040)
D_{FEB}	0.129	0.119	0.085	0.099	0.083	0.080	0.124	0.055	0.038
	(3.548)**	(3.599)**	(2.248)*	(3.304)**	(3.108)**	(2.845)**	(3.755)**	(2.391)*	(1.938)
Adj-R^2	0.049	0.050	0.018	0.042	0.037	0.030	0.055	0.020	0.012
Std Error	0.152	0.138	0.157	0.125	0.112	0.117	0.138	0.095	0.082
F-stats	12.587	12.952	5.054	10.916	9.662	8.093	14.100	5.715	3.756
D-W stat	1.695	1.607	1.802	1.817	1.832	1.923	1.862	1.876	1.949

Notes: Each cell contains the coefficient (and the *t*-statistics) of the respective explanatory variable. The asterisks ** and * indicate significance at 1per cent and 5 per cent, respectively. The Durbin-Watson statistics (D-W ~ 2.00) suggest that the autocorrelations in error terms are immaterial.

month dummy (D_{FEB}) is always greater than two standard errors from zero except for portfolio BL ($t(d_{BL})$=1.938) and therefore, the first null hypothesis is rejected. This evidence is strong and consistent with the findings of earlier studies (Abdul-Rahim and Harjito 2006; Abdul-Rahim *et al.* 2006) in relation to the existence of seasonality effect in Malaysian stock market. With respect to the magnitude of the effect, the fact that the strongest effect is consistently in portfolios that exhibit high (H) B/M ratio (d_{SH} = 0.129 and d_{BH} = 0.124) suggest that the abnormal returns in February is an attribute of firms that are relatively in distress. Similarly, seasonality effect detected in this market could be explained by the behaviour of relatively smaller firms in February because a greater coefficient is associated with portfolios that are composed of small (S) size firms ($d_{SH} = 0.129$ and $d_{SM} = 0.119$). This argument is also supported with the insignificant (p-value > 0.05) coefficient of D_{FEB} for portfolios BL in that it indicates that the seasonality effect is absent in big and healthy firms.

4.3 CAPM vs. Fama-French Model Explanation of the Seasonality Effect

The second set of time-series regression tests, which results are reported in Table 5, are meant to address the second null hypothesis. Specifically, *$H2_0$: $d = 0.00$ while simultaneously $b = 0$ where d = coefficient of the seasonal dummy variable (D_S) and b = coefficient of market risk premium.* The results show that the coefficient of market risk premium (R_M - R_F) is always positive and greater than 17.00 standard errors from zero. This evidence indicates that market risk captures most but not all variations in excess returns. It still leaves some abnormal returns in February remaining unexplained when the coefficients of D_{FEB} remain significant (p-value < 5 per cent) in four of the eight test portfolios that previously (Table 4) exhibited seasonal pattern. This evidence is not sufficient to conclude that CAPM explains seasonality effect in this market. Regarding the null hypothesis, $H2_0$ can be rejected but only marginally.

It is also worth noting that the overall performance of CAPM still lends strong support to the current rather established conclusion that as an asset pricing model, CAPM empirically fails to capture all variations in returns. The resulting adjusted-R^2 of CAPM (76.45 per cent) is statistically significant to indicate a good model fit. Nonetheless, in an asset pricing context, this level of goodness-of-fit is inadequate as it suggests that CAPM leaves 23.55 per cent of variations in returns on portfolios remaining unexplained. The empirical fault of an asset pricing model can also be detected from the magnitude of its intercept (Fama and French, 1996). The intercept is a measure of Jensen's alpha when significant proximate abnormal returns cannot be explained by the existing explanatory variables in the model. In other words, a significant intercept suggests that there are omitted variables that need to be incorporated into the model in order to absorb the remaining variations in returns perfectly. As reported in Table 5, the intercept of CAPM is somewhat monotonously declining from the most risky portfolio ($\alpha_{SH} = 2.5$ per cent per month) to the least risky portfolio ($\alpha_{BL} = 0.6$ per cent per month), but it is consistently significantly different from zero (p-value < 0.05). Both the adjusted R^2 and significant intercepts support current perception (Fama and French 1992; 1993; 1996) that CAPM empirically fails as an asset pricing model. Based on similar finding, Fama and French (1993) suggest that additional risk factors that are related to size (SMB) and distress (HML) be incorporated into the asset pricing model to alleviate the weakness in CAPM.

The last regression analyses were run to test the third hypothesis which suggests that seasonality effect can be explained from the Fama-French perspective, that is, *$H3_0$: d_s = 0.00 while simultaneously $\gamma = 0$ where d = coefficient of the seasonal dummy variable (D_S) and γ = coefficients other than d_s*. The results of the Fama-French model are reported in Table 6. First, in relation to the role of the three explanatory variables in the Fama-French model. The coefficients of SMB are not always positive but consistently greater than 2.00 standard errors from zero. The role of HML could have been larger (t(h) > 3.00) if not for one case where it is insignificant ($t(h_{ML}) = 0.699$, p-value > 0.05). In addition, the incorporation of SMB and HML does not seem to absorb the role of market risk in CAPM. Market risk

Table 5: Results of regressions of CAPM, 1985:01 – 2005:12

$$R_{t,j} - R_{t,F} = \alpha_i + b_i (R_{t,M} - R_{t,F}) + s_i (SML)_t + h_i (HML)_t + d_{s,j} (D_{s,j}) + \varepsilon_{t,j} \quad \text{........ (2b)}$$

Variable	SH	SM	SL	MH	MM	ML	BH	BM	BL
Constant	0.025	0.021	0.030	0.016	0.011	0.010	0.011	0.012	0.006
	(4.111)**	(3.525)**	(4.123)**	(4.165)**	(3.439)**	(2.757)**	(2.196)*	(5.568)**	(3.008)*
D_{feb}	0.053	0.052	0.013	0.031	0.021	0.016	0.052	-0.001	-0.008
	(2.520)*	(2.556)*	0.506	(2.269)*	1.899	1.254	(3.029)**	-0.082	-1.070
R_m-R_f	1.470	1.291	1.382	1.320	1.202	1.231	1.393	1.066	0.886
	(21.76)**	(19.58)**	(17.02)**	(30.19)**	(33.74)**	(30.23)**	(25.43)**	(45.74)**	(36.45)**
Adj-R^2	0.692	0.647	0.569	0.809	0.840	0.808	0.755	0.904	0.856
S. Error	0.086	0.084	0.104	0.056	0.046	0.052	0.070	0.030	0.031
F-Stat	256.263	209.080	150.620	483.038	598.418	477.304	350.643	1075.57	677.137
D-W Stat	1.592	1.749	1.951	1.918	1.928	1.993	1.861	1.742	1.929

Notes: Each cell contains the coefficient (and the t-statistics) of the respective explanatory variable. The asterisks ** and * indicate significance at 1 per cent and 5 per cent, respectively. The Durbin-Watson statistics (D-W ~ 2.00) suggest that the autocorrelations in error terms are immaterial.

Table 6. Results of regressions of Fama-French Model, 1985:01 –2005:12

$$R_{t,j} - R_{t,F} = \alpha_i + b_i (R_{t,M} - R_{t,F}) + s_i (SML)_t + h_i (HML)_t + d_{s,j} (D_{s,j}) + \varepsilon_{t,j} \quad \text{........ (3b)}$$

Variable	SH	SM	SL	MH	MM	ML	BH	BM	BL
Constant	0.015	0.009	0.019	0.010	0.006	0.005	0.007	0.013	0.009
	(3.840)**	(2.698)**	(3.562)**	(4.542)**	(2.608)**	(1.596)	(1.889)	(6.600)**	(5.495)**
D_{feb}	0.007	0.011	-0.012	-0.002	-0.001	-0.001	0.023	-0.002	0.006
	(0.483)	(0.868)	(-0.637)	(-0.258)	(-0.089)	(-0.114)	(1.712)	(-0.310)	(1.108)
R_m-R_f	1.109	0.957	1.173	1.065	1.031	1.093	1.177	1.058	0.998
	(22.98)**	(22.50)**	(18.08)**	(39.50)**	(36.47)**	(31.15)**	(25.12)**	(45.44)**	(49.42)**
SMB	0.844	1.144	1.284	0.514	0.451	0.582	0.173	-0.160	-0.200
	(12.56)**	(19.32)**	(14.22)**	(13.70)**	(11.46)**	(11.92)**	(2.65)**	(-4.93)**	(-7.11)**
HML	0.763	0.345	-0.348	0.619	0.310	0.035	0.786	0.193	-0.298
	(10.95)**	(5.63)**	(-3.72)**	(15.90)**	(7.61)**	(0.699)	(11.63)**	(5.749)**	(-10.2)**
Adj-R^2	0.871	0.879	0.774	0.940	0.917	0.883	0.853	0.922	0.919
S. Error	0.056	0.049	0.075	0.031	0.033	0.041	0.054	0.027	0.023
F-Stat	384.027	415.149	195.360	897.875	631.779	428.568	329.555	668.963	641.466
D-W Stat	1.432	1.958	2.137	1.948	2.085	1.963	1.737	1.833	1.747

Notes: Each cell contains the coefficient (and the *t*-statistics) of the respective explanatory variable. The asterisks ** and * indicate significance at 5 per cent and 10 per cent, respectively. The Durbin-Watson statistics (D-W ~ 2.00) suggest that the autocorrelations in error terms are immaterial.

continues to produce coefficients that are always positive and in fact greater than 18.00 standard errors from zero. This finding supports the Fama and French (1996) claim that SMB and HML explain variations in stock returns separate from those that are explained by market risk.

Second, with strong evidence that all three risk factors in the Fama-French model play a significant role in explaining stock returns, the focus is on their effectiveness in absorbing the variations due to February returns. Table 6 shows that none of the coefficients of D_{FEB} remain significant when returns are explained by this model. In other words, because the significance of the D_{FEB} disappears simultaneously with consistently significant coefficients of R_M-R_F, SMB and HML, $H3_0$ can be totally rejected. Furthermore, given the significance of D_{FEB} as a result of adding SMB and HML into the model, it may be surmised that these factors completely absorb the explanatory power of D_{FEB}. This finding is of critical importance to the existing evidence on seasonality effect because it intuitively suggests that seasonality effect is not significant in this market. The evidence indicates that evidence on seasonality effect found in earlier studies in this market could be due to the inefficiency of the model being employed. Evidently, the seemingly seasonal effect disappears once a more efficient model is used to explain variations in returns. Note also that compared to CAPM, the Fama-French model produces an average adjusted-R^2 (88.42 per cent) which is 15.66 per cent higher. The evidence from this study lends strong support to the Fama and French's (1996) proposition that the Fama-French model is an equilibrium model because it explains CAPM anomalies which in this study are in the form of February effect.

5. Conclusion and Implications

This study investigated whether seasonality effect (if any) in the Malaysian equity market can be explained by the Fama-French model as opposed to the standard CAPM. The tests were conducted using multiple time-series regression analyses on monthly returns data of 9 double-sorted ME/BM test portfolios constructed from 220 to 500 stocks listed on Bursa Malaysia over the period of 21 years from 1985:01 to 2005:12. The preliminary results which indicate persistent February effect in the Malaysian equity market explain the attention from researchers as well as participants of stock markets on the seasonality effect. However, the regression analyses of the Fama-French model produced contradicting evidence. The multifactor model eliminates entirely the significant evidence on an anomaly which apparently has been wrongly believed to warrant exceptional attention from investors. To be more specific, February effect which continues to remain persistent when returns are explained by the CAPM disappears when the Fama-French model is used.

The results from the Fama-French model so far indicate that the seemingly abnormal returns in February are merely an attribute of rational behaviour among investors during the month of February when investors are trading their stocks more actively than they normally do in other months. The returns are higher as a result of rational investors who demand premiums to compensate greater risks from their investment in companies that are relatively smaller and in distress. In other words, returns are in general higher in February because the additional trading activities during this month create greater price pressure on stocks of such companies than they do on other bigger and healthier companies. Future studies should look into the trading activities in February as opposed to the rest of the months to confirm this argument. For the time being, the results of this study can be taken as support to the proponent of stock market efficiency in Malaysia.

The effectiveness of the Fama-French model in explaining seasonality effect is attributed to the role of additional risks related to size (SMB) and distress (HML) in absorbing those variations of returns that are left unexplained by the market risk premium (R_M - R_F) of CAPM.

To a great extent, our approach of looking at seasonality effect from the perspective of the Fama-French model works to explain why the February effect remains so persistent throughout these years. Despite the widespread belief that the abnormally high returns should be a subject of exploitation by investors, the present study suggests that existence of February effect in the Malaysian stock market, if any, is mainly contributed by returns of small and distressed firms. In other words, the so-called February anomaly appears to be unexploited merely because it only reflects rational reactions of investors to the higher risks imposed by firms that are relatively smaller and in distress. The finding of this study also implies that the seemingly abnormal returns in February are another evidence of the empirical weakness of the CAPM. However, because this study is limited to common shares that are listed on the Main Board of Bursa Malaysia, the results may not be generalisable to explain the phenomenon of the market in general. Future research should address this limitation by including portfolios that have components of all stocks listed on the Bursa Malaysia.

Finally, even though the results of this study tend to suggest the superiority of the Fama-French model relative to CAPM, they are not sufficient to suggest that the former is a perfect alternative to the latter. This is because a perfect model will generate an intercept that is not significantly different from zero ($\alpha = 0.00$) beside an adjusted-R^2 of 1.00 (Fama and French 1993). Compared to that of CAPM, the range of remaining abnormal returns (α) of the Fama-French model reported in Table 6 reduces to between 1.5 per cent for the riskiest portfolio SH to 0.9 per cent for the least risky portfolio BL. However, this reduction is not small enough because in seven out of nine portfolios, the intercepts remain significant. There are several explanations for this result, one of which may be attributed to the number of stocks that form the test portfolios. As reported in Table A.1 in the Appendix, prior to 1994 more than half of the test portfolios were composed of less than 30-minimum required stocks of a well-diversified portfolio (Copeland and Weston 1988; Womack and Zhang 2003). The implication is that a good part of the unsystematic variations in the portfolio returns remain undiversified. This problem can be addressed in future studies by choosing sample markets that offer a sufficient number of sample firms. Another possible explanation for the significant intercepts is that there are variables that have been omitted that need to accounted for in the multifactor model. While the question on what the omitted factors are needs to be addressed in future studies, the findings of this study add support to the Fama and French's (1996) conclusion that like any other asset pricing model, the Fama-French model still has important gaps that require consistent efforts to fill up.

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Appendix

Table A.1. Composition of portfolios double-sorted on size and book-to-market ratatio, 1985 - 2005

Year	SH	SM	SL	MH	MM	ML	BH	BM	BL	Total
1985	23	30	15	31	33	26	13	27	27	220
1986	19	18	29	32	22	34	15	26	25	225
1987	22	32	15	36	33	23	11	27	31	230
1988	25	26	21	36	36	21	9	32	29	235
1989	26	24	22	40	35	21	6	37	29	240
1990	23	29	22	35	38	25	15	31	27	245
1991	19	33	23	43	36	21	13	31	31	250
1992	22	33	23	42	42	20	14	29	35	260
1993	22	22	39	42	47	21	18	24	40	275
1994	19	37	30	38	48	28	28	29	28	285
1995	18	33	41	46	46	30	27	35	29	305
1996	27	40	29	45	54	29	24	34	38	320
1997	29	38	34	51	54	29	20	42	38	335
1998	63	56	26	32	46	22	10	38	57	350
1999	35	44	30	56	59	32	18	43	48	365
2000	37	47	29	59	59	35	18	46	50	380
2001	44	66	34	72	68	52	28	58	58	480
2002	49	53	42	70	76	46	25	63	56	480
2003	60	51	33	66	83	44	18	58	67	480
2004	67	50	27	62	82	48	25	60	69	490
2005	60	54	36	45	80	75	34	39	77	60
Average	34	39	29	47	51	32	19	38	42	331

Note: The first letter in the column heading indicates the size category (i.e. S = small, M = medium, and B = big) whereas the second letter indicates the distress category (i.e. H = high, M = medium, and L = low).

Table A.2. Average monthly returns by month, January 1985 to December 2005

Month	SH	SM	SL	MH	MM	ML	BH	BM	BL
January	0.0832	0.0426	0.0935	0.0363	0.0358	0.0288	0.0378	0.0344	0.0222
February	0.1539	0.1414	0.1149	0.1166	0.0963	0.0915	0.1351	0.0693	0.0488
March	-0.0023	-0.0020	0.0032	-0.0232	-0.0128	-0.0153	-0.0198	-0.0081	-0.0092
April	0.0245	0.0387	0.0305	0.0422	0.0384	0.0223	0.0348	0.0307	0.0168
May	0.0104	0.0153	0.0132	0.0171	0.0108	0.0067	0.0041	0.0249	0.0189
June	0.0319	0.0135	0.0376	0.0173	0.0067	0.0175	0.0096	0.0086	0.0029
July	0.0623	0.0271	0.0654	0.0424	0.0308	0.0251	0.0281	0.0202	0.0183
August	-0.0255	0.0008	-0.0135	-0.0263	-0.0279	-0.0259	-0.0389	-0.0301	-0.0280
September	0.0117	0.0391	0.0155	0.0120	0.0142	0.0136	0.0119	0.0008	0.0119
October	-0.0021	0.0133	0.0285	0.0221	0.0014	0.0006	0.0001	0.0096	0.0102
November	0.0248	0.0330	0.0405	0.0220	0.0188	0.0220	0.0180	0.0134	-0.0013
December	0.0541	0.0222	0.0217	0.0269	0.0275	0.0358	0.0400	0.0577	0.0591
Minimum	-0.0255	-0.0020	-0.0135	-0.0263	-0.0279	-0.0259	-0.0389	-0.0301	-0.0280
Month	Aug	Mar	Aug	Aug	Aug	Aug	Aug	Aug	Aug
Maximum	0.1539	0.1414	0.1149	0.1166	0.0963	0.0915	0.1351	0.0693	0.0591
Month	Feb	Feb	Feb	Feb	Feb	Feb	Feb	Feb	Dec
Average	0.0356	0.0321	0.0376	0.0255	0.0200	0.0186	0.0217	0.0193	0.0142

Note: The first letter in the column heading indicates the size category (i.e. S = small, M = medium, and B = big) whereas the second letter indicates the distress category (i.e. H = high, M = medium, and L = low).