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Understanding Equity REITs Returns: An Investment-Based Approach

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ABSTRACT

I find that a model that includes the market factor, the investment factor and the profitability factor (Hou, Xie, and Zhang, 2015) is capable of capturing average returns on Equity REIT (EREIT). The real estate premium in a two-factor model including the market factor and the real estate factor disappears when I account for premiums associated with the investment factor and profitability factor. The results from testing integration and segmentation hypotheses suggest that the market for EREIT may be only partially integrated with the general stock market. Based on integration R-squares (Cotter, Gabriel, and Roll, 2015), I report that the degree of integration between EREIT and the general stock market moves counter-cyclically. The integration is positively related to the unemployment rate and the default spread but negatively related to real estate loan growth.

JEL Classifications: G12, R3

Keywords: asset pricing models; GARCH-means model; integration R-square; financial crisis

I. INTRODUCTION

Real estate is considered a distinct and essential asset class for investors and mutual managers. Among the real estate assets, the stock exchange-listed real estate investment trusts (REITs) offer a convenient investment vehicle. According to the National Association of Real Estate Investment Trusts, more than 70 million Americans invest in REITs through their retirement savings and other investment funds. Despite the prevalence of using REITs as an investment vehicle, whether a real estate portfolio like an EREIT portfolio earns a risk premium and whether the real estate market is integrated with or segmented from the general stock market remain controversial issues in the real estate and finance literature.

In this article, I first test the market integration and segmentation hypotheses using investment-based risk factors. Hou, Xie, and Zhang (2015) document that a model including an investment factor and a profitability factor performs at least comparable to, and in many cases better than that of the models of Fama-French (1993) and Carhart (1997). Fama and French (2016) find that a model including profitability and investment factors captures several average-return anomalies. For the sake of tractability, I use a three-factor model that include the investment factor and the profitability factor used by Hou et al. (2015) or Fama and French (2016) and a four-factor model that includes EREIT as an additional factor to study the existence of the real estate premium and market integration or segmentation.

I test the market integration hypothesis in two ways. First, I examine the significance of the risk price associated with the EREIT. Second, I test the law of one price by testing restrictions on risk prices across factor portfolios and EREIT portfolio. I find that, when the investment factor and the profitability factor are included, the risk price associated with EREIT is insignificant and the law of one price is not rejected. The evidence therefore does not reject the market integration hypothesis. The result is also consistent with the hypothesis that the expected return on EREIT is explained by the risk premiums associated with the stock market, the investment factor and profitability factor.

Market segmentation requires that expected returns on factor portfolios are only related to their covariances with the factor portfolios in the three-factor model while the expected return on EREIT is only related to its own volatility. By testing restrictions on the risk prices in that four-factor model, I find that, when the investment factor and the profitability factor are included, the market segmentation hypothesis is not rejected. As a result, the results from testing integration and segmentation hypotheses together suggest that the EREIT market may be partially integrated with the stock market. The results are in contrast with the previous literature that documents that the market for securitized real estate assets like EREITs are fully integrated with the general stock market (Liu, et al., 1990; Ling and Naranjo, 1999; and Li, 2016).

Next, I measure time-varying integration. Cotter, Gabriel, and Roll (2015) obtain a measure of integration by employing R-squares from regressions with moving windows. The housing markets are regarded as highly integrated if identical U.S. national factors explain a large portion of the variance in metro-specific housing returns. Extending this idea, I use a GARCH-in-means method to conduct tests of market integration and segmentation hypotheses and estimate the integration R-squares in a unified way. The GARCH-in-means method used in this paper produces time-varying betas (factor loadings) based on a structural model for the second moments (variance-covariance matrix), unlike regressions with moving windows in which the time-varying betas are estimated under the inconsistent assumption of constant betas.

Despite the difference in the estimation techniques, I find that the integration of the EREIT market with the general stock market trends upward during the boom period 2001-2010 and then trends downward afterwards, similar to what Cotter, Gabriel, and Roll (2015) observe in the housing markets. However, I observe more frequent and larger swings in the level of integration in the EREIT market than the housing markets. For example, I uncover 3-4 cyclical movements of the integration R-square in the EREIT market in the 1972-2001 period (before the latest 2001-2010 cycle), while the national housing market integration is relatively stable in the 1992-2001 period. In addition, the integration varies between 5-10 and 80 percent in the EREIT market but between 35 and 65 percent in housing markets. Especially in the years surrounding the latest financial crisis, the market for REITs is extremely highly integrated with the general stock market.

Finally, to understand the economic causes of the time variation of market integration, I examine the relation between the integration R-square and business-cycle related macroeconomic variables, including the unemployment rate, real estate loan growth and the default spread (Fama and French, 1989; Bernanke, Gertler, and Gilchrist 1996). The variables represent the state of the real economic activity, the change in credit availability and the risk in credit markets. I find that integration is positively related to the unemployment rate and the default spread but negatively related to real estate loan growth.

The reminder of the article is organized as follows. After a brief literature review, I describe the conditional multifactor asset pricing models and the multivariate GARCH model. Then I describe the monthly data of the EREIT index and the risk factors. I present empirical results of testing market integration and segmentation and analyzing the implied time-varying integration before conclusions.

II. LITERITURE REVIEW

Using CAPM or multifactor asset pricing models, the literature produces mixed evidence on real estate premiums and market segmentation. Liu, et al. (1990) make a first attempt to link real estate risk premiums with market segmentation. Liu and Mei (1992) and Mei and Lee (1994) point out that whether a real estate portfolio is a priced risk factor has important implications for investment and portfolio management. If a real estate portfolio such as an EREIT portfolio is a priced risk factor, some kind of real estate exposure is needed for capturing the real estate factor premiums. Otherwise, if returns on real estate portfolios are explained by systematic risks associated with other factor portfolios that form mean-variance efficiency portfolios (Huberman and Kandel, 1987), these portfolios are only useful in diversifying away some idiosyncratic risk. In a different approach, Ambrose, Ancel, and Griffiths (1992), Okunev and Wilson (1997), Chaudhry, Myer, and Webb (1999) study non-linear dependency or co-integration but drew differing conclusions. Liow and Yang (2005) report results supporting for fractional cointegration between securitized real estate price, stock market price and key macroeconomic factors in some economies. Lin and Lin (2011) conduct co-integration tests and report some evidence of partial integration between securitized real estate and equity markets in several Asian markets. Loo, Anuar, and Ramakrishnan (2016) investigate the long-run relationship and short-term linkage between the Asian REIT markets and their respective macroeconomic variables, using Johansen cointegration test and Granger causality test. They find that he emerging REIT markets are more sensitive towards the change in macroeconomic environment in relative to the developed REIT markets.

III. MODELS

In this section, I describe the conditional multifactor asset pricing model and the multivariate GARCH model. Merton (1973) shows that, in an intertemporal setting, investors need to hedge against changing investment opportunities. As a result, the expected excess return on any asset is a function of its covariances with returns on the market portfolio and a number of hedging portfolios. I present the covariance-based pricing model for the purpose of empirical estimations of the expected returns of REITs.

A. The Covariance-based Model

Consider the following covariance-based pricing model:

$$E_{t-1}(R_{it}^{e}) = \sum_{j=1}^{K} \lambda_j \mathrm{cov}_{t-1}(R_{it}^{e}, F_{jt}), \tag{1}$$

where R_{it}^{e} is the excess return for period t on the i[th] asset for $i = 1,2, N, F_{jt}$ is the j[th] risk factor for $j = 1, , K$, and λ_j represents the price of the covariance risk of each asset with the j[th] risk factor. In the covariance-based model in equation (1), the risk premium for each asset associated with a factor is the covariance of the asset's return with the factor multiplied by the price of the covariance risk. For simplicity, the risk prices are assumed to be constant. The expected returns and covariances in equation (1) are both conditional on information at time t-1. Equation (1) is specialized to the conditional CAPM if the return on the market portfolio is the single factor (K=1).

The excess return refers to the return on an asset minus the risk-free rate, or the difference between returns on two portfolios, such as the small and big stock portfolios. In the remainder of the paper, I assume that the first three excess returns are returns on three factor portfolios and the fourth portfolio (N = 4) is the EREIT. Let ε_t be the $N \times 1$ vector of unexpected returns with the i[th] element ε_{it} and the $N \times N$ conditional variance-covariance matrix H_t, whose (i, j)th element is $h_{ij,t}$. The return in equation (1) for each of the first three portfolios is given by

$$R_{it}^{e} = \sum_{j=1}^{k} \lambda_j h_{ij,t} + \varepsilon_{it}, \ i=1,2,3 \tag{2}$$

and the excess return on the EREIT portfolio is given by

$$R_{N,t}^{e} = a_{t\leq T} I_{t\leq T} + a_{t>T} I_{t>T} + \sum_{j=1}^{k} \lambda_j h_{Nj,t} + \varepsilon_{N,t} \tag{3}$$

where $a_{t\leq T}$ is the intercept for the period before date T and $a_{t>T}$ is the intercept for the period after date T. I include the indicator (dummy) variable $I_{t\leq T}(I_{t>T})$, which takes the value of unity for $t \leq T(t > T)$ and zero otherwise, to detect changes in returns resulting from a possible structural break or change in regimes unexplained by the pricing model.

The conditional variance-covariance matrix H_t follows the asymmetric BEKK GARCH specification:

$$\mathbf{H_t} = \mathbf{C'C} + \mathbf{A'H_{t-1}A} + \mathbf{B'\varepsilon_{t-1}\varepsilon_{t-1}'B} + \mathbf{D'\eta_{t-1}\eta_{t-1}'D}, \tag{4}$$

where **A**, **B** and **D** are $\mathrm{N} \times \mathrm{N}$ coefficient matrices, **C** is a lower triangular matrix with $\mathrm{N} \times (\mathrm{N}+1)/2$ parameters, and $\boldsymbol{\eta}_{\mathbf{t-1}}$ is a $\mathrm{N} \times 1$ vector with i^{th} element given by $\eta_{\mathrm{i,t-1}} = -\varepsilon_{\mathrm{i,t-1}}$ if $\varepsilon_{\mathrm{i,t-1}} < 0$ and zero otherwise. Without the last term, equation (4) is the BEKK process proposed by Engle and Kroner (1995). I choose the asymmetric BEKK process rather than the asymmetric dynamic conditional correlation process (Engle, 2002), since researchers (Guo et al., 2009) find insignificant differences between the two processes.

The specification in equation (4) is appealing because it directly imposes positive definiteness on the variance-covariance matrix. However, the estimation becomes difficult if the multivariate GARCH model is applied to multiple assets. Since this paper uses four assets (N = 4), it is necessary to make simplifying assumptions to limit the dimension of parameter space. For this reason, **A**, **B** and **D** are restricted to be diagonal matrices, following De Santis and Gerard (1997) and Hardouvelis, Malliaropulos, and Priestley (2006). I estimate the multivariate GARCH-in-means equations (2)-(4) using the method of quasi-maximum likelihood, assuming a normal conditional density. The standard errors are robust to non-normality of disturbance terms.

B. Testing Capital Market Integration

Testing capital market integration is often characterized as a test of the joint hypothesis about an asset pricing or other economic model and the market integration hypothesis. I consider one- to four-factor asset pricing models. The one-factor model is the CAPM with the return on a value-weighted stock market portfolio (MKT) as the risk factor. The MKT is included as the first factor in each of the two- to four-factor models. The two-factor model includes the EREIT return as a second risk factor. The three-factor model includes an investment factor and a profitability factor as the second and third factors. Lastly, the four-factor model enhances the three-factor model with the EREIT factor as the fourth factor. Here I include EREIT as a factor in the two- and four-factor models in order to facilitate the test of the integration hypothesis, following Liu et al. (1990) and Li (2016).

Under the assumption that the CAPM is correctly specified, the equity market integration hypothesis implies that a two-factor model should be rejected in favor of the CAPM. Alternatively, if one assumes that the expected EREIT return is determined by the three-factor model, the four-factor model should be rejected in favor of the three-factor model. I test the CAPM against the two-factor model or the three-factor model against the four-factor model: $\lambda_4 = 0$ The rejection of the two- or four-factor model is also a rejection of the market integration hypothesis.

An alternative way to test capital market integration is based on idea of the law of one price. Under the assumption that the expected EREIT return is determined by one of the four asset pricing models, the risk price(s) associated with the EREIT return should be equal to that(those) of other portfolio returns. Following Ling and Naranjo (1999), Guo et al. (2009) and Li (2016), I test restrictions on risk prices across portfolios.

To accomplish this, I replace the previous model with equal risk prices in equations (2)-(3) with a model in which the risk prices associated with the EREIT return are free from those associated with other portfolio returns. To this end, I use equation (2) and the following equation for the EREIT return:

$$R^e_{N,t} = a_{t\le T} I_{t\le T} + a_{t>T} I_{t>T} + \sum_{j=1}^{k} \lambda_{EREIT,j} h_{Nj,t} + \varepsilon_{N,t} \quad (5)$$

where the risk price, $\lambda_{EREIT,j}$, is the price of the j[th] risk factor associated with the EREIT. Given equation (5), another way of testing the market integration hypothesis is to test the restrictions:

$$\lambda_j = \lambda_{EREIT,j}, j = 1,2,,K, \quad (6)$$

where λ_j's are from equations (2) and $\lambda_{EREIT,j}$'s are from (5). The rejection of the restrictions in equation (6) is the rejection of the market integration hypothesis.

Using the unrestricted model given by equations (2) and (5), I test both the market integration and segmentation hypotheses. On the one hand, market integration requires that $\lambda_4 = \lambda_{EREIT,4} = 0$ in the two-factor or four-factor model so the two-factor model is reduced to the CAPM or the four-factor model is reduced to the three-factor model. On the other hand, market segmentation requires that expected returns on the first three factor portfolios are only related to their covariances with the aggregate stock market return in the CAPM or their covariances with the three factor portfolios in the three-factor model. Similarly, the expected EREIT return is only related to its own volatility. Together, market segmentation implies the restriction that $\lambda_4 = \lambda_{EREIT,1} = 0$ in the two-factor model, or the restrictions that $\lambda_4 = \lambda_{EREIT,1} = \lambda_{EREIT,2} = \lambda_{EREIT,3} 0$ in the four-factor model. If both the market integration and segmentation hypotheses are rejected or neither hypothesis is rejected, the market may be partially segmented.

C. Measuring Time-varying Integration

There is substantial research on the integration of international equity markets. There is also considerable variation in method of measuring the level of integration. For instance, Carrieri, Errunza, and Hogan (2007) use GARCH-in-mean to assess correlation in returns and volatility among markets, whereas Bekaert, Harvey, and Ng (2005) and Bekaert et al. (2011) use multiple economic fundamental factors. Integration is often interpreted based on cross-country correlations in stock returns. When multiple factors drive returns, Pukthuanthong-Le and Roll (2009) show that correlation may be a misleading measure of integration. While perfect integration implies that identical global factors fully explain index returns across countries, some countries may differ in their sensitivities to those factors and accordingly not exhibit perfect correlation. Cotter, Gabriel, and Roll (2015) extend this idea to U.S. housing markets and employ R-square from a regression of metropolitan housing returns on an identical set of national variables. The housing markets are regarded as highly integrated if identical U.S. national factors explain a large portion of the variance in metro-specific housing returns. They use the R-square from the multifactor model fit to a sequence of moving windows to obtain time-varying measure of integration.

There is substantial evidence that returns on a few factors explain the cross section of returns on the U.S. stock market. If the EREIT market is integrated with the general stock market, the time variation of the EREIT return should be explained by the same set of factors rather than the stock market portfolio alone. As a result, the correlation between the EREIT return and the stock market return may be an imperfect measure of the integration between the two markets. Instead, the R-square from a multifactor model should be a better measure of integration.

Unlike Cotter, Gabriel and Roll (2015), who employ regressions with moving windows, I use the dynamic GARCH-in-mean model employed in this paper to derive a time-varying measure of integration. I consider the following multifactor factor model for the disturbance terms in equation (3):

$$\varepsilon_{Nt} = \boldsymbol{\beta}'_{Nt}[\mathbf{F}_t - E_{t-1}(\mathbf{F}_t)] + e_{Nt} \tag{7}$$

where F_t is a $K \times 1$ vector of factors. Under the assumption that the idiosyncratic component e_{Nt} is conditionally uncorrelated with each of the factors, the $K \times 1$ vector of β_{Nt} for the N[th] asset is given by

$$\boldsymbol{\beta}'_{Nt} = \mathrm{cov}_{t-1}(\varepsilon_{Nt}, \mathbf{F}'_t)\mathrm{cov}_{t-1}^{-1}(\mathbf{F}_t, \mathbf{F}'_t) = \mathbf{h}'_{Nt}\mathbf{H}_{Kt}^{-1}, \tag{8}$$

where $\mathbf{h}_{Nk,t}$ is a vector of covariances between the EREIT return and the factors, and $\mathbf{H}_{Kt}$ as a submatrix of $\mathbf{H}_t$ containing the first K rows and columns, is the variance-covariance matrix of the K factors.

The R-square of the factor model in equation (7) for the period t is given by

$$R_t^2(K) = \frac{\boldsymbol{\beta}'_{Nt}\mathbf{H}_{Kt}\boldsymbol{\beta}_{Nt}}{\mathbf{h}_{NN,t}} \tag{9}$$

By substituting equation (8) into equation (9), one has the following

$$R_t^2(K) = \frac{\mathbf{h}'_{Nt}\mathbf{H}_{Kt}^{-1}\mathbf{h}_{Nt}}{\mathbf{h}_{NN,t}} \tag{10}$$

The integration R-square in equation (10) measures the level of integration of the EREIT market with the general stock market at time t. While this measure shows the contribution of all factors, the following shows the contribution from the market portfolio alone:

$$R_t^2(1) = \frac{\mathbf{h}_{N1,t}^2}{\mathbf{h}_{11,t}\mathbf{h}_{NN,t}} \tag{11}$$

The R-square in equation (11) is simply the squared correlation coefficient between the EREIT return and the return on the stock market portfolio. I also compute a similar R-square, $R_t^2(K)$ for k<K, to measure the contribution of first *k* factors.

For a given point in time, if the EREIT return is perfectly correlated with one of the factor portfolio returns or is a linear combination of the factor portfolio returns, the conditional variance of the idiosyncratic term vanishes and the value of $R_t^2(K)$ attains unity. Otherwise, the $R_t^2(K)$ measures the portion of the conditional volatility of the EREIT return that is explained by the K factors. At time t, the EREIT market is more integrated with the general stock market when $R_t^2(K)$ is higher.

IV. DATA AND METHODS

Given the limited availability of historical REITs data dating back to the 1970s, this study uses monthly returns on EREITs along with the monthly data of the factors. Monthly returns on the EREIT indices are obtained from the National Association of Real Estate Investment Trusts (NAREIT) for the sample period from January 1972 to July 2014. Approximately as of the end of 2014, stock exchange-listed Equity REITs account for 70 percent of all U.S. listed REIT assets, and EREITs represent 90 percent of the approximately $700 billion equity market capitalization of the listed REIT marketplace. There are approximately 150 listed EREITs, almost all of which are traded on the New York Stock Exchange. There are 26 listed residential Mortgage REITs with a market capitalization of $42.3 billion.

The first data set of factors includes returns on two factor portfolios in the *q*-factor model: the investment factor and the profitability factor. The monthly q-factor data for the period from January 1972 to December 2014 are provided by Hou et al. (2015), who construct the factors from a triple 2-by-3-by-3 sort on market equity, investment-to-assets, and returns on equity. When constructing the q-factors, Hou et al. (2015) form the investment portfolios annually but the ROE portfolios monthly. The investment (I/A) factor is the difference between the return on a low I/A portfolio and the return on a high I/A portfolio. The profitability (ROE) factor is the difference between the return on a high ROE portfolio and the return on a low ROE portfolio.

The second data set of factors includes the investment factor and profitability factor constructed by Fama and French (2016) and available in Kenneth French's website. Unlike the triple 2-by-3-by-3 sort used by Hou et al. (2015), the alternative factors are based on a double II-by-3 sort on size and investment (or profitability). Investment is measured by the growth of total assets for a fiscal year divided by total assets at the end of the year. The profitability is measured by annual revenues minus cost of goods sold, selling, general, administrative expenses and interest expense divided by book equity. CMA (conservative minus aggressive) is the return on the conservative investment portfolio minus the average return on the aggressive investment portfolio. RMW (robust minus weak) is the return on the robust operating profitability portfolios minus the return on a weak operating profitability portfolio.

The third date set includes the three factors used by Fama and French (1993): the market excess return (MKT), the size factor (small-minus-big, SMB), and the value factor (high-minus low, HML). SMB and HML are constructed from a double II-by-3 sort on size and book-to-market ratio. SMB is the monthly return on a portfolio of small stocks minus and the return on a portfolio of big stocks, and HML is monthly return on a portfolio of stocks with high book-to-market ratios (value stocks) minus and the return on a portfolio of stocks with low book-to-market ratios (growth stocks).

To understand the time variation of integration of EREIT with the general stock market, I employ a number of real economic and financial variables. The selection of the variables is motivated by the availability of monthly data and previous studies on real estate prices. The first is the civilian unemployment rate from the U.S. Bureau of Labor Statistics. This series is seasonally adjusted, representing the number of unemployed as a percentage of the labor force. This variable is used primarily to capture the weakness of the economy and the stage of the business cycle. Quan and Titman (1999) find that the correlation between the returns of stocks and real estate increases if business-cycle variables simultaneously affect corporate profits and rents. The second variable is the growth rate (difference in logs) of real estate loans from all commercial banks, from the Board of Governors of the U.S. Federal Reserve System. This variable helps to capture the change in the credit availability. Bernanke, Gertler, and Gilchrist (1996) point out that adverse shocks to the economy may be amplified by worsening credit-market conditions as resulting from endogenous changes over the business cycle in the agency costs of lending. The third variable is the default spread, computed as the difference between Moody’s Baa-rated corporate bond yield and Aaa-rated corporate bond yield. This variable is intended to capture the risk in the credit market. Fama and French (1989) find that the predictability of stock returns is related to long-term business cycles, as captured by the default spread. All three series are available from the St. Luis Fed’s FRED database. Other variables, such as the federal funds rate and the term spread (Moody's Aaa corporate bond yield minus yield on 10-year Treasury constant maturity bond) are not included as they are found to be statistically insignificant. Plazzi, Torous, and Valkanov (2008) document that macroeconomic variables such as the term and default spreads help explain cross-sectional dispersions of returns and growth in rents for commercial real estate.

V. EMPIRICAL RESULTS

A. Summary Statistics

In Table 1, I present the summary statistics for the EREIT index, along with those for the factors used in the paper. The average of EREIT return is 0.481 percent with a standard deviation of 4.928 percent over the sample period 1972-01 through 2014-12. The average is lower than those of MKT (0.536) and ROE (0.555), but higher than those of I/A (0.433), CMA (0.358), RMW (0.277). The standard deviation of the EREIT is slightly higher than that of MKT (4.572), but much higher than those of factors (1.859–2.992).

Interestingly, the EREIT return is negatively correlated with the profitability factors ROE (-0.224) and RMW (-0.132) and weakly correlated with the investment factors I/A (-0.028) and CMA (-0.045). However, the EREIT return is positively correlated with SMB (0.373) and HML (0.133). Among the investment and profitability factors, the I/A factor is very highly correlated the CMA factor with a correlation of 0.903 and the ROE factor is highly correlated with the RMW factor with a correlation of 0.674. However, the correlations between any investment factor and any profitability factor are low.

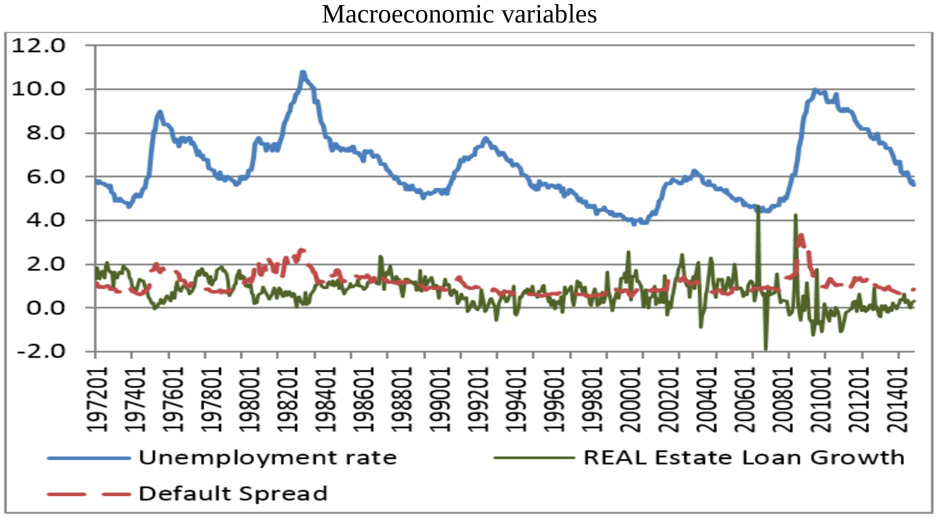
Among the macroeconomic variables, the real estate loan growth has relatively high standard deviation relative to its mean compared with two other variables. The unemployment rate is positively correlated with the default spread (0.583), but is negatively correlated with the real estate loan growth (-0.382), implying that a weak economy tends to be associated with a high credit risk and tightening credit market conditions. However, the correlation between the real estate loan growth and the default spread is moderate (-0.094), suggesting a negative but weak relation between the credit availability and the credit risk. Figure 1 illustrates the time series fluctuations of the three variables. While both the unemployment rate and the default spread appear counter-cyclical, the latter clearly leads the former for most and especially the latest boom-and-bust cycle of the economy and real estate prices. As a result, the inclusion of both variables should help capture different stages of the cycles. The real estate loan growth exhibits a pattern of procyclical movements, in contrast more to the counter-cyclical variation of the unemployment rate than to that of the default spread.

Table 1
Summary statistics

	EREIT	MKT	I/A	ROE	CMA	RMW	UNEM	RELG	DEF
Mean %	0.481	0.536	0.433	0.555	0.358	0.277	6.447	0.738	1.100
S.D. %	4.928	4.572	1.859	2.588	1.981	2.240	1.561	0.674	0.465
Correlations									
MKT	0.592								
I/A	-0.028	-0.366							
ROE	-0.224	-0.188	0.063						
CMA	-0.045	-0.394	0.903	-0.062					
RMW	-0.132	-0.225	0.099	0.674	-0.013				
UNEM	0.130	0.110	0.032	-0.080	-0.001	-0.033			
RELG	-0.132	-0.073	-0.082	0.042	-0.050	-0.075	-0.382		
DEF	0.031	0.053	0.008	-0.089	-0.010	-0.020	0.583	-0.094	

Notes: MKT is the market excess return. I/A and ROE are investment factor and profitability factor, respectively, constructed from a triple 2-by-3-by-3 sort, in the *q*-factor model. CMA and RMW are alternative investment factor and profitability factor, respectively, constructed from double 2-by-3 sorts in the Fama-French model. UNEM is the lagged civilian unemployment rate. RELG is the lagged growth rate (difference in logs) of real estate loans from all commercial banks. DEF is the lagged default spread, calculated as the difference between Moody's Baa-rated corporate bond yield and Moody's Aaa-rated corporate bond yield.

Figure 1
Macroeconomic variables



The civilian unemployment rate is the number of unemployed as a percentage of the labor force. The real estate loan growth is the growth rate (difference in logs) of real estate loans from all commercial banks. The default spread is the difference between Moody's Baa-rated corporate bond yield and Aaa-rated corporate bond yield.

B. The Role of I/A and ROE Factors

I now report the results of estimating the models in equations (2)-(4) in Table 2 (including I/A and ROE factors). In panel A, I present the results of estimating the CAPM (one-factor) and the two-factor model. In panel B, I report the results of estimating a three-factor model and a four-factor model. The size factor is excluded since estimating a five-factor model is computationally difficult. To conserve space, I do not report the estimates of the constant matrix **C** throughout the paper. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic.

Table 2
Estimates of models with equal risk prices including I/A and ROE factors

	MKT	I/A	ROE	EREIT	MKT	I/A	ROE	EREIT
Panel A. One and two-factor models								
$a_{t<93}$,%				-0.507				**-1.204**
				(0.200)				(0.239)
$a_{t\geq 93}$,%				-0.234				**-1.100**
				(0.235)				(0.178)
λ	**2.563**				-0.930			**5.871**
	(0.762)				(0.842)			(0.708)
Diag. **A**	**0.926**	**0.900**	**0.876**	**0.897**	**0.933**	**0.894**	**0.881**	**0.907**
	(0.012)	(0.014)	(0.018)	(0.021)	(0.003)	(0.006)	(0.005)	(0.003)
Diag. **B**	**0.233**	**0.313**	**0.412**	**0.283**	**0.232**	**0.312**	**0.406**	**0.286**
	(0.021)	(0.027)	(0.052)	(0.034)	(0.012)	(0.020)	(0.013)	(0.010)
Diag. **D**	**0.260**	**-0.220**	0.049	0.268	**0.231**	**-0.240**	0.038	**0.176**
	(0.059)	(0.062)	(0.055)	(0.123)	(0.017)	(0.062)	(0.096)	(0.021)
Panel B. Three- and four-factor models								
$a_{t<93}$,%				-0.411				-0.697
				(0.260)				(0.400)
$a_{t\geq 93}$,%				-0.185				-0.548
				(0.230)				(0.510)
λ	**4.777**	**15.559**	**6.972**		3.106	**14.413**	**7.606**	2.567
	(1.166)	(3.208)	(2.609)		(1.888)	(2.784)	(2.381)	(2.502)
Diag. **A**	**0.931**	**0.921**	**0.885**	**0.904**	**0.933**	**0.918**	**0.885**	**0.907**
	(0.010)	(0.019)	(0.025)	(0.024)	(0.011)	(0.011)	(0.020)	(0.020)
Diag. **B**	**0.250**	**0.257**	**0.358**	**0.316**	**0.248**	**0.256**	**0.389**	**0.316**
	(0.043)	(0.041)	(0.078)	(0.059)	(0.033)	(0.023)	(0.067)	(0.035)
Diag. **D**	**0.230**	**-0.167**	0.042	0.151	**0.221**	**-0.177**	0.037	0.127
	(0.085)	(0.038)	(0.042)	(0.253)	(0.058)	(0.037)	(0.046)	(0.146)

Panel C. Likelihood ratio tests

Integration hypothesis, $\lambda_4 = 0$	x^2	d.f.	p-value
One- vs. two-factor	6.493	1	0.011
Three- vs. four-factor	1.665	1	0.197

Notes: The return on each of the first three portfolios is given by equation (2): $R^e_{it} = \sum_{j=1}^{k} \lambda_j h_{ij,t} + \varepsilon_{it}$. The excess return on the EREIT portfolio is given by equation (3): $R^e_{N,t} = a_{t\leq T} I_{t\leq T} + a_{t>T} I_{t>T} + \sum_{j=1}^{k} \lambda_j h_{Nj,t} + \varepsilon_{N,t}$, where $a_{t\leq T}$ is the intercept for the period before date *T* and $a_{t>T}$ is the intercept for the period after date *T*. The conditional variance-covariance matrix $\mathbf{H_t}$ in follows the asymmetric BEKK GARCH specification: $\mathbf{H_t = C'C + A'H_{t-1}A + B'\varepsilon_{t-1}\varepsilon_{t-1}'B + D'\eta_{t-1}\eta_{t-1}'D}$. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic. Robust standard errors are reported in the parentheses.

The left side of panel A presents the results of estimating the CAPM. The estimated $a_{t<93}$ for EREIT is -0.507 with a standard error of 0.200, implying that the alpha in the pre-1993 period is more than two standard errors away from zero and significant at the 5 percent level. Thus, the EREIT return exhibits a significant alpha in the first half of the sample period before 1993. The estimated intercept for the post-1993 sample, $a_{t\geq 93}$, is -0.234 with a standard error of 0.235. The price of the market risk (λ) is also significant with an estimate of 2.563 and a standard error of 0.762.

In the right side of Panel A, I report the results of estimating the two-factor model in which the expected excess return on each portfolio is related to its covariance with MKT and EREIT. The estimates of $a_{t<93}$ and a_{93} are -1.204 and -1.100, respectively, with standard errors of 0.239 and 0.178, implying that the alphas are significant at the 1 percent level in both the pre- and post-1993 period. The price of market risk λ becomes insignificant, with a coefficient estimate of -0.930 and a standard error 0.842. However, the risk price associated with EREIT is significant with a coefficient of 5.871 and a standard error of 0.708. This suggests that the expected excess return on EREIT is related to its own volatility, but is unrelated to its covariance with MKT. This is against the market integration hypothesis and the CAPM. The expected returns on MKT, I/A and ROE factors are related to their covariances with the EREIT return in this setting, suggesting that the EREIT return is more important than the MKT return as a risk factor in the absence of the I/A and ROE factors. The result of estimating the two-factor model therefore suggests that the EREIT return is not explained by the CAPM or the MKT risk alone. The result here is consistent with the finding of Li (2016), who uses a similar two-factor model.

In the left side of Panel B of Table 2, I report estimates in the three-factor model. Only alpha for the pre-1993 period, with a point estimate of -0.411 and a standard error of 0.260, is insignificant at the 10 percent level. The magnitude of the alpha in each of the two periods in the three-factor model is also lower than that in the CAPM. This suggesting that the three-factor model explains the EREIT return better than the one- and two-factor models. The estimates of the risk price parameters (λ) are 4.777 for MKT (std. err. = 1.166), 15.559 for the I/A factor (std. err. = 3.208) and 6.972 (std. err. = 2.609) for the ROE factor, implying that the risk prices associated with MKT and I/A and ROE factors are all positive and significant at the 1 percent level. The results suggest that both investment and profitability factors are important.

Next, the right side of panel B of Table 2 displays the results of estimating the four-factor model, which includes EREIT as the fourth factor, in addition to the factors in the three-factor model. Like that in the three-factor model, only the alpha for the pre-1993 period, with a point estimate of -0.697 and a standard error of 0.400, is significant at the 10 percent level. The alpha in each of the two periods is approximately half of that in the two-factor model. In the presence of the EREIT factor, the prices of the I/A and ROE risk factors are similar to the risk prices estimated in the three-factor model and still significant at the 1 percent level, while the price of the MKT risk is significant at the 10 percent level. However, the price for the EREIT risk is not statistically significant even at the 10 percent level, indicating that the risk-return tradeoff on the EREIT portfolio is explained by the systematic risks associated with the market factor, the investment and profitability factor. In other words, the result suggests that the real estate premium in the two-factor model disappears when I account for premiums associated with the investment factor and profitability factor. The estimated risk prices in the four-factor are consistent with the integration hypothesis that the EREIT return is unrelated to its volatility.

Finally, I discuss the estimates of GARCH parameters. All of the estimates of diagonal elements of matrices **A** and **B** plus the first two elements of the diagonal elements of matrix **D** are significant at the 1 percent level. Since the estimates in the four models are similar with few exceptions, I mostly discuss the estimates in panel A. The estimated diagonal elements of matrix **A** (which link second moments to their lagged values) are 0.926 (MKT), 0.900 (I/A), 0.876 (ROE), and 0.897 (EREIT), implying high persistence of the volatility of each factor and the REIT return. The estimated diagonal parameters of matrix **B** (which link second moments to past innovations) are 0.233, 0.313, 0.412 and 0.283, indicating sizable effects of past innovations on the second moments. More interestingly, the diagonal parameters of matrix **D** (which measure the asymmetric effects of negative shocks) are 0.260 (MKT), -0.220 (I/A) and 0.268 (EREIT), which are significant at the 5 or 1 percent level. The sizes of the estimated parameters of **B** and **D** suggest that ignoring the asymmetric effects of negative shocks would greatly underestimate the impacts of negative news on the volatility of MKT, I/A and EREIT portfolios. The result is consistent with asymmetric betas (Chiang, Lee, and Wisen, 2004) and the asymmetric response of REIT returns to monetary policy (Chou and Chen, 2014).

In panel C, I report the results of the likelihood ratio tests of the market integration hypothesis. The test of the CAPM against the two-factor model produces a x^2 statistic of 6.493. With one degree of freedom, the CAPM and the integration hypothesis are rejected at the 5 percent level. However, the test of the three-factor model against the four-factor model produces a x^2 statistic of 1.665, with one-degree of freedom and a *p*-value of 0.197. Hence, the three-factor model and the market integration hypothesis are not rejected even at the 10 percent level. Finally, I test the CAPM against the three-factor model as a model specification test. The resulting x^2 statistic is 50.998, with two-degree of freedom and a *p*-value of 0.000. Therefore, the CAPM is rejected against the three-factor model at any conventional significance level. The results suggest that testing the market integration should be based on the three-factor model rather than the CAPM.

In Table 3, I report the results of estimating and testing the model with unequal risk prices, given by equations (2) and (5). To save space, I omit the estimates of the GARCH parameters, which are very similar to earlier estimates in the model with equal risk prices. In the CAPM, the alpha in the pre-1993 period is -0.294 with a standard error of 0.220 and the alpha in the post-1993 period is -0.033 with a standard error of 0.335. Thus with unequal risk prices, the alphas are no longer significant at the 10 percent level and the CAPM is no longer rejected. The estimated market price of risk, λ, associated with the first three portfolios is 2.707, with a standard error of only 0.859. Thus, the estimated λ is significant at the 1 percent level. The estimated market price of risk, λ_{EREIT}, associated with the EREIT is 1.010, which is less than half of the estimated λ. However, the standard error of λ_{EREIT} is 1.638, which makes the risk price estimate very imprecise and insignificant at the 10 percent level.

In the two-factor model, the alpha in the pre-1993 period is -1.248 with a standard error of 0.527 and the alpha in the post-1993 period is -1.005 with a standard error of 0.346. Thus with unequal risk prices, the alphas are significant at the 5 or 1percent level

Table 3
Estimates of models with unequal risk prices including I/A and ROE factors

	MKT	I/A	ROE	EREIT	MKT	I/A	ROE	EREIT
				Panel A. One and two-factor models				
$a_{t<93}$,%				-0.294				<u>-1.248</u>
				(0.220)				(0.527)
$a_{t\geq 93}$,%				-0.033				**-1.005**
				(0.335)				(0.346)
λ	**2.707**				-3.137			**9.587**
	(0.859)				(1.942)			(3.689)
λ_{EREIT}	1.010				-1.896			**6.397**
	(1.638)				(4.899)			(1.537)
				Panel B. Three- and four-factor models				
$a_{t<93}$,%				-0.081				-0.423
				(0.224)				(0.454)
$a_{t\geq 93}$,%				0.049				-0.407
				(0.385)				(0.498)
λ	**4.831**	**15.615**	**6.734**		1.747	**13.618**	**7.960**	4.788
	(0.886)	(2.455)	(1.486)		(3.221)	(3.048)	(2.435)	(5.269)
λ_{EREIT}	1.838	18.812	2.788		-1.974	15.063	3.244	3.483
	(2.597)	(10.611)	(3.784)		(3.852)	(12.571)	(4.638)	(2.698)

Panel C. Likelihood ratio tests

Law of one price, $\lambda_j = \lambda_{EREIT,j}$	x^2	d.f.	p-value
CAPM, $j = 1$	1.429	1	0.232
Two-factor, $j = 1,4$	5.512	2	0.064
Three-factor, $j = 1,2,3$	2.384	3	0.497
Four-factor, $j = 1,2,3,4$	3.611	4	0.461
Integration hypothesis, $\lambda_4 = \lambda_{EREIT,4} = 0$			
CAPM vs. two-factor,	10.575	2	0.005
Three- vs. four-factor	2.891	2	0.236
Segmentation hypothesis, $\lambda_4 = \lambda_{EREIT,j} = 0$			
Two-factor model, $j = 1$	9.678	2	0.008
Four-factor model, $j = 1,2,3$	5.733	4	0.220

Notes: The return on each of the first three portfolios is given by equation (2): $R^e_{it} = \sum_{j=1}^{k} \lambda_j h_{ij,t} + \varepsilon_{it}$ The excess return on the EREIT portfolio is given by equation (5): $R^e_{N,t} = a_{t\leq T} I_{t\leq T} + a_{t>T} I_{t>T} + \sum_{j=1}^{k} \lambda_{EREIT,j} h_{Nj,t} + \varepsilon_{N,t}$, where $a_{t\leq T}$ is the intercept for the period before date *T* and $a_{t<T}$ is the intercept for the period after date T. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic. Robust standard errors are reported in the parentheses.

and the two-factor model is rejected. The estimated market prices of risk, λ and λ_{EREIT}, are -3.137 (std. err. = 1.942) and -1.896 (std. err. = 4.899), respectively, implying that both estimates of risk prices are insignificant. In contrast, the estimated prices of EREIT risk, λ and λ_{EREIT}, are 9.587 (std. err. = 3.689) and 6.397 (std. err. = 1.537), respectively, implying that the risk price estimates are significant at the 1 percent level. The significance of alphas and risk prices in the model with unequal risk prices echoes that in the previous model with equal risk prices.

In the three-factor model with unequal risk prices, the pre-1993 alpha (-0.081; std. err. = 0.224) and post-1993 alpha (0.049; std. err. = 0.385) are smaller in magnitude, compared with those in the model with equal risk prices, and insignificant at the 10 percent level. In the four-factor model, the alphas are also smaller in magnitude and insignificant. Thus, the models here perform better than those with equal risk prices. The risk prices associated with the first three portfolios, λ, in the three- and four-factor models here are similar in magnitude and significance to those in the previous models with equal risk prices. The estimated risk prices λ in the three-factor model are 4.831, 15.615, and 6.734; while the estimated λ_{EREIT} are 1.838, 18.812 and 2.788. In the four-factor model, the estimates of λ are 1.747, 13.618, 7.960 and 4.788; while the estimated λ_{EREIT} are -1.974, 15.063, 3.244 and 3.483. However, most estimates of the risk prices associated with the EREIT are insignificant at the 10 percent level. The only exception is the price of the I/A risk in the three-factor model, which is 18.812 with a standard error of 10.611, so the estimate is significant at the 10 percent level.

Finally, I discuss the results of testing the market integration and segmentation. In the CAPM, the x^2 statistic is 1.429, with one degree of freedom and a *p*-value of 0.232. In the two-factor model, the x^2 statistic is 5.512, with two degrees of freedom and a p-value of 0.064. In the three-factor model, the x^2 statistic is 2.384, with three degrees of freedom and a *p*-value of 0.497. In the four-factor model, the x^2 statistic is 3.611, with four degrees of freedom and a *p*-value of 0.461. Hence, based on the law of one price, the market integration hypothesis is not rejected in the CAPM, the three- or four-factor models but rejected in the two-factor model at the 10 percent level.

The results in panel C of the table offer further results on market integration and segmentation from tests of alternative models. Similar to the results in Table 2, the market integration hypothesis is rejected in the test of the CAPM against the two-factor model with a p-value of 0.005, and the CAPM is also rejected in favor of the three-factor model with a p-value of 0.000. But the integration hypothesis is no rejected in the test of the three-factor model against the four-factor model, with a p-value of 0.236.

The tests of the market segmentation hypothesis reveal quite different facts. The market segmentation hypothesis is rejected based on the two-factor model, with a p-value of 0.008, but the hypothesis is not rejected based on the four-factor model, with a p-value of 0.220. In summary, based on the two-factor model, both the market integration and segmentation hypotheses are rejected. However, based on the four-factor model, neither the integration nor the segmentation hypothesis is rejected. Therefore, the results are consistent with the partial segmentation hypothesis.

In panel A of Table 4, I present descriptive statistics for the integration R-squares in equations (10)-(11), computed from estimates in Table 2 for the three-factor model with MKT, I/A and ROE as the factors. I choose to do so because this model is not rejected in previous tests. In Figure 2, I illustrate the time series of the three-factor and market R-squares. The average three-factor R-square is 40.2%, with a standard error of 15.6%. There is considerable variation of the R-square, with a minimum of 3.4% and a maximum of 79.4%. The average two-factor R-square (MKT, I/A) is 36.9% with a standard error of 14.8% and the average market R-square is 34.3% with a standard error of 15.2%. The results suggest that the I/A factor and the ROE factor contribute roughly equally to the overall three-factor R-square. The contributions from other factors tend to more evident around peaks and troughs of the cycles.

Table 4
Descriptive statistics and regressions of implied R-squares for I/A and ROE factors

Panel A. Descriptive Statistics

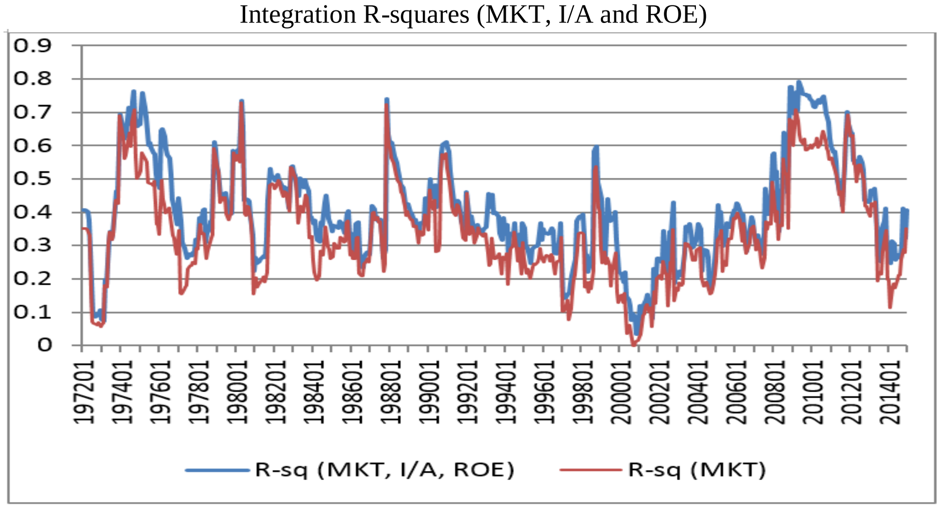
Variable	Mean	Std. Err.	Min.	Max.
R-Sq. (MKT,I/A,ROE)	0.402	0.156	0.034	0.794
R-Sq. (MKT,IA)	0.369	0.148	0.026	0.737
R-Sq. (MKT)	0.343	0.152	0.000	0.728

Panel B. Regressions of Implied R-Squares

Dependent variable	Constant	Unemployment Rate	Real Estate Loan Growth	Default Spread	Adj. R^2
R-Sq. (MKT,I/A,ROE)	0.078	**0.050**			0.250
	(0.091)	(0.015)			
	0.448		**-0.063**		0.071
	(0.045)		(0.017)		
	0.244			**0.143**	0.180
	(0.056)			(0.051)	
	0.132	<u>0.032</u>	<u>-0.029</u>	0.077	0.288
	(0.125)	(0.015)	(0.015)	(0.044)	
R-Sq. (MKT)	0.046	**0.046**			0.223
	(0.102)	(0.015)			
	0.382		**-0.053**		0.054
	(0.043)		(0.014)		
	0.194			**0.136**	0.172
	(0.057)			(0.044)	
	0.088	*0.029*	<u>-0.022</u>	<u>0.076</u>	0.258
	(0.130)	(0.016)	(0.011)	(0.035)	

Notes: The R-sq's are implied by the three-factor model in Table 2. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic. All explanatory variables are one-month lagged variables. Standard errors in parentheses are robust to heteroskedasticity and residual serial autocorrelations up to 60 lags.

Figure 2
Integration R-squares (MKT, I/A and ROE)



As shown in Figure 2, the integration of the EREIT market with the general stock market increases during the boom period and reaches the peak around the recent financial crisis. The integration R-square is about 3-4% in late 2000 but about 50-58% in early 2008. As boom turns to bust, integration continues to rise, reaching about 75-79% in mid-2009. Subsequently, as the financial crisis subsides, market integration starts trending down to 25% in early 2014. The market R-square generally follows the same pattern as the three-factor R-square throughout the full sample period. The time series property of the integration R-squares supplements the findings that the expected return and volatility of EREIT surge to unprecedented level during the recent financial crisis (Sun, Titman and Twite, 2015; Li, 2015). Compared with the integration R-squares in the housing markets (Cotter, Gabriel, and Roll, 2015), I observe more frequent and larger swings in the level of integration in the EREIT market than the housing markets. For instance, I uncover 3-4 cyclical movements of the integration R-square in the EREIT market in the 1972-2001 period (before the latest 2001-2010 cycle), while the national housing market integration is relatively stable in the 1992-2001 period. In addition, the integration varies between 5-10 and 80 percent in the EREIT market but between 35 and 65 percent, approximately, in housing markets. I also note that the level of integration in the US market during and immediately after the financial crisis (2008-2012) is higher than the long-run average level of approximately 40 percent. This result is consistent with what Chang, Chou and Fung (2012) find that, after the financial crisis, REITS returns show a stronger linkage to the overall market returns for Australia, Japan, Taiwan and the USA.

To understand the economic causes of the time variation of market integration, I report in panel B of Table 4 the results of regressions of the R-squares on three macroeconomic variables. All explanatory variables are one-month lagged variables. Standard errors in parentheses are robust to heteroscedasticity and residual serial autocorrelations up to 60 lags. In six univariate regressions, the point estimate of the slope coefficient of each of the three variables (unemployment rate, real estate loan growth and default spread) is significant at the 1 percent level. With the three-factor R-square as the dependent variable, the coefficient estimates are 0.050 (std. error = 0.015), -0.063 (std. error = 0.017) and 0.143 (std. error = 0.051), respectively, implying that integration is positively related to the unemployment rate and the default spread but negatively related to real estate loan growth.

The adjusted R^2 of the regressions are 0.250, 0.071 and 0.180, respectively. The results suggest that the level of integration is counter-cyclical and tends to be highest at the peak of the recessions. In a multiple regression, the first two variables (unemployment rate and real estate loan growth) enter the regression with coefficients significant at the 5 percent level, while the third variable (default spread) enters with a coefficient significant at the 10 percent level. The adjusted R^2 is 0.288, implying that close to 29 percent of the variation of the three-factor integration R-square is explained by three variables. The results for the market R-square are qualitatively similar. Approximately 26 percent of the variation of the market R-square is explained by the three factors. The noticeable difference is that second and third variables (real estate loan growth and default spread) are significant at the 5 percent level, while the first (unemployment rate) is significant the 10 percent level. Thus the unemployment rate is more important for explaining the contributions of I/A and ROE factors while the default spread is more important for explaining the contribution of the MKT factor.

C. The Role of CMA and RMW Factors

In panel A of Table 5, I present the results of estimating the one- to four-factor models in which the I/A and ROE factors are replaced with CMA and RMW factors. First, I find that the results of estimating the one- and two-factor models are remarkably similar to those reported in Table 2. The estimated a_{93} for EREIT is -0.612 with a standard error of 0.182, implying that the alpha in the pre-1993 period is more than two standard errors away from zero and significant at the 1 percent level. Thus, the EREIT return exhibits a significant alpha in the first half of the sample period before 1993. In the two-factor model, The estimates of $a_{t<93}$ and a_{93} are -0.924 (std. error = 0.207) and -0.661 (std. error = 0.175), respectively, implying that the alphas are significant at the 1 percent level in both the pre- and post-1993 period.

The price of the market risk (λ) in the CAPM is also significant at the 1 percent level with an estimate of 2.503 and a standard error of 0.829. In the two-factor model, the price of market risk λ is insignificant, with an estimate of 0.667 and a standard error 0.885; but the price of EREIT risk is significant at the 1 percent level, with an estimate of 2.634 and a standard error of 0.745. The estimates of the GARCH parameters here are also mostly similar to those in Table 2. The only exception is that the diagonal element of D for negative shocks associated with the profitability factor (RMW) rather than the investment factor (CMA) is significant at the 1 percent level, unlike the result in Table 2. Although the factors in the one- or two-factor models used in Tables 2 and 4 are the same, two factor portfolios used in estimating the models in the two tables are different. This implies that the results from the CAPM and the two-factor model are not sensitive to the portfolios used in estimating the models.

The results of estimating three- and four-factor models in panel B of Table 5, however, are not quite similar to those in Table 2. The estimated pre-1993 alpha, $a_{t<93}$, is -0.609 with a standard error of 0.171 in the three-factor model, or -0.606 with a standard error of 0.286 in the four-factor model. Hence, the alpha is significant at the 1 percent level in either model, just like that in the CAPM or the two-factor model. This suggests that the CMA and RMW factors do not capture the average EREIT return as well as the I/A and ROE factors. This is true in spite of the fact that the prices of risks associated with the first three factors are all significant at the 1 percent level in the three- and four-factor models, but the price of the EREIT risk is not significant at the 10 percent level in the four-factor model.

The results of testing the integration hypothesis in panel C show that the CAPM is not rejected against the two-factor model with a p-value of 0.110. Thus the likelihood ratio test seems to have low power to detect the statistical significance of λ_4. Thus, the three-factor model is not rejected against the four-factor model, with a p-value of 0.952, in support of the full integration hypothesis, just like in the previous case with I/A and ROE as the second and third factors.

Table 5
Estimates of models with equal risk prices including CMA and RMW factors

	MKT	CMA	RMW	REIT	MKT	CMA	RMW	EREIT
				Panel A. One and two-factor models				
$a_{t<93}$,%				**-0.612**				**-0.924**
				(0.182)				(0.207)
$a_{t\geq 93}$,%				-0.297				**-0.661**
				(0.223)				(0.175)
λ	**2.503**				0.667			**2.634**
	(0.829)				(0.885)			(0.745)
Diag. **A**	**0.913**	**0.841**	**0.842**	**0.884**	**0.915**	**0.842**	**0.843**	**0.884**
	(0.017)	(0.015)	(0.016)	(0.019)	(0.003)	(0.009)	(0.008)	(0.003)
Diag. **B**	**0.253**	**0.432**	**0.417**	**0.183**	**0.251**	**0.434**	**0.415**	**0.178**
	(0.023)	(0.020)	(0.027)	(0.030)	(0.011)	(0.021)	(0.020)	(0.018)
Diag. **D**	**0.289**	0.068	**-0.296**	**0.416**	**0.281**	0.051	**-0.302**	**0.388**
	(0.059)	(0.123)	(0.057)	(0.051)	(0.019)	(0.059)	(0.039)	(0.016)
				Panel B. Three- and four-factor models				
$a_{t<93}$,%				**-0.609**				**-0.606**
				(0.171)				(0.286)
$a_{t\geq 93}$,%				-0.268				-0.257
				(0.181)				(0.395)
λ	**4.682**	**12.158**	**8.081**		**4.737**	**12.198**	**8.072**	-0.089
	(0.776)	(2.343)	(1.518)		(1.684)	(2.666)	(1.779)	(1.884)
Diag. **A**	**0.915**	**0.871**	**0.838**	**0.884**	**0.915**	**0.872**	**0.838**	**0.884**
	(0.017)	(0.038)	(0.028)	(0.020)	(0.017)	(0.037)	(0.027)	(0.020)
Diag. **B**	**0.256**	**0.382**	**0.410**	**0.183**	**0.256**	**0.381**	**0.410**	**0.183**
	(0.026)	(0.049)	(0.035)	(0.041)	(0.023)	(0.047)	(0.034)	(0.029)
Diag. **D**	**0.273**	0.004	**-0.266**	0.400	**0.273**	0.002	**-0.266**	**0.401**
	(0.045)	(0.108)	(0.064)	(0.035)	(0.054)	(0.071)	(0.069)	(0.046)
				Panel C. Likelihood ratio tests				
Integration hypothesis, $\lambda_4 = 0$						x^2	d.f.	p-value
One- vs. two-factor						2.551	1	0.110
Three- vs. four-factor						0.004	1	0.952

Notes: The return on each of the first three portfolios is given by equation (2): $R_{it}^{e} = \sum_{j=1}^{k} \lambda_j h_{ij,t} + \varepsilon_{it}$. The excess return on the EREIT portfolio is given by equation (3): $R_{i}^{e} = a_{t\leq T} I_{t\leq T} + a_{t>T} I_{t>T} + \sum_{j=1}^{k} \lambda_j h_{Nj,t} + \varepsilon_{N,t}$, where $a_{t\leq T}$ is the intercept for the period before date T and $a_{t>T}$ is the intercept for the period after date T. The conditional variance-covariance matrix $\mathbf{H_t}$ in follows the asymmetric BEKK GARCH specification: $\mathbf{H_t} = \mathbf{C'C} + \mathbf{A'H_{t-1}A} + \mathbf{B'\varepsilon_{t-1}\varepsilon_{t-1}'B} + \mathbf{D'\eta_{t-1}\eta_{t-1}'D}$. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic. Robust standard errors are reported in the parentheses.

In Table 6, I report the results of estimating and testing the model with unequal risk prices, with CMA and RMW as the second and third portfolios. In the CAPM, the alpha in the pre-1993 period is -0.417 with a standard error of 0.237 and the alpha in the post-1993 period is -0.132 with a standard error of 0.312. Thus, with unequal risk prices, only the alpha in the pre-1993 period is significant at the 10 percent level. The estimated market price of risk, λ, associated with the first three portfolios is 2.572, with a standard error of only 0.903. Thus the estimated λ is significant at the 1 percent level. The estimated market price of risk, λ_{EREIT}, associated with the EREIT is 0.924, with a standard error of 1.979, implying that the risk price estimate is insignificant at the 10 percent level.

Table 6

Estimates of models with unequal risk prices including CMA and RMW factors

	MKT	CMA	RMW	EREIT	MKT	CMA	RMW	EREIT
				Panel A. One and two-factor models				
$a_{t<93}$,%				-0.417			-0.653	
				(0.237)			(0.387)	
$a_{t\geq 93}$,%				-0.132			-0.522	
				(0.312)			(0.298)	
λ	**2.572**				-0.742		5.028	
	(0.903)				(2.507)		(3.744)	
λ_{EREIT}	0.924				-6.263		5.644	
	(1.979)				(4.523)		(2.981)	
				Panel B. Three- and four-factor models				
$a_{t<93}$,%				-0.433				-0.384
				(0.244)				(0.226)
$a_{t\geq 93}$,%				0.289				-0.296
				(0.381)				(0.360)
λ	**4.849**	**12.233**	**8.224**		**4.620**	**11.992**	**8.185**	0.238
	(0.768)	(2.415)	(1.697)		(1.496)	(2.192)	(1.883)	(1.659)
λ_{EREIT}	1.286	11.752	-17.046		-1.974	8.651	-10.721	2.105
	(1.914)	(13.493)	(24.311)		(2.316)	(13.291)	(18.006)	(1.018)

Panel C. Likelihood ratio tests

Law of one price, $\lambda_j = \lambda_{\mathrm{EREIT},j}$	x^2	d.f.	p-value
CAPM, $j = 1$	1.357	1	0.244
Two-factor, $j = 1,4$	5.553	2	0.062
Three-factor, $j = 1,2,3$	3.036	3	0.386
Four-factor, $j = 1,2,3,4$	4.185	4	0.382
Integration hypothesis, $\lambda_4 = \lambda_{\mathrm{EREIT},4} = 0$			
CAPM vs. two-factor	6.747	2	0.034
Three- vs. four-factor	1.153	2	0.562
Segmentation hypothesis, $\lambda_4 = \lambda_{\mathrm{EREIT},j} = 0$			
Two-factor model, $j = 1$	5.685	2	0.058
Four-factor model, $j = 1,2,3$	1.047	4	0.903

Notes: The return on each of the first three portfolios is given by equation (2): $R^e_{it} = \sum_{j=1}^{k} \lambda_j h_{ij,t} + \varepsilon_{it}$. The excess return on the EREIT portfolio is given by equation (5): $R^e_{N,t} = a_{t\leq T} I_{t\leq T} + a_{t>T} I_{t>T} + \sum_{j=1}^{k} \lambda_{\mathrm{EREIT},j} h_{Nj,t} + \varepsilon_{N,t}$, where $a_{t\leq T}$ is the intercept for the period before date *T* and $a_{t>T}$ is the intercept for the period after date *T*. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic. Robust standard errors are reported in the parentheses.

In the two-factor model, the alpha in the pre-1993 period is -653 with a standard error of 0.387 and the alpha in the post-1993 period is -0.522 with a standard error of 0.298. Thus with unequal risk prices, the alphas in the two-factor model are only significant at the 10 percent level in the posr-1993 period. The estimated market prices of risk, λ and λ_{EREIT}, are -0.742 (std. error = 2.507) and -6.263 (std. error = 4.523), respectively, implying that both estimates of risk prices are insignificant. In contrast, the estimated prices of EREIT risk, λ and λ_{EREIT}, are 5.028 (std. error = 3.744) and 5.644 (std. error = 2.981), respectively, implying that the risk price λ_{EREIT} is significant at the 10 percent level. The alphas and risk prices in the model with unequal risk prices here are less significant than those in the previous model with equal risk prices.

In the three-factor model with unequal risk prices, the pre-1993 alpha (-0.433; std. error = 0.244) and post-1993 alpha (0.289; std. error = 0.381) are less significant or insignificant, compared with those in the model with equal risk prices. In the four-factor model, the alphas are also insignificant. Thus, as the results in Table 3, the three- and four-factor models here perform better than those with equal risk prices. In the three-factor model, the estimated risk prices λ are 4.849, 12.233, and 8.224; while the estimated λ_{EREIT} are 1.286, 11.752 and -17.046. In the four-factor model, the estimates of λ are 4.620, 11.992, 8.185 and 0.238; while the estimated λ_{EREIT} are -1.974, 8.651, -10.721 and 2.105. Although the estimated risk prices λ associated with the first three portfolios are significant at the 1 percent level, most the estimated risk prices associated with the EREIT are significant at the 10 percent level. The only exception is that the estimated λ_{EREIT} associated with the EREIT portfolio is significant at the 5 percent level, offering evidence on the segmentation of the REIT market from the general stock market.

Finally, I discuss the results of testing the market integration and segmentation. In the CAPM, the x^2 statistic is 1.357, with one degree of freedom and a p-value of 0.244. In the two-factor model, the x^2 statistic is 5.553, with two degrees of freedom and a p-value of 0.062. In the three-factor model, the x^2 statistic is 3.036, with three degrees of freedom and a p-value of 0.386. In the four-factor model, the x^2 statistic is 4.185, with four degrees of freedom and a p-value of 0.382. Hence, based on the law of one price, the market integration hypothesis is not rejected in the CAPM, the three- or four-factor models but rejected in the two-factor model at the 10 percent level. All of the results here are qualitatively the same as those in Table 3.

Similar to the results in Table 3, the market integration hypothesis is rejected in the test of the CAPM against the two-factor model with a p-value of 0.034. However, the integration hypothesis is not rejected in the test of the three-factor model against the four-factor model, with a p-value of 0.562. The market segmentation hypothesis is rejected at the 10 percent level, based on the two-factor model, with a p-value of 0.058, but the hypothesis is not rejected based on the four-factor model, with a p-value of 0.903. As before, based on the two-factor model, there is some evidence against both the market integration and segmentation hypotheses. However, based on the four-factor model, neither the integration nor the segmentation hypothesis is rejected. Once again, the results are consistent with the partial segmentation hypothesis.

In Table 7, I present descriptive statistics and regressions of implied R-squares for CMA and RMW factors. The descriptive statistics in panel A here are qualitatively similar to those in Table 4. For instance, the average three-factor, two-factor and market R-squares here are 0.391, 0.381 and 0.346, respectively. The notable difference is that the two-factor R-square with CMA as the second factor here is higher than that (0.369) in the previous model with I/A as the second factor. This implies that CMA contributes more than I/A to the integration of the EREIT market with the general stock market.

Table 7
Descriptive statistics and regressions of implied R-squares for CMA and RMW factors

Panel A. Descriptive Statistics

Variable	Mean	Std. Err.	Min.	Max.
R-Sq. (MKT,CMA,RMW)	0.391	0.139	0.066	0.780
R-Sq. (MKT,CMA)	0.381	0.141	0.059	0.779
R-Sq. (MKT)	0.346	0.146	0.017	0.746

Panel B. Regressions of Implied R-Squares

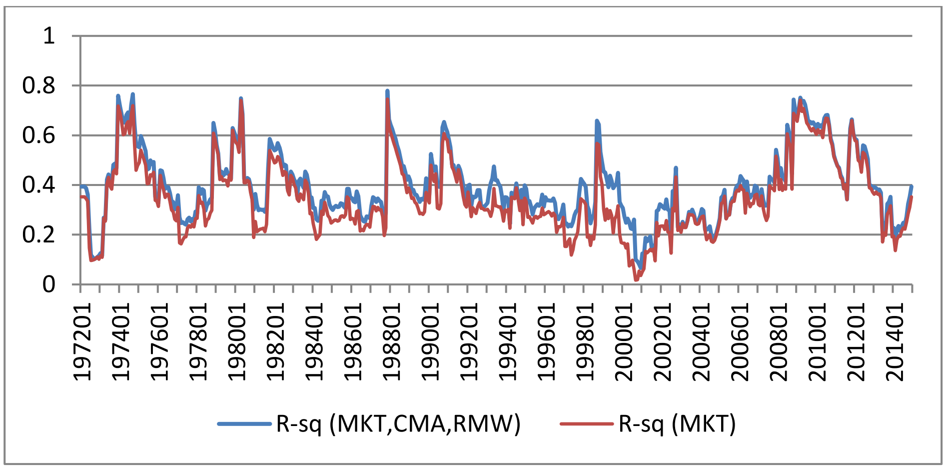
Dependent variable	Constant	Unemployment Rate	Real Estate Loan Growth	Default Spread	Adj. R^2
R-Sq. (MKT, CMA,RMW)	0.203	0.029			0.105
	(0.086)	(0.013)			
	0.423		**-0.043**		0.042
	(0.030)		(0.009)		
	0.269			0.110	0.135
	(0.048)			(0.047)	
	0.258	0.009	-0.030	**0.089**	0.168
	(0.121)	(0.014)	(0.012)	(0.029)	
R-Sq. (MKT)	**0.298**	**0.033**			0.148
	(0.080)	(0.012)			
	0.533		**-0.029**		0.019
	(0.023)		(0.015)		
	0.378			**0.122**	0.178
	(0.052)			(0.048)	
	0.315	0.017	-0.008	0.088	0.205
	(0.102)	(0.012)	(0.011)	(0.039)	

Notes: The R-sq's are implied by the three-factor model in Table 5. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic. All explanatory variables are one-month lagged variables. Standard errors in parentheses are robust to heteroskedasticity and residual serial autocorrelations up to 60 lags.

The regression results in panel B here are qualitatively similar to those in panel B of Table 4. In six univariate regressions, each of the three macroeconomic variables is significant at the 5 or 1 percent level and the sign of each coefficient is the same as before. The signs of the coefficients here in multiple regressions are also the same as before. The results echo the earlier finding that the level of integration tends to be high when the economy is weak, the credit risk is high or the credit availability is low.

The statistical significance of the coefficients in the multiple regressions here is slightly different from that in Table 4. With the market R-square as the dependent variable, only the default spread enters the multiple regression with a coefficient significant at the 5 percent level. This is in contrast to the previous result that both financial variables (real estate loan growth and default spread) are significant. With the three-factor R-square as the dependent variable, the coefficient of the unemployment rate is no longer significant here but the coefficient of the default spread turns significant at the 1 percent level. A comparison of the results from the regressions with the market R-square and three-factor R-square as the dependent variables reveals that the real estate loan growth captures the time variation of the three-factor R-square contributed by the additional factors (CMA and RMW). The time series patterns of the market R-square and three-factor R-square, as plotted in Figure 3, are similar to those in Figure 2. One noticeable exception is that CMA and RMW do not add as much as the I/A and ROE to the measured level of integration around the peak of the recent financial crisis.

Figure 3
Integration R-squares (MKT, CMA, RMW)



The R-squares are based on estimates in Table 5 for the three-factor model with MKT, CMA and RMW as factors.

D. The Role of SMB and HML Factors

Previous research suggests that the investment factor helps to explain the HML factor (see introduction). In this section, I replace the investment factor and profitability factor with the SMB (size) factor and the HML (value) factor and re-estimate the one- to four-factor models. I report the results in Table 8. The results of estimating the CAPM and the two-factor model are similar to those reported earlier. Once again, this suggests that the estimation results of the CAPM and the two-factor model are robust to the portfolios included in the estimation. The results of estimating the three- and four-factor models reveal some interesting findings. First, the estimated $a_{t<93}$ for EREIT is -0.733 with a standard error of 0.146, implying that the alpha of EREIT in the pre-1993 period is significant at the 1 percent level. In the two-factor model, the estimates of $a_{t<93}$ is -0.771, with a standard error of 0.258, implying that the alpha of EREIT in the pre-1993 period is also significant at the 1 percent level. Thus the models that include the I/A and ROE factors perform better than models with CMA and RMW factors or the models with SMB and HML factors. While the results of estimating the CAPM and the two-factor models here are similar to what Li (2016) finds, the finding here on the significance of the pre-1993 alpha in the three- and four-factor models are different.

Table 8

Estimates of models with equal risk prices including SMB and HML factors

	MKT	SMB	HML	EREIT	MKT	SMB	HML	EREIT
				Panel A. One and two-factor models				
$a_{t<93}$,%				**-0.784**				**-0.952**
				(0.150)				(0.192)
$a_{t\geq 93}$,%				-0.167				-0.378
				(0.165)				(0.149)
λ	**3.351**				1.965			**2.082**
	(0.901)				(0.871)			(0.789)
Diag. **A**	**0.916**	**0.904**	**0.906**	**0.885**	**0.919**	**0.905**	**0.906**	**0.885**
	(0.015)	(0.037)	(0.009)	(0.020)	(0.003)	(0.005)	(0.004)	(0.004)
Diag. **B**	**0.288**	**0.307**	**0.306**	**0.229**	**0.286**	**0.306**	**0.309**	**0.229**
	(0.040)	(0.090)	(0.028)	(0.038)	(0.011)	(0.013)	(0.015)	(0.016)
Diag. **D**	**-0.207**	**-0.253**	**-0.225**	**-0.370**	**-0.197**	**-0.240**	**-0.210**	**-0.345**
	(0.064)	(0.073)	(0.060)	(0.061)	(0.020)	(0.028)	(0.030)	(0.015)
				Panel B. Three- and four-factor models				
$a_{t<93}$,%				**-0.733**				**-0.771**
				(0.146)				(0.258)
$a_{t\geq 93}$,%				-0.156				-0.206
				(0.275)				(0.396)
λ	**4.405**	1.180	**5.989**		4.084	0.974	5.694	0.464
	(0.977)	(1.350)	(1.372)		(2.053)	(1.843)	(2.356)	(2.447)
Diag. **A**	**0.925**	**0.910**	**0.913**	**0.902**	**0.925**	**0.910**	**0.913**	**0.901**
	(0.007)	(0.029)	(0.011)	(0.006)	(0.013)	(0.021)	(0.015)	(0.022)
Diag. **B**	**0.260**	**0.336**	**0.292**	**0.331**	**0.260**	**0.336**	**0.292**	**0.331**
	(0.025)	(0.066)	(0.028)	(0.038)	(0.033)	(0.039)	(0.053)	(0.040)
Diag. **D**	**-0.186**	0.074	**-0.190**	-0.084	**-0.184**	0.075	**-0.190**	-0.081
	(0.029)	(0.142)	(0.053)	(0.061)	(0.062)	(0.210)	(0.069)	(0.103)
				Panel C. Likelihood ratio tests				
Integration hypothesis, $\lambda_4 = 0$						x^2	d.f.	p-value
One- vs. two-factor						2.239	1	0.135
Three- vs. four-factor						0.045	1	0.207

Notes: The return for each of the first three portfolios is given by equation (2): $R_{it}^e = \sum_{j=1}^{k} \lambda_j h_{ij,t} + \varepsilon_{it}$ The excess return on the EREIT portfolio is given by equation (3): $R_i^e = a_{t\leq T} I_{t\leq T} + a_{t>T} I_{t>T} + \sum_{j=1}^{k} \lambda_j h_{Nj,t} + \varepsilon_{N,t}$ where $a_{t\leq T}$ is the intercept for the period before date T and $a_{t>T}$ is the intercept for the period after date T. The conditional variance-covariance matrix $\mathbf{H_t}$ in follows the asymmetric BEKK GARCH specification: $\mathbf{H_t} = \mathbf{C}'\mathbf{C} + \mathbf{A}'\mathbf{H_{t-1}}\mathbf{A} + \mathbf{B}'\boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}_{t-1}{}'\mathbf{B} + \mathbf{D}'\boldsymbol{\eta}_{t-1}\boldsymbol{\eta}_{t-1}{}'\mathbf{D}$. Coefficients that are significant at the 1, 5 or 10 percent level are highlighted in bold, underlined or italic. Robust standard errors are reported in the parentheses.

Second, the estimated price of the EREIT risk in the four-factor model is insignificant with an estimate of 0.464 and a standard error of 2.447. The result is qualitatively similar to what I obtain from the other four-factor models studied earlier. Finally, the estimated prices of the MKT risk and HML risk are both significant at the 1 percent level in the three-factor model and at the 5 percent level in the four-factor model. However, the estimated price of the SMB risk in neither the three-factor model nor the four-factor model is significant at the 10 percent level. I find similar results about the insignificance of the price of the SMB risk when the investment factor or the profitability factor is replaced with SMB factor in the previous models. Other results about the tests and estimates of time-varying integration based on the SMB and HML factors are quite similar to what I discussed earlier.

VI. CONCLUSIONS

I find that a three-factor model that includes the market factor, the investment factor and the profitability factor used by H XZ (2015) is capable of capturing average returns on EREIT. The real estate premium in the two-factor model including the market factor and the real estate factor disappears when I account for premiums associated with the investment factor and profitability factor. This result implies that an EREIT portfolio is useful as a factor portfolio in forming mean-variance efficiency portfolios only when other factor portfolios such as the investment and profitability portfolios are excluded.

The results from testing market integration and segmentation hypotheses using three and four-factor asset pricing models suggest that the market for EREIT may be only partially integrated with the general stock market, consistent with the literature based on nonlinear cointegration tests. I also find that the integration is positively related to the unemployment rate and the default spread but negatively related to real estate loan growth. I find that the level of integration in the US market during and immediately after the financial crisis is much higher than the long-run average level of integration. The results suggest that the benefits of diversification between EREITs and the stock market tend to be lower during and immediately after the financial crisis when the economy is weak, the credit risk is high, or the credit availability is low. This result has important implication for mutual fund managers in asset allocation and security selection decisions.

The R-squares implied by the parameters in the dynamic GARCH-in-means model suggest that the degree of integration between EREIT and the general stock market moves counter-cyclically, which is consistent with the consumption-based asset pricing model with time-varying investor risk aversion (Campbell and Cochrane, 1999; Li, 2001). As the investor risk aversion moves counter-cyclically, prices of all risky assets including the stock market portfolio and EREIT tend to decline together during a recession or financial crisis. Such declines may be induced by investors' consumption needs as a high unemployment rate is associated with low income, slow real estate loan growth means less credit availability, and a high default spread predicts a weak economy and low future cash flows.

I compare the results from the three-factor model mentioned above with alternative models in which the investment and profitability factors (I/A and ROE) used by Hou et al. (2015) are replaced with similar factors (CMA and RMW) used by Fama and French (2016) or the size and value factors (SMB and HML) used by the Fama and French (1993). I find that alternative three-factor models fail to capture the average return on EREIT. However, the results on testing and measuring market integration from the alternative models are similar to those obtained from our main three- and four-factor models used in the paper.

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