# TIME-VARYING RISK PREMIA FOR SIZE EFFECTS ON EQUITY REITS

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**ABSTRACT**

*We examine if the risk premia of the size effect on equity REITs (EREITs) are time-varying by using GARCH models. We also investigate how macroeconomic factors affect the size premia. We reexamine the size effect by using Fama-French three-factor model to demonstrate that the size effect exists in EREITs market. We investigate time-varying volatility for size effect by using a sample of publicly traded EREITs with GARCH family models. We find variation of the size premia partially results from volatility the bond market term spread and the volatility of short-term interest rates. The unexpected shock from fluctuation of the bond market term spread lowers the volatility of the size premia on EREITs return. We also find the big-sized EREITs are a good investment when default risk premium fluctuates dramatically.*

**JEL**: E52; G12; G32

**KEYWORDS:** equity REITs, Size Effect, GARCH, VAR, Volatility, Leverage Effect

## INTRODUCTION

The market capitalization of U.S. market Real Estate Investment Trusts (REITs) has grown rapidly since 1992. According to the National Association of Real Estate Investment Trusts (NAREIT), the market capitalization of the REIT industry grew from \$15.9 billion in 1992 to \$ 438 billion in 2006. Much of this growth is driven by investors who became disappointed with the downturn in performance of private property investments in the early 1990s. They turned to public real estate assets (such as REITs) for a more liquid way to get exposure to real estate markets.

Existing research suggests that portfolios of REITs outperform portfolios of common stocks. There might be characteristics of REITs that lead to the excess returns. One potential characteristic may be size. The size effect, which argues that average returns on small size firms exceed those on large size firms was first documented by Banz (1981). Fama and French (1992, 1993) show the size effect remains significant even after controlling for beta. In REITs literature, only a few papers refer to the size effect. Colwell and Park (1990) find evidence that there are size-related seasonality and a January effects in the REIT market. McIntosh, Liang, and Tompkins (1991) show that there is a “positive” small-firm effect for REITs over the period of 1974 to 1988.

In this paper, we examine size effects on Equity REITs (EREITs) returns by using the Fama-French three-factor model for the period of July 1995 to December 2006. Our result indicates that there is a “negative” size effect (reverse size effect) during our research period. This opposite effect leads us to believe the size premia may change over time. To identify time series fluctuations we investigate the size effect on EREIT returns using a family of GARCH models.

According to the capital asset pricing model (CAPM), there is a positive relationship between expected security returns and market betas, and the market betas are sufficient to describe equity expected returns. The CAPM states that investors price only market risk, beta. However, more recent literature has identified several non-market risk factors that should be considered in pricing assets. A long literature has shown that various firm characteristics, for example, size, price-to-earnings ratios, book-to-market equity ratios, and leverage degrees also have explanatory power in cross sectional expected equity returns (Banz, 1981; and Bhandari, 1988).

The size effect is one of the most distinguished anomalies in finance. Banz (1981) finds that market equity (ME), price of stock times its shares outstanding, adds to the explanation of cross-sectional expected returns provided by market betas. Fama and French (1993, 1995) use a time-series regression approach and find that two non-market risk factors SMB (the difference between the return on a portfolio of small stocks and the return on a portfolio of large stocks) and HML (the difference between the return on a portfolio of high book-to-market-value stocks and the return on a portfolio of low book-to-market-value stocks) are significant in explaining the cross sectional equity returns.

In REIT literature, Fama-French three-factor model has been applied to REIT return analysis. Peterson and Hsieh (1997) find that two non-market risk variables along with stock market data, SMB and HML, are also significant for EREIT returns. That is, Fama-French factors help explain EREIT pricing and performance. Chiang, Lee, and Wisen (2005) use the Fama-French factors to test if EREITs beta has remained constant over time. Their findings suggest the Fama-French three factors can explain EREITs pricing better than the CAPM framework. However, anomalies such as size effects in the equity market may arise from inadequacies in asset pricing models. One of the possible inadequacies in asset pricing models is the assumption of constant variance (homoskedasticity).

Because the variance of equity returns may change over time, previous models, such as the CAPM, may not capture the essence of “heteroskedasticity” in equity returns. Thus, we reexamine EREITs returns by using GARCH family models to verify if the size effect still exists when taking “heteroskedasticity” into account. In the next section, we provide a brief review of relevant literature. We then describe our data and methodology and discuss our empirical results. The final section concludes.

## LITERATURE REVIEW

Many empirical studies report different kinds of equity market irregularities. Researchers call these irregularities “anomalies”. The existence of anomalies indicates that the equity market is inefficient. The size premium for smaller companies is a well-known market anomalies. Banz (1981) argues that the CAPM is misspecified and small NYSE firms have significantly larger risk adjusted returns than large NYSE firms for the 1936-1977 periods. The findings question market efficiency and spawn hundreds of subsequent studies on the size effect.

Many previous papers have investigated the equity returns generating process and some market research provides explanations of the size effect. One explanation is the bid-ask spread. Lower-priced stocks tend to be small size stocks and have larger bid-ask spreads (Keim, 1989). This result partially explains the size effect. Another explanation is liquidity. Small size stocks tend to have less liquidity, so they would have liquidity premium to compensate the liquidity risk.

Other explanations such as transaction costs, default risk, return seasonality are documented by some papers. In addition, several papers show the size effect may depend on the macroeconomic environment. Jensen, Johnson, and Mercer (1997) indicate the importance of the size and price-to-book ratio effects depend heavily on the monetary status of the Federal Reserve from 1965 to 1994. Fama and French (1995) find that market and size factors in earnings help explain the market and size factors in returns. They indicate the size effects in earnings may be the source of the size effect in returns.

Because of infrequent trading, real estate markets are generally thought to be less efficient than the stock market. Recent papers, such as Bhasin, Cole, and Kiely (1997), and Clayton and MacKinnon (2000), provide evidence that REIT market capitalization and trading volume grow rapidly and the liquidity of REITs increase through the early and mid 1990s as a result of the growth in scope and size of the real estate securities market.

Using a regression test and a mean test, Scott (1990) finds that REIT prices do not always track market fundamentals. Kuhle and Alvayay (2000) examine 108 EREITs using autocorrelation and run tests in the period of 1989 to 1998 and find evidence of inefficiency in EREIT pricing in the form of one and two-day price dependencies. Colwell and Park (1990) examine 28 EREITs and 22 mortgage REITs (MREITs) and find evidence to support a reverse size effect and a January effect during the 1964-1986 periods. McIntosh, Liang, and Tompkins (1991) examine REIT price performance and find evidence that small firms earn higher average rates-of-return than large firms from 1974 to 1988 after taking risks into account. These results show that there exist a size effects in REIT markets.

Peterson and Hsieh (1997) use a five-factor model from Fama and French (1993) and find that risk premia on EREITs returns are significantly related to both the excess returns of market portfolios and the risk premia on mimicking portfolios for size and book-to-market equity factors. Chiang, Lee, and Wisen (2005) use the Fama-French three factors and find that the Fama-French three-factor model is more useful than the CAPM, the single-factor model, in explaining variation in EREIT returns.

Previous studies have employed time-series data to investigate how macroeconomic factors influence stock returns over time. For example, Chen, Roll and Ross (1986) find that five macroeconomic variables significantly contribute to the expected stock returns. These factors are the unanticipated inflation rate, the change in expected inflation, the unanticipated change in default risk premium, the unanticipated change in term spread, and the unanticipated change in the growth rate in industrial production. Several papers have documented that real estate returns are inversely related to interest rate movements (Chen and Tzang, 1988; Ling and Naranjo, 1998; and Devaney, 2001) and positively related to market conditions (Sagalyn, 1990; Ling and Naranjo, 1998; and Glascock, Lu and So, 2002).

Similarly, Chan, Hendershott, and Sanders (1990) specify some macroeconomic factors such as inflation, the term and risk structure of interest rates, and industrial production. They find that bond market risk premia such as the term spread, the difference between returns on long-term Treasury bonds and short-term Treasury bills, and the default risk spread, the difference between returns on high-yield corporate bonds and returns on long-term Treasury bonds, as well as changes in market capitalization of stock equities are important for explaining the average variation in REIT returns.

McCue and Kling (1994) find that prices, nominal rates, output and investment all significantly influence real estate returns. In particular, nominal interest rates contribute to the majority of variation in EREIT returns. Karolyi and Sanders (1998) conclude that changes in the prices of economic risk are more important than changes in beta in explaining the predictable variation of REITs returns.

Based on Chen et al. (1986) and Chan et al. (1990), Chen, Hsieh, Vines, and Chiou (1998) find that size and unanticipated changes in the term structure command a risk premium in EREIT pricing. Johnson and Jensen (1999) show that REIT returns are affected by monetary policy. Their results show that all kinds of NAREIT index returns during expansive periods are more than quadruple those realized during restrictive periods. Ewing and Payne (2005) indicate that shocks to monetary policy, economic growth, and inflation all lead to lower REIT returns, while a shock to the default risk premium raises REIT returns.

The size effect may arise from bid-ask spread, liquidity premium, transaction costs, default risk, return seasonality and macroeconomic factors. Besides these microeconomic and macroeconomic variables, the size effect itself may also result from inadequate pricing models. Engle (1982) develops autoregressive conditional heteroskedasticity (ARCH) model which is stimulated by the observation that the volatility of asset returns generally fluctuate. In Bollerslev (1986), the GARCH model allows the conditional variance to follow an ARMA (autoregressive moving average) process. The popularity of GARCH and ARCH models stems from the ability to capturing the so-called volatility clustering phenomenon, that is, the fact that equal intensity for volatility periods tend to cluster.

Young and Graff (1995) investigate the return distribution of individual properties in the Russell-NCREIF database. They find evidence of time-varying heteroskedasticity and skewness over the period of their study. Engle, Lilien and Robins (1987) propose the GARCH-in-mean (GARCH-M) model in order to better fit financial data which generally depend directly on its own conditional variances (volatility). Devaney (2001) finds that interest rate movements have significant influences on REITs returns. However, he also finds that in most cases the results for EREITs are not significant. Najand, Lin and Fitzgerald (2006) utilize both a GARCH and a GARCH-M model and the conditional CAPM in their analysis of daily REIT volatility. They find that GARCH-M terms are not relevant in determining the EREIT returns.

To account for such asymmetric effects, Nelson (1991) proposes the exponential GARCH (EGARCH) model. In our paper of examining size premia, the negative shocks and positive shocks indicate the size effect and the reverse size effect, respectively. We utilize EGARCH model to investigate if there is a leverage effect on size premia of EREITs returns.

## DATA

We examine monthly EREIT returns for the period of July 1995 to December 2006, which provides 138 monthly observations. According to the National Association of Real Estate Investment Trusts (NAREIT), there exists a total of 183 REITs which included 138 EREITs, 38 MREITs and 7 Hybrid REITs at the end of 2006. All EREITs in our research are selected from publicly traded REITs listed on the CRSP database, which are also members of the National Association of Real Estate Investment Trusts (NAREIT).

The EREITs are traded on NYSE, AMEX, and NASDAQ. Monthly market data and the yearly book value data of EREITs are obtained from CRSP and COMPUSTAT databases, respectively. Downs (2000) argues that asset pricing models without mimicking REITs-based portfolios may not capture enough information in REIT prices. Consequently, stock market data used in Peterson and Hsieh (1997) may not reflect proper information in REIT pricing and returns. Therefore, instead of mimicking Fama-French three factors with stock market data, our paper restructures Fama-French three factors with EREIT data.

According to the NAREIT, NAREIT equity index is more correlated with Russell 2000 index than other indexes. Thus, we use Russell 2000 index as our market return proxy. The Russell 2000 index data is from DataStream database.

Table 1: Data Description of Macroeconomic Variables, July 1995 to December 2006

| Symbol | Definition | Data Measurement |
|---|---|---|
| Panel A: Basic Data and Source | | |
| $Baa_t$ | High-yield corporate bond yield | |
| $LTB_t$ | Long-term Treasury bond yield (10-year) | |
| $STB_t$ | Short-term Treasury bill yield (3-month) | |
| $CPI_t$ | Consumer Price Index * | |
| Panel B: Derived Macroeconomic Variables | | |
| $DEF_t$ | Default risk premium | $Baa_t$ - $LTB_t$ |
| $TERM_t$ | Term structure premium | $LTB_t$ - $STB_t$ |
| $INF_t$ | Inflation rate | $(CPI_t – CPI_{t-1})/CPI_{t-1}$ |
| $STB_t$ | Short-term nominal interest rate | |

*This table shows data details of macroeconomic variables from July 1995 to December 2006. Source: Federal Reserve Statistical Release and U.S. Department of Labor (website). * Consumer Price Index is for all urban consumers. The base period is 1984=100.*

We use one of the Fama-French three factors, SMB, as a proxy for the excess returns of EREITs size. The portfolios are formed based on Fama and French (1992). Fama and French (1993) find that book-to-market equity ratio has a stronger role than size on average stock returns. Therefore, they decide to sort firms into three groups based on book-to-market equity ratios (BE/ME) and only two based on size (market value).

We categorize EREITs into three groups based on size and into two groups based on BE/ME. In each month from July 1995 to December 2006, the 33 EREITs are ranked on size (market capitalization). The ranking of the 33 EREITs is based on breakpoints for the bottom 30% (Small), middle 40% (Medium), and top 30% (Big) of market capitalization values.

The median book-to-market equity ratio is used to split EREITs into two groups, low and high (L and H). We construct 6 portfolios (S/L, S/H, M/L, M/H B/L, and B/H) from the combinations of the three size groups and two BE/ME groups. Our portfolio SMB (small minus big) is meant to mimic the risk factor in EREIT returns related to size. The SMB is measured by the difference between the simple average of the returns on the two small EREIT portfolios (S/L and S/H) and the simple average of the returns on the two big EREIT portfolios (B/L and B/H).

The portfolio HML (high minus low) is meant to mimic the risk factor in EREIT returns related to book-to-market equity ratio. The HML is evaluated by the difference between the simple average of returns on the three high BE/ME portfolios (S/H, M/H, and B/H) and the average of returns on the three low BE/ME portfolios (S/L, M/L, and B/L). The correlation between SMB and HML is 0.204 in our research. Success of the portfolio formation procedure is evident because when BE/ME (ME) is considered, ME (BE/ME) should be free to any degree of BE/ME (ME).

In choosing macroeconomic variables to include in our analysis, we find that REIT returns are affected by bond market factors and monetary policy (McCue and Kling, 1994). The variables included in our research are: default risk premium, inflation, nominal short-term interest rate, and term spread. Table 1 presents a description of these variables.

To capture the changes in corporate default risk imbedded in bond returns, we use the default risk premium. Default risk premium ($DEF_t$) is defined as the monthly returns on high-yield corporate bonds (Baa rating) less the monthly returns on long-term (10-year) U.S. Treasury bonds (Chen et al., 1986; Jensen, Mercer and Johnson, 1996). This variable may act as a signal of the future market condition expectation for investors.

The term spread ($TERM_t$) is measured by the difference between the monthly yield on long-term (10-year) Treasury bonds and the monthly yields on short-term (3-month) Treasury bills. The consumer price index (CPI) is the measure of aggregate price level (Park and Ratti, 2000). We obtain inflation rate ($INF_t$) by using the CPI series ($INF_t = [(CPI_t - CPI_{t-1})/CPI_{t-1}]$). McCue and Kling (1994) identify inflation as a key contribute to REITs returns. We use the three-month Treasury bill yield ($STB_t$), the proxy for monetary policy, to represent the nominal short-term rate, which is also used in McCue and Kling (1994). McCue and Kling (1994) indicate that shocks to nominal interest rate have a negative effect on the EREITs returns.

Descriptive statistics of monthly returns for these six portfolios, Fama-French three factors, and macroeconomic variables are outlined in Table 2. In Panel A of Table 3-2, it is interesting to note that except B/H portfolio, the other five portfolios are leptokurtic. For example, the kurtosis is 1.713684 for S/L portfolio. Eagle (1982) shows that a possible reason of leptokurtosis in the unconditional distribution is conditional heteroskedasticity.

We find that the average EREIT returns are increasing when book-to-market equity ratio is high, while low book-to-market equity ratios accompany decreasing average EREIT returns. In the high book-to-market equity ratio group, the average EREIT returns are 1.15%, 1.50%, and 1.43% for small, median and big size, respectively. In contract, in low book-to-market equity ratio group, the average EREITs returns are 1.23%, 1.82%, and 1.77% for small, median, and big size, respectively.

In Panel B of Table 2, the mean monthly size premia (SMB) is -0.41% for EREITs returns, which is equivalent to an annual premia of -4.93%. We examine SMB, the size premia proxy, according to the Jarque-Bera test. The result rejects the null hypothesis of a normal distribution. The rejection of normality implies that ARCH and GARCH models are appropriate for analyzing the size effect of EREIT data. According to Panel C, the average default risk premium, inflation rate, short-term interest rate, and term structure premium are 2.17%, 0.20%, 3.85%, and 1.35%, respectively.

Next, we examine SMB, the proxy for size premia, on EREITs returns. All of the returns are time varying with evidence of volatility clustering, which tends to concentrate together in the high levels of volatility. Table 3 shows the correlation matrix for SMB and our four macroeconomic variables. We choose Pearson correlation method to test the significance of all variables. The SMB is positively correlated with the term structure premium. The SMB is negatively correlated with the short-term interest rate.

Table 2: Summary Statistics of Monthly Returns

| | Mean | Median | Std. Deviation | Variance | Skewness | Kurtosis |
|---|---|---|---|---|---|---|
| Panel A: Six portfolios | | | | | | |
| S/L | 0.0124 | 0.0122 | 0.0356 | 0.0013 | -0.3901 | 1.7137 |
| S/H | 0.0115 | 0.0099 | 0.0446 | 0.0020 | -0.0498 | 1.1511 |
| M/L | 0.0182 | 0.0173 | 0.0410 | 0.0017 | -0.3117 | 2.3550 |
| M/H | 0.0150 | 0.0161 | 0.0434 | 0.0019 | -0.4455 | 0.7808 |
| B/L | 0.0177 | 0.0192 | 0.0406 | 0.0017 | -0.6938 | 2.2266 |
| B/H | 0.0143 | 0.0183 | 0.0409 | 0.0017 | -0.6105 | 1.8800 |
| Panel B: Fama-French three-factor model | | | | | | |
| SMB | -0.0041 | -0.0050 | 0.0370 | 0.0014 | -0.1164 | 0.4025 |
| HML | -0.0073 | -0.0132 | 0.0576 | 0.0033 | 0.5208 | 1.6558 |
| MER* | 0.0031 | 0.0018 | 0.0618 | 0.0038 | 0.0811 | 0.9786 |
| NAR** | 0.0127 | 0.0167 | 0.0386 | 0.0015 | -0.7847 | 2.2371 |
| Panel C: Macroeconomic Variables | | | | | | |
| DEF | 0.0217 | 0.0208 | 0.0058 | 0.0000 | 0.7133 | -0.4983 |
| TERM | 0.0135 | 0.0110 | 0.0116 | 0.0001 | 0.4177 | -0.9026 |
| INF | 0.0020 | 0.0019 | 0.0030 | 0.0000 | -0.0383 | 1.0683 |
| STB | 0.0385 | 0.0466 | 0.0172 | 0.0003 | -0.5438 | -1.2563 |

*This table shows summary statistics of monthly returns for different models. The sample period is July 1995 to December 2006, which provides 138 observations. * MER (market excess returns) is Russell 2000 index monthly returns minus risk-free rate (3-month Treasury bill yield). **NAR is equity NAREIT index returns.*

Table 3: Correlation Matrix for SMB and Macroeconomic Variables

| Z | SMB | DEF | TERM | INF | STB |
|---|---|---|---|---|---|
| SMB | 1 | - | - | - | - |
| DEF | -0.2723 (0.00124)* | 1 | - | - | - |
| TERM | 0.0375 (0.6625) | -0.0429 (0.6173) | 1 | - | - |
| INF | -0.2466 (0.0050)* | -0.0332 (0.7099) | 0.0163 (0.8553) | 1 | - |
| STB | 0.0397 (0.6439) | -0.1909 (0.0250)** | -0.2013 (0.0179)** | 0.1703 (0.0546)*** | 1 |

*This table shows correlation matrix for SMB and macroeconomic variables. SMB = small size EREITs returns minus big size EREITs return; DEF = default risk premium; INF = inflation rate; STB = short-term interest rate; and TERM = term structure premium. * Significant at the 1% level; ** Significant at the 5% level; and *** Significant at the 10% level.*

## METHODOLGY

Our study investigates volatility behavior of size effect, the major anomaly, on EREIT returns over the period of July 1995 to December 2006 by using GARCH family models. We implement the unit root test to identify the appropriate forms of our variables. The results are reported in Table 4. We find that both SMB and the market excess returns (MER) are stationary (no unit root) variables at the 1% significant level. We also, we find that the residual conforms to white noise process in 0 lagged differences.

Table 4: Results for Unit Root Test

| Augmented Dickey-Fuller Test Statistic | | | | |
|---|---|---|---|---|
| Variables | *level (k=0)* | *level (k=1)* | *level (k=2)* | *level (k=3)* |
| SMB | -13.28 | -8.388 | -6.359 | -5.298 |
| | (0.0000)* | (0.0000)* | (0.0000)* | (0.0000)* |
| MER | -10.92 | -8.079 | -7.027 | -6.194 |
| | (0.0000)* | (0.0000)* | (0.0000)* | (0.0000)* |

*This table shows the results for unit root test. k is the lagged differences. The selected p-value method is MacKinnon (1996) one-sided p-values. The null hypothesis in the Augmented Dickey-Fuller test is that there exists a unit root in time series, that is, the time series is non-stationary. The test critical values are the Mackinnon critical values for rejection of hypothesis. * Significant at the 1% level.*

The size effect anomaly may result from using the incorrect asset pricing model because the variance of size premia changes over time. The change of size effects implies the risk premia of size effects may vary over time. Therefore, we use the GARCH model to investigate whether the size effect in EREIT markets still exists when heteroskedasticity is taken into account (Equation 1) and if there are time-varying risk premia due to the volatility of size effect (Equation 2). The GARCH process model can be formulated as follows:

$$NAR_t = a_0 + bR_t + cHML_t + \varepsilon_t \tag{1}$$

$$R_t = a_0 + bR_{t-1} + \varepsilon_t, \tag{2}$$

Where

$$\varepsilon_t^2 | \Phi_{t-1} \ \mathrm{GED}(0, h_t^2, v)$$

$$h_t^2 = \alpha_0 + \sum_{i=1}^{q} \alpha_i \varepsilon_{t-i}^2 + \sum_{j=1}^{p} \beta_j h_{t-j}^2$$

$NAR_t$ is the equity NAREIT index returns. $R_t$, the proxy for SMB, is the small size risk premium computed by taking the difference between equally-weighted small size portfolio returns and equally-weighted big size portfolio returns in period of *t*, $\Phi_{t-1}$ is the information set at time *t-1*, $a_0$ is a constant term, $\varepsilon_t$ is a random error, $h_t^2$ is the conditional variance of $\varepsilon_t$, $p$ is the order of the GARCH term and $q$ is the order of the ARCH term. The generalized error distribution (GED) is written as follows:

$$f(\varepsilon_t) = \frac{v\, exp\left[-(1/2)\left|\varepsilon_t h_t^{-1/2}/\lambda\right|^v\right]}{\lambda 2^{[v+1/v]} \Gamma(1/v)} h_t^{-1/2}, \tag{3}$$

where $\lambda = \left[\frac{2^{(-2/v)} \Gamma(1/v)}{\Gamma(2/v)}\right]^{1/2}$.

$\Gamma(\cdot)$ is the gamma function and $\nu$ is a scale parameter controlling the shape of GED distribution, that is,

$\nu$ is a measure of tail thickness, which is equal to 2 for the normal distribution and less than 2 for the leptokurtic distribution.

We further investigate the relationship between SMB, the size effect proxy of EREITs and its expected volatility using GARCH-M model. GARCH-M model captures thick tailed returns and volatility clustering. GARCH-M model is different from the standard GARCH method in that it expresses present volatility as a function of past volatility and returns behavior by introducing past conditional variance into the mean equation. GARCH-M model formation can be expressed as follows:

$$R_t = a_0 + bR_{t-1} + ch_t^2 + \varepsilon_t, \quad (4)$$

where

$\varepsilon_t^2 | \Phi_{t-1}$ GED$(0, h_t^2, \nu)$

$$h_t^2 = \alpha_0 + \sum_{i=1}^{q} \alpha_i \varepsilon_{t-i}^2 + \sum_{j=1}^{p} \beta_j h_{t-j}^2$$

Nelson (1991) develops an exponential GARCH (EGARCH) model in which the conditional variance is allowed to depend on the magnitude and sign of the innovations (i.e., error term).

Negative and positive innovations can have different impacts on expectation errors (leverage effect). Cotter and Stevenson (2007) examine asymmetry in REITs volatility using a univariate EGARCH model. They find no evidence of a leverage effect on REIT returns. We use an EGARCH model to examine whether the sign of the past size effect has an influence on predicting future volatility of the size premia on EREIT returns. The EGARCH model is outlined as follows:

$$R_t = a_0 + bR_{t-1} + \varepsilon_t, \quad (5)$$

Where

$\varepsilon_t^2 | \Phi_{t-1}$ GED$(0, h_t^2, \nu)$

$$\ln h_t^2 = \alpha_0 + \sum_{i=1}^{q} \alpha_i \left( \frac{\varepsilon_{t-i}}{h_{t-i}^{0.5}} \right) + \sum_{i=1}^{q} \gamma_i \left| \frac{\varepsilon_{t-i}}{h_{t-i}^{0.5}} \right| + \sum_{j=1}^{p} \beta_j \ln\left(h_{t-1}^2\right).$$

Evidence that the size premium conditional volatility for EREITs changes over time leads us to investigate why it changes. We use the unrestricted VAR model to test whether the variance of size premium, SMB, for EREITs returns results from the changes in macroeconomic factors. The unrestricted VAR model can be defined using five equations as follows:

$$\Delta SMB_t = \sum_{i=1}^{k} \lambda_{11}^i \Delta SMB_{t-i} + \sum_{i=1}^{k} \lambda_{12}^i \Delta DEF_{t-i} + \sum_{i=1}^{k} \lambda_{13}^i \Delta TERM_{t-i} + \sum_{i=1}^{k} \lambda_{14}^i \Delta INF_{t-i} + \sum_{i=1}^{k} \lambda_{15}^i \Delta STB_{t-i} + \varepsilon_{1t} \quad (6)$$

$\Delta DEF_t =$

$$\sum_{i=1}^{k}\lambda_{21}^{i}\Delta SMB_{t-i}+\sum_{i=1}^{k}\lambda_{22}^{i}\Delta DEF_{t-i}+\sum_{i=1}^{k}\lambda_{23}^{i}\Delta TERM_{t-i}+\sum_{i=1}^{k}\lambda_{24}^{i}\Delta INF_{t-i}+\sum_{i=1}^{k}\lambda_{25}^{i}\Delta STB_{t-i}+\varepsilon_{2t} \quad (7)$$

$$\Delta TERM_t=\sum_{i=1}^{k}\lambda_{31}^{i}\Delta SMB_{t-i}+\sum_{i=1}^{k}\lambda_{32}^{i}\Delta DEF_{t-i}+\sum_{i=1}^{k}\lambda_{33}^{i}\Delta TERM_{t-i}+\sum_{i=1}^{k}\lambda_{34}^{i}\Delta INF_{t-i}+\sum_{i=1}^{k}\lambda_{35}^{i}\Delta STB_{t-i}+\varepsilon_{3t} \quad (8)$$

$$\Delta INF_t=\sum_{i=1}^{k}\lambda_{41}^{i}\Delta SMB_{t-i}+\sum_{i=1}^{k}\lambda_{42}^{i}\Delta DEF_{t-i}+\sum_{i=1}^{k}\lambda_{43}^{i}\Delta TERM_{t-i}+\sum_{i=1}^{k}\lambda_{44}^{i}\Delta INF_{t-i}+\sum_{i=1}^{k}\lambda_{45}^{i}\Delta STB_{t-i}+\varepsilon_{4t} \quad (9)$$

$$\Delta STB_t=\sum_{i=1}^{k}\lambda_{51}^{i}\Delta SMB_{t-i}+\sum_{i=1}^{k}\lambda_{52}^{i}\Delta DEF_{t-i}+\sum_{i=1}^{k}\lambda_{53}^{i}\Delta TERM_{t-i}+\sum_{i=1}^{k}\lambda_{54}^{i}\Delta INF_{t-i}+\sum_{i=1}^{k}\lambda_{55}^{i}\Delta STB_{t-i}+\varepsilon_{5t} \quad (10)$$

where $\lambda_{nm}^{i}$ is parameter to be estimated, $k$ is the maximum distributed lag length, $\varepsilon_{nt}$ is independent and identically distributed error term, and $\Delta$ means the conditional variable volatility at time $t$. The Fama-French three-factor model is defined as:

$$NAR_t=\lambda_0+\lambda_1(R_{mt}-R_{ft})+\lambda_2 SMB_t+\lambda_3 HML_t+\varepsilon_t, \quad (11)$$

where $NAR_t$ is equity NAREIT index returns, $R_{mt}$ is Russell 2000 index returns, $R_{ft}$ is three-month Treasury bill yield, and $\lambda$ is the parameter which will be estimated.

## EMPIRICAL RESULTS

The regression result is shown in Table 5. While using a single variable, SMB or HML is not significant. SMB is significant when using three factors together. The $R^2$ is 0.5075, which is substantially greater than the $R^2$ obtained using each of the three factors individually. The results indicate that the Fama-French three-factor model is an appropriate model for EREIT markets. SMB, the proxy for the size premia, is significant at the 1% level and the coefficient of SMB ($\lambda_2$) is negative in the Fama-French three-factor model for EREITs market period.

Table 5: Results from Fama-French Three-Factor Model, July 1995 to December 2006

| | Coefficient | Std. Error | t-Statistic | Sig. |
|---|---|---|---|---|
| $\lambda_0$ | 0.0201 | 0.0027 | 7.5623 | 0.0000* |
| $\lambda_1$ | 0.2991 | 0.0395 | 7.5780 | 0.0000* |
| $\lambda_2$ | -0.5720 | 0.0642 | -8.914 | 0.0000* |
| $\lambda_3$ | 0.0580 | 0.0539 | 1.0764 | 0.2837 |

*This table shows the regression result. The dependent variable is equity NAREIT index (NAR). The sample size is 138 observations. * Significant at the 1% level and ** Significant at the 5% level.*

Our finding suggests that large cap EREITs significantly outperformed small cap EREITs during July 1995 to December 2006. The significance of SMB supports evidence of a size effect on EREIT returns.

Our results are in accordance to McIntosh, Liang, and Tompkins (1991). Previous studies show that book-to-market equity ratio is not significant for EREIT returns from 1978 to 1994 (Chen, Hsieh, Vines, and Chiou, 1998).

The results for equation (1) are shown in Table 6. We examine whether the size premium on EREIT returns still exists when heteroskedasticity is taken into account. The coefficient $b$ is significant at the 1% level and the coefficient is negative. This result proves that the size effect still exists when heteroskedasticity is taken into account.

Table 6: Results from GARCH Model for REITs Returns, July 1995

| | Coefficient | Std. Error | z-Statistic | Sig. |
|---|---|---|---|---|
| Mean equation | | | | |
| $a_0$ | 0.0115 | 0.0028 | 4.1058 | 0.0000* |
| $b$ | -0.5237 | 0.0653 | -8.015 | 0.0000* |
| $c$ | 0.1050 | 0.0428 | 2.451 | 0.0143** |
| Variance equation | | | | |
| $\alpha_0$ | 0.0000 | 0.0000 | 8.240 | 0.0000* |
| $\alpha_1$ | -0.0522 | 0.0216 | -2.415 | 0.0158** |
| $\beta_1$ | 1.036 | 0.0244 | 42.42 | 0.0000* |

*This table shows the results from GARCH model for REITs Returns in July 1995. The sample size contains 138 observations. * Significant at the 1% level and ** Significant at the 5% level.*

Table 7 reports the parameter estimates and $p$-value for SMB, the size premium proxy, on EREIT returns based on the models presented in equations (2) to (5). These models test the relationship between excess returns and the size premia. Bollerslev, Chou, and Kroner (1992) find that GARCH (1, 1) fits well for most financial time series data. Thus, we adopt GARCH ($p$, $q$) which sets $p = q = 1$ in our study. In each GARCH family model, we find that all GED parameters ($\nu$) are less that 2, which indicate that the distribution of SMB is leptokurtic. This result is in accordance with our summary statistics. From the mean equation the of GARCH (1,1) model, we note that the first-order autoregressive coefficient ($b$) is significant during our sample period. The result indicates that the risk premia for the size effect may be affected by its lagged values.

The result also implies that EREITs market is not efficient from July 1995 to December 2006. The long-term variance for size premia is 0.0001 during our sample period. The GARCH-estimated parameter $\beta_1$ is statistically significant in the conditional variance equation. This result supports our assumption that the size premia on EREITs returns may be time-varying and may be affected by its past volatility. We also find that $\beta_1$ (0.8780) is larger than $\alpha_1$ (0.0750). This finding suggests the current volatility to risk premia for size effect leads only small revisions to future volatility.

The sum of the ARCH and GARCH term $(\alpha_1 + \beta_1)$ can be interpreted as the measure of volatility consistence. In a GARCH (1, 1) model, $\alpha_1 + \beta_1$ (0.9530) is less than one and is very close to 1, indicating the volatility shocks to the size premia on EREIT returns are quite persistent. The result of persistent volatility implies that the risk premia for the size effect of the same magnitude tend to cluster over time.

If the expected size premia for EREITs can be predicted based on the forecasted variance, the coefficient $c$, the measurement of risk premia, should be significantly positive. In our test, the coefficient $c$ is 0.6251. From July 1995 to December 2006, this result supports that there exists a positive relationship between the expected size premia and conditional variance in EREIT markets. This result also implies that the conditional volatility has explanatory power for EREIT returns.

Table 7: Estimates for GARCH Family Models for SMB on EREITs returns

| | GARCH(1,1) | GRACH-M (1,1) | EGARCH (1, 1) |
|---|---|---|---|
| | Coefficient (*p*-value) | | |
| Mean equation | | | |
| $a_0$ | -0.0047 (0.0712)*** | -0.0267 (0.2608) | -0.0051 (0.0479)** |
| $b$ | -0.1669 (0.0524)*** | -0.1752 (0.0375)** | -0.1177 (0.1538) |
| $c$ | | 0.6251 (0.3400) | |
| Variance equation | | | |
| $\alpha_0$ | 0.0001 (0.5976) | 0.000 (0.5959) | -10.93 (0.0000)* |
| $\alpha_1$ | 0.0750 (0.3515) | 0.0617 (0.3081) | -0.1666 (0.1745) |
| $\beta_1$ | 0.8780 (0.0000)* | 0.8993 (0.0000)* | -0.5904 (0.0158)** |
| $\gamma_1$ | | | 0.3603 (0.0664)*** |
| GED parameter | | | |
| $\nu$ | 1.580 | 1.461 | 1.810 |
| $\alpha_1 + \beta_1$ | 0.9530 | 0.9610 | -0.7570 |
| Log likelihood ratio | 263.6 | 264.3 | 265.0 |

*This table shows the estimates for GARCH family models for SMB on EREITs returns. The conditional variance process for GARCH model is* $\sigma_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 \sigma_{t-1}^2$. *The sample period is from July 1995 to December 2006. There are 138 observations. * Significant at the 1% level; ** Significant at the 5% level; and *** Significant at the 10% level.*

Devaney (2001) finds that the GARCH-M specification is more suitable for the mortgage REIT portfolios returns than for EREIT portfolio returns. However, according to our results, the GARCH-M model may be appropriate for the size premia on EREIT returns. In the EGARCH (1, 1) model, $\alpha_1$ is not significant at the 1%, 5% or 10% significant level. This result suggests the leverage effect does not exist on the size premia for EREIT returns. This finding implies the size effect and reverse size effect do not have different impacts on predicting future size premia volatility on EREIT returns.

We perform several diagnostic statistics for each GARCH family model. The results are reported in Table

8. We employ the Ljung-Box Q test for the standardized residual ($\varepsilon_t/h_t^{0.5}$) and the squared-standardized residual ($\varepsilon_t^2/h_t$). All series are free of serial correlation for 24 lags at the 1% significance level except for the EGARCH model. The Jarque-Bera test all shows that the null hypothesis of normality is rejected. Finally, we carry out a Lagrange Multiplier test (LM test) to examine whether the standardized residuals exhibit additional autoregression in the conditional variance. We find that in all GARCH family models, there are no additional autoregressive structures remaining in the standardized residuals for conditional variance equation.

Table 8: Diagnostic Statistics of the Residuals of GARCH Family Models

| | GARCH (1, 1) | GARCH-M (1, 1 ) | EGARCH |
|---|---|---|---|
| $\varepsilon_t / h_t^{0.5}$- standardized residuals | | | |
| Ljung-Box (24 lags) | 25.78 | 26.67 | 30.59 |
| | (0.3110) | (0.2700) | (0.1330) |
| Jarque-Bera test | 3.144 | 5.987 | 0.0640 |
| | (0.2080) | (0.0520)*** | (0.9690) |
| Lagrange Multiplier test | 0.9490 | 0.3977 | 0.0088 |
| | (0.3320) | (0.5290) | (0.9250) |
| $\varepsilon_t^{2}/ h_t$ - squared standardized residuals | | | |
| Ljung-Box (24 lag) | 24.58 | 19.17 | 32.90 |
| | (0.3720) | (0.6910) | (0.0830)*** |

*This table shows the diagnostic statistics of the residuals of GARCH family models. The null hypothesis of Ljung-Box is no serial correlation; the null hypothesis of Jarque-Bera test is normal distribution; and the null hypothesis of LM test is no autoregressive structure for residuals in the conditional variance equation. * Significant at the 1% level.*

We begin with first-order VAR model at $k$ = 1. We then use Ljung-Box Q test to examine whether the residuals are auto-correlative. From the results for Ljung-Box Q test, we find that when using a 3-order VAR model, the residuals show no autocorrelation. Therefore, we choose the 3-order unrestricted VAR model in our study. Table 9 shows the parameter estimates and $t$-statistics of the VAR model.

It is important to note that these results provide information as to how the dependent variable (△SMB) is influenced by state variables (△DEF, △TERM, △INF, and △STB) in the EREIT market. We find that there are only three lagged significant values: the 1st lagged and the 3rd lagged conditional volatility of size premia (△SMB), and the 1st lagged conditional volatility of default risk premium (△DEF). All the 1st lagged and the 3rd lagged of △SMB and the 1st lagged of △DEF have positive effect on the △SMB.

Although the lagged conditional volatilities of △TERM, △INF and △STB are insignificant in the equation for △SMB, it would be misleading to conclude that these two macroeconomic variables have no influence on the volatility of the size premia for the EREIT market. That is, the VAR result does not imply that a shock (unexpected change) has no impact on the change in size premia (△SMB).

Table 9: The Estimators for 3-order Unrestricted VAR Model

| | △SMB | △DEF | △TERM | △INF | △STB |
|---|---|---|---|---|---|
| C | 0.0001 | -0.0000 | 0.0000 | 0.0000 | -0.0000 |
| | (1.777)** | (-0.0378) | (1.151) | (1.510)*** | (-1.636) |
| △SMB(-1) | 0.9308 | -0.0019 | -0.0008 | -0.0011 | -0.0009 |
| | (10.22)* | (-1.080) | (-0.2328) | (-1.297) | (-0.1763) |
| △SMB(-2) | -0.2252 | 0.0013 | 0.0002 | 0.0007 | 0.0052 |
| | (-1.823) | (0.5394) | (0.0378) | (0.5525) | (0.7942) |
| △SMB(-3) | 0.2118 | 0.0016 | 0.0016 | 0.0000 | 0.0000 |
| | (2.381)* | (0.9457) | (0.4951) | (0.1172) | (0.0147) |
| △DEF(-1) | 12.52 | 0.7949 | 0.3154 | 0.0193 | 0.9381 |
| | (2.478)* | (8.295)* | (1.736)** | (0.4152) | (3.523)* |
| △DEF(-2) | -13.78 | -0.2252 | -0.2814 | -0.0406 | -0.8387 |
| | (-2.197) | (-1.893) | (-1.248) | (-0.7051) | (-2.537) |
| △DEF(-3) | 4.469 | 0.0257 | 0.2363 | 0.0602 | 0.0853 |
| | (0.8417) | (0.2554) | (1.238) | (1.236) | (0.3048) |
| △TERM(-1) | 3.761 | 0.0045 | 0.1492 | -0.0511 | 0.3262 |
| | (1.127) | (0.0713) | (1.244) | (-1.670) | (1.855)** |
| △TERM(-2) | -1.231 | 0.0276 | -0.2528 | -0.0232 | -0.2264 |
| | (-0.3645) | (0.4312) | (-2.082) | (-0.7494) | (-1.272) |
| △TERM(-3) | 0.0080 | -0.0711 | -0.0089 | 0.0381 | -0.1557 |
| | (0.0024) | (-1.110) | (-0.0734) | (1.228) | (-0.8743) |
| △INF(-1) | -0.0986 | -0.1757 | -0.0381 | 0.8771 | 0.0268 |
| | (-0.0106) | (-0.9977) | (-0.1140) | (10.29)* | (0.0548) |
| △INF(-2) | -3.605 | 0.2861 | 0.4383 | 0.4960 | 1.161 |
| | (-0.3078) | (1.289)*** | (1.041) | (4.615)* | (1.880)** |
| △INF(-3) | -1.342 | -0.0812 | -0.0770 | -0.4156 | -0.6274 |
| | (-0.1435) | (-0.4581) | (-0.2291) | (-4.843) | (-1.273) |
| △STB(-1) | -2.479 | -0.0707 | 0.0371 | 0.0601 | 0.2002 |
| | (-1.073) | (-1.616) | (0.4469) | (2.836)* | (1.645)** |
| △STB(-2) | 2.459 | 0.0323 | 0.0223 | -0.0124 | 0.0323 |
| | (1.036) | (0.7174) | (0.2612) | (-0.5677) | (0.2580) |
| △STB(-3) | -1.563 | -0.0157 | 0.0006 | -0.0363 | -0.0413 |
| | (-0.72310 | (-0.3828) | (0.0073) | (-1.832) | (-0.3624) |
| F-statistic | 67.18 | 6.571 | 2.017 | 155.7 | 4.924 |

*This table shows the estimators for 3-order unrestricted VAR model. * Significant at the 1% level; ** Significant at the 5% level; and *** Significant at the 10% level.*

## CONCLUSION

Our study identifies a reverse size effect in the EREIT market (Colwell and Park, 1990) from July 1995 to December 2006 by employing the Fama-French three-factor model. However, McIntosh, Liang, and Tompkins (1991) argue that there is a size effect on REITs returns from 1974 to 1988. As changes in size effects are uncertain, we adopt the GARCH family models, namely GARCH, GARCH-M and EGARCH models to examine if there are time-varying risk premia due to volatility of the size effect. These models, with heteroskedasticity, avoid possible measurement errors which may occur in models with homoskedasticity by assuming a time-varying structure of non-systematic risk. We also adopt the VAR model to examine the relationship of conditional volatility between four macroeconomic variables and the size premia.

Our results show the GARCH effect is significant for the size premia on EREIT returns. This provides evidence that the risk premia for size effect on EREITs returns are not stable over time. This implies that the volatility of size premia on EREITs returns have no regular patterns. We also find that the past size effects or the reverse size effects have the same impact on predicting future size effect volatility on EREIT returns. This evidence indicates that there is no leverage effect in risk premia of size effects on EREIT returns.

Moreover, the result of our VAR model shows that the variation of the size premia partially results from volatility of the term spread in the bond market and short-term interest rate. In addition, the unexpected shock from the fluctuation of term spread in bond market lowers the volatility of the size premia on EREIT returns. Investors may benefit from investing in large size EREITs when bond market participants have exceptionally different expectations about interest rate movements. Finally, the large size EREITs are a good investment vehicle when default risk premium fluctuates dramatically.

For future research, the regime-switch GARCH model (Gray, 1996) which has been improved by Klaassen (2002) and Haas, Mittnik, and Paolella (2004) may be used for solving the problem of misestimating GARCH parameters such that they imply too much persistence in volatility due to the structural change in the unconditional variance (Lamoureux and Lastrapes, 1990).

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