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Long-term Cointegrative and Short-term Causal Relations among U.S. Real Estate Sectors

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We investigate long-term cointegrative and short-term causal relations among seven U.S. sectoral REITs. First, cointegration tests identify one long-term cointegrative relation among five of the sectors, which suggests that two of the sectors are outside the cointegrative space. Second, short-term Granger causality tests identify three leading and two following cointegrated sectors. Third, a proposed vector autoregressive model indicates that a stronger cointegrating effect is induced by declining real estate markets and a multivariate sensitivity regression model shows that unexpected inflation significantly and negatively influences the cointegrative disequilibrium. Lastly, our cointegration-based portfolio performance analyses show that the inferior performance of the all-sector market portfolio stems from containing the redundant cointegrated sectors which shatter portfolio diversification.

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Keywords

Cointegration; Domestic Real Estate Sector; Error Correction Model; Portfolio Construction and Diversification; Vector Autoregressive Model

1. Introduction

There has been a recent research interest in the integration and cointegration of global real estate markets and international real estate portfolio diversification by Yunus and Swanson (2007, 2013), Glascock and Kelly (2007), Yunus (2009), Gallo and Zhang (2010), and Gallo, Lockwood, and Zhang (2013). The resulting extension of knowledge in these areas has been facilitated in particular by two developments. First, Yunus (2009) and Gallo and Zhang (2010) show that econometric advances have improved upon correlation analysis, and cointegration analysis has emerged as a productive tool for detecting and understanding long-term time series relations. At the same time, public real estate indices in developed global markets have matured to the point where they are now available and appropriate for cointegration analysis to reveal underlying long-term relations. Although prior studies, such as Chaudhry, Myer and Webb (1999), Yunus (2009), Yunus and Swanson (2013), Gallo and Zhang (2010), and Gallo, Lockwood and Zhang (2013), report the existence of cointegrating real estate markets, none of these studies focus on the causes of the cointegrative disequilibrium within the cointegrative space. In this study, we investigate both the long-term cointegrative and short-term causal relations among seven U.S. sector real estate investment trusts (REITs). Importantly, we propose a vector autoregressive model and a multivariate sensitivity regression model to study the causes of the cointegrative disequilibrium.

This study offers three important findings to the real estate literature. First, Chaudhry, Myer, and Webb (1999) report a long-term cointegration relation among four real estate sectors: office, retail, research and development office, and warehouse. In this study, we extend their work by examining both long-term cointegrative and short-term causal relations among seven real estate sectors. Johansen cointegration tests identify one long-term cointegrative relation among five of the sectors, which suggests that two of the sectors are outside the cointegrative space. Granger causality tests identify a short-term causal relation among the five cointegrated sectors, and three leading and two following cointegrated sectors within the cointegrative vector (CIV, hereafter). Second, we study the causes of cointegrative disequilibrium. To this end, we first design a novel vector autoregressive model and report that a stronger cointegrative effect (smaller cointegrative disequilibrium) is induced by declining real estate markets. We then conduct a multivariate sensitivity regression model to examine whether macroeconomic factors, real estate specified factors, and U.S. stock market returns have impact on the cointegrative disequilibrium. The results show that unexpected inflation significantly and negatively influences the cointegrative disequilibrium. Third, by eliminating redundant real estate sectors and constructing a portfolio with two segmented and three leading cointegrated sectors which are considered essential sectors, we conclude a cointegration-based real estate sector portfolio (COI, hereafter) has excess as well as risk adjusted returns greater than the all-sector CRSP/Ziman benchmark portfolio (MKT, hereafter). Tested with a REIT-based four-factor model, our COI strategy provides significant abnormal returns to investors under down markets in which investors most need diversification. We conclude that the inferior performance of the all-sector MKT portfolio stems from containing the two superfluous cointegrated sectors which shatter portfolio diversification.

2. Literature Review

Concern about the reliability of correlation structures in studying asset returns has led to studies that adopt the alternative of cointegration analysis. Tarbert (1998) applies this to two different series of UK direct property returns (1971-1995 and 1977-1995) and find, in general, a tendency for returns to be cointegrated across both region and property type, which indicates restricted diversification potential. By using a similar cointegration method, Myer, Chaudry, and Webb (1997) examine the 1987-1992 time-series of both aggregated and disaggregated direct property values for the U.S., U.K., and Canada, and find a common factor between these markets, which indicates reduced cross-diversification benefits in the long-term. Wilson and Okunev (1996), also by using a cointegration approach, demonstrate segmentation between the indirect property markets and stock markets of the U.S., U.K., and Australia, which indicates diversification potential across these assets within the national markets. They also show, in support of international diversification, that the same is true for indirect property markets across these countries[1].

Yunus and Swanson (2007) investigate long-term relations and short-run linkages across the U.S. and several Asia-Pacific indirect real estate markets over the period 2000-2006.[2] Over this period, they find that these markets are not cointegrated. They conclude that this is progressively becoming less so, although it is not happening across all sub-markets, which means some long-term diversification opportunities still exist. By using a longer time period (1990 through 2007), Yunus (2009) extends this analysis, and adds the Netherlands, France and the U.K. to a similar U.S./Asia-Pacific market set.[3] She finds that most international public real estate markets are cointegrated,

[1] Wilson and Okunev (1996) use different sets of times series, dictated by availability. The longest is 1969-1993.
[2] Australia, Japan, Hong Kong and Singapore.
[3] Yunus (2009) does not include Singapore.

and this trend to integration is increasing over time.[4] Other studies that have adopted the cointegration method in the real estate literature include Chaudhry, Christie-David, and Sackley (1999), Kleiman, Payne and Sahu (2002), Gallo and Zhang (2010), Gallo, Lockwood, and Zhang (2013), and Yunus and Swanson (2013).

There are some studies in the real estate literature that evaluate diversification benefits across property types. Hartzell, Hekman, and Miles (1986) focus on region and property type within a domestic commercial real estate opportunity set and argue that property type offers certain diversification benefits within the real estate portfolio. Similar results are reported by Eichholtz, Hoesli, MacGregor and Nanthakumaran (1995) who analyze data from both the U.S. and the U.K. Domestically, however, Gyourko and Nelling (1996) provide no evidence that REIT diversification across property type or broad geographic regions actually result in meaningful diversification. Chaudhry, Myer and Webb (1999) conduct cointegration tests to analyze interactions among four U.S. real estate sectors. They find one long-term cointegrative relation among these four sectors which implies limited diversification benefit of holding cointegrated sectors. Lee and Byrne (1998) and Lee (2001) study data on real estate regions and sectors in the U.K. and conclude that the performance of real estate is largely property type-driven which suggests that the property type composition of the real estate fund should be the first level of analysis in constructing and managing the real estate portfolio. Young (2000) reports that equity-REITs grouped by property-type sectors have become more integrated over the 1989 to 1998 period as evidenced by increasing correlation over time. More recently, Glascock and Kelly (2007) examine and test the merits of diversifying portfolios of real estate securities internationally and across property type, and find that property type effects are smaller than country effects.

The causes of disequilibrium/disequilibria within the cointegration space have been studied in stock markets (Arshanapalli and Doukas 1993, Richards 1995, and Wood 1995), foreign exchange markets (Kim 2003, Narayana and Smytha 2004), and commodity markets (Mananyi and Struthers 1997, Swaray 2008). However, none of the aforementioned real estate studies look into the causes of disequilibrium within the cointegration relation. In an attempt to fill this gap, we design a vector autoregressive model and a multivariate regression model to investigate the causes of disequilibrium in the cointegration relation among real estate sectors.

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[4] The exceptions are France and the Netherlands, a circumstance that Yunus (2009) suggests may be associated with their convergence with the real estate markets of the Euro zone.

3. Data

We examine the period from March 1984 to December 2009. We obtain monthly returns and market capitalization series for seven REIT sectors [5] (healthcare, industrial/office, residential, lodging/resort, retail, self-storage, and unclassified) from the Center for Research in Securities Prices (CRSP)/Ziman U.S. Real Estate Data Series. [6] Unlike other real estate databases, such as the FTSE NAREIT, SNL REIT, and S&P/Citigroup REIT, the CRSP/Ziman database provides the most comprehensive property type data at both firm and index levels. [7] In addition, all REITs in the Ziman database are publicly traded in the U.S. stock markets so that trading barriers and illiquidity biases are non-issues. The CRSP/Ziman database has been widely used in recent REIT-related studies, such as Chiang (2010), Chen, Downs and Patterson (2012), Ling, Naranjo and Ryngaert (2012), among others.

Importantly, the CRSP/Ziman database which includes all REITs that have traded on the NYSE, AMEX and NASDAQ exchanges since 1980, contains a series of indices categorized by property type, underlying individual REIT information for each index, and qualitative measures important for evaluating information in the thinly populated index series. [8] With all U.S. individual

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[5] Although we focus on seven property type indices, each specialized property type index contains all individual REITs categorized in this particular property type in the CRSP/Ziman database. A portfolio that contains seven property indices would include the universe of all individual REITs that are trading on the NASDAQ, New York Stock Exchange, and American Stock Exchange.

[6] A note on the unclassified sector is warranted. We contacted the CRSP to ask about the definition and composition of the 'unclassified' category. The CRSP informed us that this is a catch-all category for asset returns that do not fit into the any of the six explicit classifications. No further information was available. By following Ro and Ziobrowski (2011) who also use the CRSP/Ziman data series, we include the 'unclassified' category in our analysis.

[7] Note that CRSP/Ziman contains no direct property data. Certain REITs have international real estate in their portfolios. This paper, however, stands from the perspective of U.S. domiciled investors and all prices are dollar denominated so that the currency effect is not a concern.

[8] Combining stock price and return data with carefully researched information with regard to the population, characteristics, and history of REITs, the CRSP/Ziman database provides firm-specific information and indices essential to analyses that involve this important asset class. This database includes several qualitative measures which detail market capitalizations, concentrations, and changes in index composition particularly important for evaluating the information in thinly populated index series which were common during the 1980s. The CRSP/Ziman database includes all REITs that have traded on the NYSE, AMEX and NASDAQ exchanges since 1980, a series of indices based on REIT and property type, underlying individual security information for the indices, and qualitative measures important for evaluating information in thinly populated index series: market capitalization, concentration, and changes in index composition.

REITs available in the CRSP/Ziman database, we are able to construct U.S. REIT-based risk factors. By following Hartzell, Mühlhofer and Titman (2010) and Cici, Corgel and Gibson (2011), we create Fama-French-Carhart REIT-based four factors from the universe of all individual REITs over the 1984 to 2009 period. First, we use the value-weighted CRSP/Ziman REIT market index as the market portfolio (MKT). The other factors are the return differentials between the small cap and large cap REITs (Size), high and low book-to-market REITs (Book-to-Market), and positive and negative prior year return-momentum REITs (Momentum). The method for constructing these factors is based exactly on Fama and French (1993) and Carhart (1997).[9] To further examine the causes of cointegrative disequilibrium, we follow Peterson and Hsieh (1997) and conduct a multivariate sensitivity regression model with seven factors from Lee and Chiang (2004) which include four macroeconomic factors, two real estate related factors, and one stock market factor. In particular, we obtain four U.S. macroeconomic factors used in a study by Liu and Zhang (2008): industrial production growth rate, unexpected inflation rate, term structure, and default risk premium, which is taken from Professor Laura Xiaolei Liu's website (http://www.bm.ust.hk/~fnliu/research.html); a 30-year conventional mortgage rate from the Federal Reserve Bank of St. Louis; new, privately owned housing unit start data from the U.S. Census Bureau; and the S&P500 returns from the Center for Research in Securities Prices.

4. Research Methods

4.1 Unit Root and Cointegration Tests

Cointegration methods, developed by Engle and Granger (1987), are based on error correction models (ECMs, hereafter), rather than correlations, to identify long-term equilibrium among a set of non-stationary variables (REIT price indices in this study). Cointegrated indices share a linear combination of non-stationary variables. The stationarity of indices can be identified with unit root tests.

First, we conduct five unit root tests, augmented Dickey-Fuller (1981, ADF hereafter), Phillips-Perron (1988, PP hereafter), Kwiatkowski, Phillips, Schmidt, Shin (1992, KPSS hereafter), Zivot-Andrews (1992, ZA hereafter), and Ng and Perron (2001, NP hereafter) to test for price series stationary. The REIT monthly returns of each sector are converted into price series, which is then converted into natural logarithms for the unit root tests. Second, we employ Johansen (1988, 1992a) cointegration rank tests to assess long-term cointegrative relations among non-stationary real estate sector indices. Johansen (1991) exclusion tests are then performed to identify sectors

[9] Our four REIT-based factors are different from the U.S. equity-based factors used in Ro and Ziobrowski (2011).

independent of the cointegrative space. Third, we conduct the Granger causality test to elucidate the short-term causal relation and lead-lag linkage among cointegrated sectors.

Similar to Johansen, Mosconi and Nielsen (2000), we employ likelihood-ratio (L-R) tests to identify any possible structural breaks on the price level.[10] Dummy variables are added as controls for large residual shocks (significant at the 0.01 level) that could induce non-normally distributed residuals. The Bartlett small sample correction test suggested by Johansen (2002) is included in each cointegration rank test to mitigate potential small sample size bias. We report both the Johansen trace and the Bartlett-corrected trace statistics.

4.2 Causes of Cointegrative Disequilibrium

Cointegration ECM residuals generated from the above long-term cointegration tests allow us to further examine both market conditions and economic factors that possibly drive the cointegrative disequilibrium among cointegrated real estate sectors. First, Liow, Ho, Ibrahim and Chen (2011), Stevenson (2002), Wilson, Stevenson and Zurbruegg (2007), and Hoesli and Reka (2011)examine co-movement and volatility spillovers across real estate markets and reveal a greater degree of co-movement and increased volatility spillover among real estate markets under real estate down markets.

To study the causes of cointegrative disequilibrium, we first propose a novel vector autoregression (VAR) analysis with an interactive dummy to examine the influence that down markets have on cointegrative disequilibrium. As described by Engle and Granger (1987), the CIV of the real estate sector indices, Y_{it}, takes a long-term equilibrium form:

$$\beta_1^* Y_{1t} + \beta_2^* Y_{2t} + \ldots + \beta_n^* Y_{nt} + \delta^* D = 0 \tag{1}$$

where $Y_{1t} \ldots Y_{nt}$ are the price levels of the real estate sector indices, $\beta_1 \ldots \beta_n$ are the eigenvector coefficients and D is the deterministic component (i.e., constant, linear time trend, etc.). The CIV residual, ϑ_t, or cointegrative disequilibrium, is calculated as follows:

$$\vartheta_t = \mathrm{B}^* \mathrm{Y}_t' = (\beta_1, \beta_2, \ldots, \beta_n, \delta)^* (Y_{1t}, Y_{2t}, \ldots, Y_{nt}, D)' \tag{2}$$

where B and transposed Y_t' denote the vectors $(\beta_1, \beta_2, \ldots, \beta_n, \delta)$ and $(Y_{1t}, Y_{2t}, \ldots, Y_{nt}, D)'$ respectively. Although not reported, the cointegrative disequilibrium or the deviation from the long-term cointegration relation, ϑ_t, is proven to be a stationary *I(0)* process by our unit root tests.

[10] In addition to the likelihood-ratio (L-R) test from Johansen, Mosconi and Nielsen (2000), we also conduct the multiple structural break test from Bai and Perron (2003). We find a consistent result in that no structure break is detected for the log of price of any sector.

Given that a previous residual ϑ_{t-1} is positive (negative) at time t-1, this means that the sector deviated from its bonded long-term equilibrium with a positive (negative) error. Therefore, the sector price level Y_t at current time t should move down (up) with a negative (positive) return towards its long-term equilibrium to correct its past error should the cointegrative relation hold. Therefore, we expect that there is an inverse relation between the current index return at time t and its past cointegrative disequilibrium at time t-1. Up (down) markets are collectively defined as months in which the excess return of the broad real estate market index is positive (non-positive), as specified by Fabozzi and Francis (1977). Specifically, the VAR model is as follows:

$$\Delta Y_t = \rho_0 + \rho_1 \vartheta_{t-1} + \rho_2 \varpi_t + \rho_3 (\vartheta_{t-1} * \varpi_t) + \sum_{\eta=1}^{k} \rho_{\eta+3} * \Delta Y_{t-\eta} + \psi_t , \quad (3)$$

where ΔY_t is the vector of the real estate sector returns in a VAR system at time t, ϑ_{t-1} is the lagged cointegrative disequilibrium at time t-1, ϖ is a binary market condition dummy variable at time t (1 = declining market and 0 = rising market), $(\vartheta_{t-1} * \varpi_t)$ is the interactive term between past cointegrative disequilibrium and the current market condition dummy, $\sum_{\eta=1}^{k} \rho_{\eta+3} * \Delta Y_{t-\eta}$ is the sum of k lagged real estate sector returns in the VAR system, and ψ represents the error term of the VAR system. When ϖ =0 (ϖ =1), ρ_1 depicts how sector REIT returns respond to the past disequilibrium in a rising (declining) market condition. In general, a positive ρ_1 implies that real estate sectors tend to become more segmented in rising real estate markets. On the other hand, a negative $(\rho_1 + \rho_3)$ implies that real estate sectors tend to become more cointegrated in declining real estate markets. ρ_0 is the intercept when ϖ =0 while $(\rho_0 + \rho_2)$ is the intercept when ϖ =1.

Second, prior studies show that there is a strong linkage between macroeconomic factors and the U.S. real estate market, such as Chan, Hendershott, and Sanders (1990), McCue and Kling (1994), and Ling and Naranjo (1997). To examine the causes of contemporaneous cointegrative disequilibrium, we follow Peterson and Hsieh (1997) and conduct a multivariate sensitivity regression model with the absolute value of the white noise CIV residual ϑ_t from the ECM model as the endogenous variable. As one can deviate from its bonded cointegrative relation from above or below, the absolute value of the residual captures the absolute extent of the cointegration error. With regard to exogenous variables, we follow the seven factors from Lee and Chiang (2004) that include four macroeconomic factors, two real estate related factors, and one stock market factor. We test the impact of changes in each of the seven economic factors on the contemporaneous cointegrative disequilibrium. We perform the following multivariate sensitivity regression model:

$$|\vartheta_t| = \lambda_0 + \lambda_1 INDPRO_t + \lambda_2 INFL_t + \lambda_3 TS_t + \lambda_4 DEF_t + \lambda_5 MTGED_t + \lambda_6\, HSTARTS_t + \lambda_7\, SP500R_t + \tau_{pt}, \qquad (4)$$

where $INDPRO_t$ is the industrial production growth rate, $INFL_t$ is the unexpected inflation rate which is defined as the change of the seasonally adjusted Consumer Price Index, TS_t is the term structure which is defined as the (20-yr)-(1-yr) yield, DEF_t is the risk premium which is defined as the BAA-AAA yield, $MTGED_t$ is the change of the 30-year conventional mortgage rate, $HSTARTS_t$ is new private housing starts, and SP500R is the S&P 500 index return at time t. All seven factors are on a monthly basis. For rigorous purposes, we conduct a robust regression to compute heteroskedastic robust standard errors. We also test for multicollinearity by calculating the mean variance inflation factors (VIFs) for the exogenous variables. In addition, we perform Durbin-Watson testing (d-statistics) to monitor any residual autocorrelation.

4.3 Implications of the Cointegrative Relation

The investment value of applying a cointegration approach to domestic real estate sectors can be measured by testing the performance of a portfolio that consists of only the essential real estate sectors (segmented plus leading cointegrated sectors). In line with Pesaran, Shin, and Smith (2000), Gallo, Phengpis and Swanson (2007), Yunus (2009), and Gallo, Lockwood, and Zhang, (2013), we argue that the cointegrated real estate sectors offer limited diversification potential and should be excluded in diversified portfolios while the only cointegrated sectors that deserve allocations are leading cointegrated markets which are the source of common trends among cointegrated markets.

On the portfolio level, we examine two risk-adjusted performance measures. First, we perform significance tests of Sharpe ratio (SHP, hereafter) differences across portfolios by using the Jobson and Korkie (1981) z-statistics. Second, we examine portfolio performance after controlling for the portfolio's exposure to the REIT-based market, size, style and momentum factors:

$$R_{pt} - R_{ft} = a + b_1(ReitR_{MKT} - R_{ft}) + b_2 ReitSMB_t + b_3 ReitHML_t + b_4 ReitUMD_t + e_{pt}, \qquad (5)$$

where R_{pt} is the monthly raw return of the real estate portfolio, R_{ft} is the monthly Citigroup 3-month Treasury bill return, $ReitR_{MKT}$ is the value-weighted CRSP/Ziman REIT market index, ReitSMB is the REIT size factor, HML is the REIT style factor, and ReitUMD is the REIT momentum factor.

We test the performance of each portfolio by examining the sign and significance of the estimate of the intercept. The intercept measures the incremental performance (abnormal return) of the portfolio after controlling for exposure to the broad market, size, style, and momentum factors. A significant positive (negative) *t*-statistic for the intercept estimate indicates superior (inferior) risk-adjusted performance.

Next, we hypothesize that cointegrated real estate sector indices become more cointegrated in down markets so that cointegrated redundant sectors lack any diversification benefit during down markets while essential ones should outperform. We therefore conduct a performance test by running Equation (5) over varying market conditions based on the performance of the benchmark MKT. As defined above, an up market is the months in which the excess MKT return is positive. A down market is defined as the months in which the excess MKT return is non-positive. Between 01/1995 and 12/2009, our sample is demarcated with 108 months of up markets and 72 months of down markets. The first portfolio formation in 01/1995 is based on cointegration tests over the 03/1984 to 12/1994 window. The portfolio is then held for 12 months and rebalanced every January thereafter.[11]

Finally, similar to Chaudhry, Myer, and Webb (1999) and Yunus, Hansz, and Kennedy (2012), we present the model specification for price discovery and cointegration forecasting purposes based on the entire sample period from 1984 to 2009.

5. Empirical Results

Descriptive statistics are calculated over a 310-month sample period, March 1984 to December 2009, for each REIT sector index and are reported in Table 1. Other than the means, standard deviations, market capitalizations, and SHP, we also report the Jobson-Korkie statistics for equality of the SHP of the REIT sector versus that of the MKT.

The mean monthly return of the MKT equaled 0.82% with a 0.09 SHP over the full sample period. The mean returns of the REIT sectors exhibit substantial variation, ranging from a low of 0.25% for LODG to a high of 1.33% for HEAL. The Jobson-Korkie statistics indicate that only the SHP of HEAL is significantly larger than that of the MKT (z-stat=2.24). The SHP of the other sector indices are either smaller than or indifferent from that of the MKT. The mean market capitalization shows that INDU had the highest market value with a 27.39% market share among all other REIT sectors while

[11] For example, the second portfolio formation in 01/1996 is based on cointegration tests over the 01/1986 to 12/1995 window and the third portfolio formation in 01/1997 is based on cointegration tests over the 01/1987 to 12/1996 window, and so on and so forth.

Table 1 Descriptive Statistics

INDEX	RET(%)	SD (%)	SHP	J-K z-stat (vs. MKT)	Mkt. Cap.	Mkt. Share	Count	ConRatio	R-Sqr	P-value
Market Index	0.82%	4.95%	0.09		$124,000	100%	170.19	17.31%		
3MTB	0.38%	0.19%								
REIT Type										
Unclassified (UNCL)	0.54%	5.21%	0.03	-1.36	$5,790	5.41%	18.17	73.71%	0.52	0.00***
Health Care (HEAL)	1.33%	5.56%	0.17	2.24**	$8,480	7.93%	10.88	70.21%	0.63	0.00***
Industrial/Office (INDU)	0.57%	6.30%	0.03	-2.16**	$29,300	27.39%	29.43	53.72%	0.79	0.00***
Lodging/Resorts (LODG)	0.25%	9.10%	-0.01	-2.41**	$9,214	8.61%	10.01	85.69%	0.52	0.00***
Residential (RESI)	0.97%	5.23%	0.11	0.84	$19,900	18.60%	22.16	58.79%	0.77	0.00***
Retail (RETL)	1.04%	5.79%	0.11	1.31	$28,700	26.83%	31.19	51.11%	0.88	0.00***
Self Storage (SELF)	1.11%	6.00%	0.12	0.76	$5,577	5.21%	7.26	88.04%	0.50	0.00***

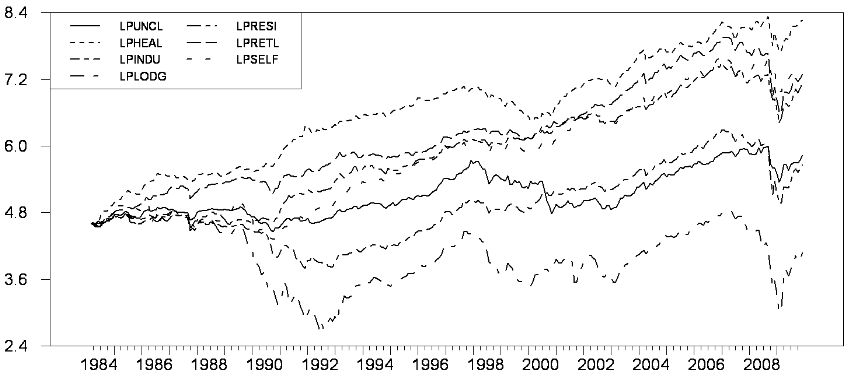
Notes: This exhibit summarizes the monthly performance of the value-weighted REIT indices examined over a 310-month period,03/1984-12/2009. Market Index is the CRSP/Ziman value-weight market index. 3MTB is the 3-month Treasury bill. For each index, we report the raw mean return (RET), standard deviation of returns (SD), Sharpe ratio (SHP), and Jobson-Korkie statistic (J-K z-stat) for the equality of the SHP with the SHP of the MKT. Mkt. Cap. represents the average market value of each REIT in millions. Count represents the average number of REITs eligible for inclusion in the index. ConRatio is the concentration ratio that is the ratio of the market value of the largest four securities in the portfolio versus the market value of the entire portfolio computed by using the beginning of period market caps. To examine by how much the market return can be explained by a certain real estate type return, we adopt a single factor model as follows:

$$RtMKT = AlphaP + BetaP*RtP + Errp$$

where RtMKT represents the REIT market raw return and RtPis the raw return of the real estate type index P. R-Sqr shows the goodness of fit of the single factor model. P-value is for the F-test of model fitness. ***and ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

SELF with a 5.21% market share is ranked the lowest. The count ranged from 7.26 (SELF) to 31.19 (RETL), and on average, there was about 170 REITs included in the market index over the sample period. The ConRatio indicates that the market value of the largest four REITs occupied 88.04% (highest) of the SELF portfolio while it was 51.11% (lowest) for the RETL portfolio. In the spirit of Glascock and Kelly (2007), we run the MKT returns against each real estate sector return to see by how much the MKT can be separately explained by each sector. We find large R-squares under each real estate sector regression with significant model F-test statistics which mean that real estate sectors explain a large portion of the return variation in the domestic real estate MKT. These findings are different from those of Glascock and Kelly (2007), in which there is a 6% explanatory power of the sectors in the international real estate markets. We conclude that real estate sectors provide important diversification within the U.S. REIT market. The time series plot of the log of price for these seven real estate sectors is presented in Figure 1.

Figure 1 Log of Price for Seven Real Estate Type REITs (03/1984-12/2009)



Notes: Test period is from 03/1984 to 12/2009 which includes 310 monthly observations. Figure 1 plots the log of price for seven real estate type REITs: LPUNCL is the log price of unclassified REIT index; LPHEAL is the log price of healthcare REIT index; LPINDU is the log price of industrial/office REIT index; LPLODG is the log price of lodging/resort REIT index; LPRESI is the log price of residential REIT index; LPRETL is the log price of retail REIT index; and LPSELF is the log price of self-storage REIT index.

The cointegration methodology begins with unit root tests on the seven sector price indices. We conduct five unit root tests, ADF, PP, KPSS, ZA, and NP, to test the stationarity of each sectoral price index. Except for the KPSS test, all the other unit root tests examine the null hypothesis of a unit root and non-stationary against the alternative hypothesis that no unit root is present and the data series is stationary. As shown in Table 2, we find that each of these seven sector price indices has a unit root representation which is a non-stationary

I(1) process. No structure break is detected by the ZA test on the price level. Therefore, unit root tests merit the implementation of the cointegration methodology to further detect long-term equilibrium among non-stationary sectoral price indices.

Table 2 Unit Root Tests from 03/1984 to 12/2009

Sector Index	ADF	PP	KPSS (mu)	ZA (Break)	$MZ_{\alpha}^{(a)}$	$MZ_{t}^{(b)}$
UNCL	-2.04	-1.00	4.35***	-3.92 (2000:08)	-0.03	-0.02
HEAL	-1.38	-1.35	5.90***	-4.14 (1998:04)	1.27	2.14
INDU	-1.60	-0.76	4.41***	-3.88 (1990:07)	-1.55	-0.68
LODG	-7.42	-1.89	0.79***	-4.62 (1989:10)	-7.35	-1.89
RESI	-1.01	-0.53	6.17***	-3.70 (2006:01)	1.11	1.44
RETL	-2.13	-1.22	5.90***	-3.16 (2004:05)	0.79	0.92
SELF	-0.04	-0.02	6.03***	-3.54 (1989:10)	1.41	1.82

Note: Test period is the 310-month period from 03/1984 to 12/2009. Five unit root tests are performed on the price levels of each data series: augmented Dicky-Fuller (ADF), Phillips-Perron (PP), Kwiatkowski, Phillips, Schmidt, Shin (KPSS), Zivot-Andrews (ZA) and Ng-Perron (NG). ADF, PP, ZA, NG tests all examine the null hypothesis of a unit root and nonstationary against the alternative hypothesis that no unit root is present and the data series is stationary. The KPSS tests the null hypothesis of no unit root present and the data is stationary. All tests allow for a maximum of 12 lags. The ZA test uses AIC criteria to decide the lag length from a maximum of 12 lags. ADF joint test of unit root critical values: 1%= 6.47, 5%= 4.61, and 10%= 3.79; PP unit root test critical values: 1%= -3.468, 5%= -2.878, and 10%= -2.575 (reported statistics are with 4 lags); ZA unit root test (Model C which allows break=both, maximum 12 lags) critical values: 1%= -5.57, 5%= -5.08, and 10%= -4.82; KPSS unit root test with mu statistics (H0: stationary around a level) critical value: 1%= 0.739 and 5%= 0.463 and 10%=0.347 (reported statistics are with 4 lags). Ng and Perron (2001) MZ_a statistic critical value: 1%= -23.80, 5%= -17.30, and 10%= -14.20. Ng and Perron (2001) MZ_t statistic critical value: 1%= -3.42, 5%= -2.91, and 10%= -2.62. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

The results of the cointegration tests are presented in Table 3. To ensure that our cointegration tests are not look-ahead biased and suitable for an out-of-sample portfolio formation, we follow Phengpis and Swanson (2010) and run cointegration tests by using a 120-month rolling window approach with the first window that ranges from 03/1984 to 12/1994 which instructs us to form the COI portfolios on 01/1995 and hold them for the next 12 months. The second window is from 01/1986 to 12/1995 for the 01/1996 formation, and so on and so forth until the last window is from 01/1999 to 12/2008 for the 01/2009 formation. From the cointegration tests for fifteen separate rolling windows, we reach consistent results in terms of cointegrated/segmented and leading/following sectors. This finding is in line with the essence of the cointegration framework, that in general, a long-term stable cointegrative relation is warranted should the cointegrating ECM hold. Furthermore, we present a recursive cointegration test of CIV stability to show that the cointegration test results are stable over time from 1995-2009. To conserve space, we only report the results from the cointegration tests for the first window. Other subsequent windows and the full period results are available from the authors upon request.

Although the price series for all seven REIT sectors are with unit root representation over the entire sample period as reported in Table 2, it is necessary to further confirm if they are an *I(1)* process under each sub-period in which we conduct the cointegration tests. Table 3A first reports chi-square testing with stationary as the null hypothesis given the cointegration space. We confirm that all sector price indices are non-stationary. Table 3B shows a significant CIV among the seven real estate sector indices (Bartlett λ_{trace}=152.22, p-value=0.04) with insignificant G(r) common linear trend components (p-value=0.17). The results from the cointegration exclusion tests are reported in Table 3C. Sector indices with significant (insignificant) L-R test statistics are cointegrated (segmented). The findings indicate that five sectors: HEAL ($L\text{-}R_{HEAL}$=5.81), INDU ($L\text{-}R_{INDU}$=7.15), LODG ($L\text{-}R_{LODG}$=5.49), RETL ($L\text{-}R_{RETL}$=3.62) and SELF ($L\text{-}R_{SELF}$=4.27), are cointegrated and share one CIV.[12] However, the L-R statistics are insignificant for UNCL ($L\text{-}R_{UNCL}$=0.55) and RESI ($L\text{-}R_{RESI}$=0.39), which imply their segmentation from the cointegrative space. The robustness test, presented in Table 3D, confirms there is one and only one significant CIV among all five cointegrated real estate sectors (Bartlett λ_{trace}=72.07, p-value=0.03) with a similar insignificant G(r) common linear trend component (p-value=0.28). Table 3E further confirms that none of the cointegrated REITs are segmented from the cointegrative space. Chaudhry, Myer, and Webb (1999) find one cointegative relation among four real estate sectors and there is a positive long-term cointegrative relation between retail and office. Our results support the findings of Chaudhry, Myer, and Webb (1999). From the Johansen cointegration rank and exclusion tests, we conclude that HEAL, INDU, LODG, RETL, and SELF are cointegrated while UNCL and RESI areoutside the cointegrating space.

[12] The residual of our ECM model is stationary with 1 lag. The length of our lag is determined by Ljung-Box statistics. Insignificant Ljung-Box statistics imply no autocorrelation once 1 lag is imposed in the model. Although not reported, the test result with 2 lags consistently finds 1 cointegrating vector. Since the use of more lags sacrifices the degree of freedom, we keep the most parsimonious model with 1 lag.

Table 3A Stationarity Tests on all Indices (from 03/1984 to 12/1994)

	UNCL	HEAL	INDU	LODG	RESI	RETL	SELF
L-R statistic	31.13	36.20	44.54	36.40	38.60	39.76	38.38
P-value	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***

Table 3B Cointegration Rank Tests on all Indices (from 03/1984 to 12/1994)

I(1)-Analysis (n=7, lag=1)	G(r)	p-r	r	Eig. Value	Trace	Bartlett Trace	P-Value	Bartlett P-Value
	1.90	7	0	0.29	156.63	152.22	0.02**	0.04**
	P-value=0.17	6	1	0.25	111.63	108.99	0.11	0.16

Table 3C Exclusion Tests among all Indices (from 03/1984 to 12/1994)

Sector Index (n=7)	r	DF	5% C.V.	UNCL	HEAL	INDU	LODG	RESI	RETL	SELF
L-R statistic	1	1	3.84	0.35	5.81	7.15	5.49	0.76	3.62	4.27
P-value				0.55	0.02**	0.01**	0.02**	0.39	0.06*	0.04*

Table 3D Cointegration Rank Tests on Five Cointegrated Indices (from 03/1984 to 12/1994)

I(1)-Analysis (n=5, lag=1)	G(r)	p-r	r	Eig. Value	Trace	Bartlett Trace	P-Value	Bartlett P-Value
	1.15	5	0	0.26	73.49	72.07	0.02**	0.03**
	P-value=0.28	4	1	0.11	34/86	34.34	0.46	0.49

(Continued…)

(Table 3 Continued)

Table 3E Exclusion Tests among Five Cointegrated Indices (from 03/1984 to 12/1994)

Sector Index (n=5)	r	DF	5% C.V.	HEAL	INDU	LODG	RETL	SELF
L-R statistic	1	1	3.84	8.20	13.49	10.38	3.28	4.83
P-value				0.00***	0.00***	0.00***	0.07*	0.03**

Notes: The G(r) statistic is distributed chi-square with r degrees of freedom and used to detect common linear trends among any of the real estate sector indices, p is the number of dimensional vectors of real estate sector indices, r is the number of cointegrated vectors, Eig Value is the eigenvalue obtained from maximum likelihood estimation of the ECM. Trace is Johansen trace statistic for cointegration rank test; Bartlett Trace is the Bartlett-small-sample-corrected trace statistics, P-value is the probability value for the Johansen trace statistics. The final column reports the p-value for the Bartlett trace statistic. Table 3A reports chi-square stationary test as the null hypothesis given the cointegration space. Table 3C presents the exclusion test results based on the rank tests in . Table 3B. Insignificant likelihood-ratio (L-R) statistics affirm the null hypothesis that the index is independent from cointegrative relations. Insignificant L-R test statistics are associated with leading countries within the CIV. Corresponding p-values are shown under the L-R test statistics. DF is the degree of freedom, while 5% C.V. represents the critical value at a 5% level. ***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

We contend that cointegrated real estate sectors share common underlying trends and thus co-move temporally, which undermines diversification benefits. However, leading indices within a CIV do not respond to deviations from the cointegrative relation, and although cointegrated, may still offer diversification benefits (Pesaran, Shin, and Smith 2000, Yunus 2009, Gallo, Lockwood, and Zhang 2013). To differentiate leading sectors from subordinate following sectors, we conduct a Granger (1969) causality test.

The Granger-causality test is to identify if one variable improves the forecasting performance of another variable. If the price indices are cointegrated, as suggested by Engle and Granger (1987), an error correction term needs to be imposed in the Granger-causality relation as follows:

$$\Delta X_t = \upsilon + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_k \Delta X_{t-k} + \lambda \vartheta_t + \varepsilon_t, \qquad t=1,\ldots,T \tag{6}$$

where ΔX_t is a (5×1) matrix of the returns of the five cointegrated sectors, υ is a (5×1) vector of the constants, Γ_i is a (5×5) matrix of the beta coefficients, ΔX_{t-i} is a (5×1) matrix of the lagged endogenous return variables, and ε_t is a (5×1) matrix of the white noise error terms. ϑ_t isthe error correction term that measures the response to deviations from the cointegrative equilibrium among the five cointegrated sectors. As mentioned in Granger (1988), the causation of the endogenous variable by the exogenous variables within the ECM can be either through the lagged values of the exogenous variables or the error correction term, ϑ_t. To test if one sector Granger-causes the other in the short-term, we test if the coefficient on the lagged endogenous sector returns are jointly significant as measured by the F-statistic or if the coefficient of the ϑ_t is significant as measured by the T-statistic.

The last column in Table 4 shows that all of the sectors are not affected by the error correction term in the short-term as shown by the insignificant T-stat in each sector. In addition, other than its own lags (F-stat=24.98), HEAL is not significantly Granger-caused by any other cointegrated sector. Similar results can be found on the RETL (F-stat=12.69) and SELF (F-stat=12.35) sectors. INDU, however, is not only Granger caused by its own lags (F-stat=2.89), but also significantly caused by HEAL (F-stat=1.69), RETL (F-stat=2.40) and SELF (F-stat=1.99). LODG is also significantly Granger-caused by its own lags (F-stat=22.68) and RETL (F-stat=1.72). The Granger causality tests suggest the leading role of HEAL, RETL, and SELF and subordinate role of INDU and LODG in the CIV. Since cointegrated leading indices still offer diversification benefits, we conclude that diversified real estate sector portfolios should consist of two segmented sectors (UNCL and RESI) and the three leading cointegrated sectors (HEAL, RETL, and SELF) as the essential real estate sectors.

Table 4 Ganger Causality within the CIV (from 03/1984 to 12/1994)

	HEAL(X)	INDU(X)	LODG(X)	RETL(X)	SELF(X)	ϑ (T-stat)
HEAL(Y)	24.98***	0.85	1.19	1.02	0.92	0.01
INDU(Y)	1.69*	2.89***	0.89	2.40**	1.99*	0.01
LODG(Y)	0.94	1.12	22.68***	1.72*	0.70	-0.02
RETL(Y)	1.06	1.08	1.59	12.69***	0.94	-0.01
SELF(Y)	0.69	0.71	0.90	0.78	12.35***	-0.01

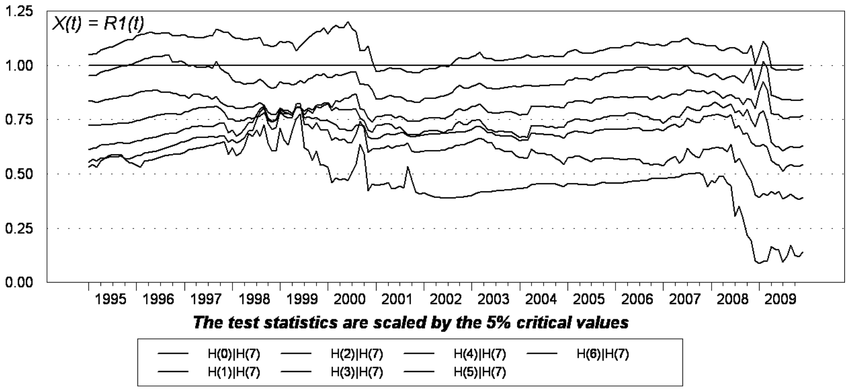
Notes: Table 4 shows Granger (1969) causality F-test statistics with maximum of 12 lags, (X) represents explainable variables; and (Y) represents dependent variables. ϑ is the error correction term from the cointegration test and under a T-test. Significant F-test statistics and/or T-test statistics represent Granger causality.

From the essence of the cointegration theory and ECM model, we know that in general the cointegrative relation is stable over time. To show that our findings of cointegrated/segmented and leading/following sectors in the first window are not time dependent, we present a recursive cointegration analysis in Figure 2. The Hansen and Johansen (1999) recursive stability test of the cointegration parameters is performed. Based on the notion that the parameters should be stable if the model is valid and useful, the test is implemented by assuming that short-term parameters are constantly held at their full-sample estimates, but that the long-term relations are allowed to change over time, then calculating the trace statistic over the fixed base period (03/1984– 12/1994)and increasing one additional observation at each iteration to re-estimate the trace statistic until the last trace statistic is computed over the full sample period. In Figure 2, conditional on the presence of one CIV in the partial VAR, the recursive likelihood ratio trace test statistics (scaled by a 5% critical value) are plotted against the end of each estimation window. The number of lines above the critical value line of 1.00 would indicate the number of CIVs determined at the 5% significance level. [13] Figure 2 shows that one and only one CIV is consistently above the line of 1.00 from 1995 to 2009, which indicates stable cointegrative relations. This shows us that all our subsequent cointegration tests are consistent and unbiased. [14]

[13] Note that the 1.00 line represents a more restrictive 5% significant level. The two short periods: 1) Jan. 2001 to Mar. 2002 and 2) Mar. 2009 to Dec. 2009, although the upmost line is slightly below 1.0, are still significant at the 10% level, so that the cointegration relation is consistent over time and not time dependent.

[14] Although not reported, we conduct a beta constancy test to evaluate whether the cointegration rank is stable over the time period. We find a null hypothesis in that a single CIV which is stable over the sample cannot be rejected, which suggests the CIV is stable over time. This is consistent with our reported recursive stability test.

Figure 2 Trace Test Statistics (01/1995-12/2009)



Notes: Recursive cointegration tests of stability of cointegrating vectors (CIVs) in the group of REIT sector indices: healthcare, industrial/office, residential, lodging/resort, retail, self-storage, and unclassified (recursive trace tests). Figure 2 shows the plots of recursive trace statistics scaled by 5% critical values against time. 03/1984– 12/1994 is the base period. The number of lines above 1.00indicate the number of CIVs determined at the 5% significance level. The two short periods: 1) Jan. 2001 to Mar. 2002 and 2) Mar. 2009 to Dec. 2009, although the upmost line is below1.00, they are still significant at the 10% level.

Next, we conduct analyses on the causes of the cointegrative disequilibrium. As presented in Table 3, there is one CIV among five sector indices. A time series of CIV residuals is computed and saved to proxy for the deviation from the bonded cointegrative relation among these five cointegrated sectors. We are therefore interested in identifying the driving factors that affect the cointegrative disequilibrium.

First, we present a VAR analysis with an interactive dummy with the past CIV residuals. The purpose of this test is to examine the effects of down markets on the cointegrative disequilibrium. The influx of downside risk questions the importance of a domestic real estate sector investment strategy in down markets, precisely the market conditions in which investors most need diversification. Within the cointegrative relation, we record the CIV residuals and lag the residuals by one month to account for the past disequilibria of the CIV, and then regress the returns of each cointegrated sector against the past disequilibria of the CIV, a market condition dummy, an interactive term, and other lagged returns. [15]

[15] We also test the VAR system with up to 12 lagged market returns and these results are consistent with the original model.

Table 5 Cointegration under Different Market Conditions

ΔY_t	ρ_0	ρ_1	ρ_2	ρ_3	$\rho_{HEAL(t-1)}$	$\rho_{INDU\,(t-1)}$	$\rho_{LODG(t-1)}$	$\rho_{RETL(t-1)}$	$\rho_{SELF(t-1)}$
$\Delta Y_{HEAL(t)}$	0.0142	0.0378	-0.0302	-0.0505	-0.1327	-0.1715	-0.0473	0.2137	0.1661
T-stat	0.69	1.70*	-0.99	-2.03**	-1.76*	-2.09**	-1.14	2.24**	2.60***
$\Delta Y_{INDU(t)}$	0.0548	-0.0173	0.0088	-0.0771	-0.1321	0.0788	-0.0682	0.1818	-0.0352
T-stat	2.40**	-0.77	0.26	-2.35**	-1.59	0.87	-1.49	1.73*	-0.50
$\Delta Y_{LODG(t)}$	-0.0156	-0.0254	0.0447	-0.1295	-0.2212	0.4152	0.0272	-0.0472	0.0119
T-stat	-0.45	-1.25	0.87	-2.61***	-1.75*	3.03***	0.39	-0.30	0.11
$\Delta Y_{RETL(t)}$	0.0058	0.0350	-0.0019	-0.0651	-0.1234	-0.0011	-0.0282	0.1599	0.0141
T-stat	0.27	1.70*	-0.06	-2.16**	-1.61	-0.01	-0.67	1.65	0.22
$\Delta Y_{SELF(t)}$	0.0398	0.0031	-0.0423	-0.0160	-0.0180	-0.2104	0.0212	0.2273	-0.0188
T-stat	1.71*	0.13	-1.23	-1.88*	-0.21	-2.28**	0.45	2.13**	-0.26

Notes: This exhibit presents a vector autoregression (VAR) analysis with an interactive dummy with the past CIV residual. The VAR model is as follows:

$$\Delta Y_t = \rho_0 + \rho_1 \vartheta_{t-1} + \rho_2 \varpi_t + \rho_3 (\vartheta_{t-1} * \varpi_t) + \sum_{\eta=1}^{k} \rho_{\eta+3} * \Delta Y_{t-\eta} + \psi_t$$

where ΔY_t is the vector of real estate sector returns in a VAR system at time *t*; ϑ_{t-1} is the CIV residual (disequilibrium) in the past at time *t-1*; ϖ is a binary dummy variable with 1 = declining market and 0 = rising market at time *t*; $(\vartheta_{t-1} * \varpi_t)$ is the interactive term with past disequilibrium and current market condition; $\sum_{\eta=1}^{k} \rho_{\eta+3} * \Delta Y_{t-\eta}$ is the sum of *k* number of one period of lagged returns in the real estate sector in the VAR system and ψ represents the error term of the system. ***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

As presented in Table 5, two (three) out of five of the returns of the real estate sectors significantly (insignificantly) and positively responded to their past disequilibria in up markets, as indicated by the dummy variable ρ_1. In particular, in up markets, HEAL (ρ_1=0.0378, t-stat $_{HEAL,\rho1}$=1.70) and RETL (ρ_1=0.0350, t-stat$_{RETL,\rho1}$=1.70) deviated from their bonded long-term equilibrium. However, all five cointegrated sectors significantly and negatively reacted to their past disequilibria in down markets, captured by the interactive term $\rho_{1+}\rho_3$. For instance, in down markets, HEAL moved towards its long-term equilibrium by ($\rho_{1+}\rho_3$) =0.0378-0.0505=-0.0127 (t-stat$_{HEAL,\rho3}$=-2.03). These findings suggest that cointegrated sectors become more cointegrated in down markets as evidenced by movement towards their long-term equilibrium. We report that the cointegrative relation among cointegrated real estate sectors is in fact induced by down markets. This finding suggests that cointegrated sectors lack diversification benefit during down markets while essential ones should provide better downside risk protection. This implies that we ought to look into different market conditions when examining a cointegration-based portfolio strategy.

Second, we further examine the causes of the cointegrative disequilibrium by running the absolute value of the CIV disequilibria against seven contemporaneous economic factors proposed by Lee and Chiang (2004) who find that industrial production affects both equity and mortgage REIT returns. Besides that, Chan, Hendershott, and Sanders (1990) report that unexpected inflation, changes in the risk premium, and term structures of interest rates consistently drive equity REIT returns. Ling and Naranjo (1997) argue that unexpected inflation and term structures of interest rates also systematically affect real estate returns. Myer, Chaudry, and Webb (1997) also find that a common factor, inflationary expectations, affects property values across countries. Although we do not directly test the determinants of REIT returns, the returns are closely related to the ECM residuals through the REIT prices. Therefore, we expect that certain economic factors have impact on ECM residuals as they affect real estate asset returns.

In Table 6A, we present the description, data source, and expected sign of each of the seven exogenous variables in Equation 5. The data sources are similar to those of Lee and Chiang (2004). Importantly, we list the anticipated impacts of the exogenous variables on the absolute value of the cointegrative disequilibrium. In particular, INDPRO is expected to have a positive sign because a higher industrial production growth rate signals a better market condition in general. According to our finding in Table 5 that real estate sectors tend to be less cointegrated under up market conditions, we expect that a higher industrial production growth rate will mean a larger deviation from their bonded cointegrative relation. By the same token, a higher term structure means a more upward sloping yield curve which signals a better market condition so we anticipate a positive sign for the TS coefficient. Similarly, higher new private housing starts mean better real estate market conditions, so that a positive sign for the HSTARTS coefficient is expected as well. Ling and

Naranjo (1999) and Okunev, Patrick and Zurbruegg (2000) find a strong positive relation between U.S. stock and the real estate market, so we expect a positive impact of the SP500R on the cointegrative disequilibrium.

On the contrary, we expect a negative sign for the INFL, DEF, and MTGED coefficients. With respect to unexpected inflation, we anticipate a negative sign in that high unexpected inflation translates into a pessimistic economic state so that real estate sectors tend to cointegrate more under down market conditions. High default risk premium also signals an unpleasant market condition so we expect a negative sign as well. Similarly, high mortgage rate means high borrowing costs to home buyers which lower real estate market returns. A negative relation is anticipated between mortgage rate and the cointegrative disequilibrium.

In Table 6B, we report the empirical results of Equation 5 and find that unexpected inflation (λ_{INFL}=-1.3890, t-stat=-1.81) significantly and negatively influences the absolute value of the cointegrative disequilibrium. When experiencing higher unexpected inflation, a smaller disequilibrium error indicates a greater cointegrating effect. The economic explanation for the negative coefficient on the unexpected inflation is that with a high-unexpected inflation shock, all of the domestic markets, including the real estate market, would drop due to such unexpected unpleasant news that gives rise to lower real estate market returns. Although the entire real estate market is declining, different sectors might have a different declining pace due to the taste of the investors.[16] Therefore, the five cointegrated sectors become more cointegrated under down real estate markets.

Table 6A Description of Exogenous Variables

Variable	Description	Data Source	Expected Sign
INDPRO	Industrial production growth rate	Federal Reserve Bank of St. Louis	+
INFL	Unexpected inflation rate	Labor Bureaus of Statistics	-
TS	Term structure	Federal Reserve Bank of St. Louis	+
DEF	Default risk premium	Federal Reserve Bank of St. Louis	-
MTGED	Change of 30-Year mortgage rate	Federal Reserve Bank of St. Louis	-
HSTARTS	New private housing starts	U.S. Census Bureau	+
SP500R	S&P 500 index return	Center for Research in Security Prices	+

(Continued…)

[16] We also conduct the same tests under sub-periods and varying market conditions, and we find similar results. Unreported results are available upon request.

(Table 6 Continued)

Table 6B Causes of Cointegrative Disequilibrium

Variable	Coefficient	Robust Standard Error	T- stat
Intercept	0.1286	0.0165	7.81***
INDPRO	-1.2049	0.9260	-1.30
INFL	-1.3890	0.7674	-1.81**
TS	0.9408	0.5897	1.60
DEF	0.0845	1.6469	0.05
MTGED	-0.0813	0.2339	-0.35
HSTARTS	-0.0594	0.0474	-1.25
SP500R	-0.1062	0.1388	-0.76
R-Sqr	0.02		
Mean VIF	1.62		
Durbin-Watson	1.86		

Notes: The absolute value of the cointegration residual $|\vartheta_t|$ as the dependent variable is regressed by using against seven macroeconomic factors proposed in Lee and Chiang (2004):

$$|\vartheta_t| = \lambda_0 + \lambda_1 INDPRO_t + \lambda_2 INFL_t + \lambda_3 TS_t + \lambda_4 DEF_t + \lambda_5 MTGED_t + \lambda_6 HSTARTS_t + \lambda_7 SP500R_t + \tau_{pt},$$

where $INDPRO_t$ is the industrial production growth rate, $INFL_t$ is the unexpected inflation rate which is defined as the change of seasonally adjusted Consumer Price Index, TS_t is the term structure which is defined as the (20-yr)-(1-yr) yield, DEF_t is the risk premium which is defined as the BAA-AAA yield, $MTGED_t$ is the change of the 30-year conventional mortgage rate, $HSTARTS_t$ is the new private housing starts, and SP500R is the S&P 500 index return at time t. All seven factors are on a monthly basis.T-stats are reported beneath each parameter estimate. We conduct the robust regression to calculate the heteroskedastic robust standard errors. We test for multicollinearity by calculating the mean variance inflation factors (VIFs) for the independent variables. We also perform Durbin-Watson testing (d-statistics) of residual autocorrelation.*** and ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Next, based on the above results from the long-term cointegration tests and short-term Granger causality tests, we contend that only five essential sectors: two segmented sectors plus three leading cointegrated sectors, should be included in diversified real estate sector portfolios. The two cointegrated sectors are redundant diversifiers which should be excluded. We form our annually rebalanced portfolios: VWCOI and EWCOI. Particularly, at the beginning of every year over the sample period, the VWCOI is constructed based on the prior December market value of five essential sectors, the portfolio is then held for the next 12 months and rebalanced annually. Table 7A presents the annual portfolio constituents of the VWCOI from 1995 to 2009.

Table 7A Annual VWCOI Portfolio Constituents (in percentage)

Country	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
UNCL	6.08	6.14	6.48	2.86	3.09	7.64	4.61	8.45	7.43	6.67	9.46	9.72	8.68	8.87	10.46
HEAL	14.85	18.20	16.35	14.83	12.67	9.06	7.71	8.69	9.57	9.60	9.36	8.92	10.17	14.16	19.24
RESI	31.07	30.57	32.19	35.91	34.78	36.17	42.30	37.50	33.35	30.46	29.04	28.18	29.16	21.42	23.81
RETL	42.51	38.18	36.13	38.47	41.12	39.56	37.95	36.85	43.15	46.19	44.96	44.80	43.00	46.58	33.03
SELF	5.49	6.91	8.84	7.93	8.34	7.57	7.42	8.51	6.49	7.07	7.19	8.39	8.99	8.97	13.47

Table 7B Annual EWCOI Portfolio Constituents (in percentage)

Country	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
UNCL	20	20	20	20	20	20	20	20	20	20	20	20	20	20	20
HEAL	20	20	20	20	20	20	20	20	20	20	20	20	20	20	20
RESI	20	20	20	20	20	20	20	20	20	20	20	20	20	20	20
RETL	20	20	20	20	20	20	20	20	20	20	20	20	20	20	20
SELF	20	20	20	20	20	20	20	20	20	20	20	20	20	20	20

Notes: Table 7A reports the annual percentage (%) allocations of the value-weighted cointegration-based VWCOI portfolio. At the beginning of every year over the sample period, we construct the VWCOI based on the prior December market value of five essential real estate type REITs, the portfolio is also held for the next 12 months. Table 7B reports the annual percentage (%) allocations of the equally weighted cointegration-based EWCOI portfolio. At the beginning of every year over the sample period, we construct the EWCOI with 20% weight on each sector, the portfolio is also held for the next 12 months.

As presented in Table 7B, at the beginning of every year over the sample period, the EWCOI is constructed based on 20% equal weight on each of the five essential sectors, the portfolio is then held for the next 12 months. The benefits of including leading sectors can be examined by comparing the performance of the COI portfolios to portfolios that consist of only segmented sectors (UNCL and RESI). Although not reported, we find that COI portfolios outperformed the portfolios that contain only two segmented sectors which support the inclusion of leading sectors into sufficiently diversified real estate sector portfolios.

To understand the implications and investment value of the cointegrative relation among real estate sectors, we compare the performance of our COI portfolios with the MKT portfolio over 01/1995 to 12/2009 in Table 8.[17] Table 8A reports the average monthly return, standard deviation, SHP, Jobson-Korkie statistics, and risk premium T-test statistics. The mean monthly return for both the VWCOI (1.02% versus 0.93%, z-stat=1.67) and EWCOI(1.05% versus 0.93%, z-stat=1.73) portfolios significantly exceeded the mean return for the MKT. Annualized returns equaled 12.95% for the VWCOI portfolio, 13.35% for the EWCOI portfolio and 11.75% percent for the MKT respectively. Both VWCOI and EWCOI risk premium were significantly positive ($t\text{-stat}_{VWCOI}$=1.71; $t\text{-stat}_{EWCOI}$=1.87) while the MKT was less attractive with an insignificant risk premium ($t\text{-stat}_{MKT}$=1.44). The Jobson-Korkiez-stat indicates that the SHP is significantly higher for both the VWCOI portfolio (12.72% vs. 10.78%) and EWCOI (13.95% vs. 10.78%) versus the MKT, which is because the COI portfolios had a higher return (RET) as well as lower risk (SD) than those of the MKT benchmark. In sum, Table 8A indicates that domestic real estate investors would prefer the COI portfolios over the all-sector MKT portfolio. These results confirm that the inclusion of two non-leading cointegrated sectors (INDU and LODG) leads to suboptimal asset allocations, thus resulting in inferior portfolio performance.

Tables 8B and 8C present the results of the four-factor model that control for exposures to the broad REIT-based market, size, style, and momentum factors. The results are consistent with the total risk findings reported in Table 8A. Importantly, both the VWCOI (a_{VWCOI}=0.14%, t-stat=1.92) and EWCOI (a_{EWCOI}=0.20%, t-stat=1.85) portfolios exhibit superior significant abnormal returns. Compounded annually, the abnormal returns equal 1.69% for the VWCOI and 2.43% for the EWCOI portfolios, respectively. Not surprisingly,

[17] Similar to the study by Ro and Ziobrowski (2011) which uses data from 1997 to 2006, this study examines the data from 1995 (the beginning of our portfolio formation) to 2009. However, unlike Ro and Ziobrowski (2011) who use equity market factors, MKTRP, SMB, HML and UMD, in their four-factor asset pricing model, we use REIT-based market factors as suggested by recent studies by Hartzell, Mühlhofer and Titman (2010) and Cici, Corgel and Gibson (2011).

Table 8A Summary Statistics (from 01/1995 to12/2009)

Portfolio	RET(%)	SD(%)	SHP(%)	J-K z-stat (vs. MKT)	T-test H0: (RET-R_f) = 0
MKT	0.93	5.90	10.78		1.44
VWCOI	1.02	5.75	12.72	1.67*	1.71*
EWCOI	1.05	5.45	13.95	1.73*	1.87*
3MTB (R_f)	0.38	0.19			

Table 8B VWCOI Performance Summary (from 01/1995 to12/2009)

Portfolio	a	b1	b2	b3	b4	R-Sqr
VWCOI	0.0014	0.9699	0.0329	-0.1581	0.0820	0.97
T-stat	(1.92)*	(67.59)***	(1.02)	(-4.68)***	(0.13)	
$VWCOI_{Up}$	-0.0006	1.0278	0.1025	-0.2159	-0.4759	0.93
T-stat	(-0.36)	(26.61)***	(2.24)**	(-4.49)***	(-0.57)	
$VWCOI_{Down}$	0.0023	0.9733	-0.0313	-0.1204	2.4491	0.97
T-stat	(1.69)*	(56.32)***	(-0.64)	(-1.90)*	(1.78)*	

Table 8C EWCIM Performance Summary (from 01/1995 to12/2009)

Portfolio	a	b1	b2	b3	b4	R-Sqr
EWCOI	0.0020	0.9215	0.0995	-0.1831	0.7479	0.94
T-stat	(1.85)*	(37.32)***	1.63	(-2.67)***	0.69	
$EWCOI_{Up}$	-0.0006	0.9090	0.1364	0.0919	-1.1144	0.83
T-stat	(-0.34)	(15.48)***	(1.20)	(0.92)	(-0.44)	
$EWCOI_{Down}$	0.0039	0.9266	0.0833	-0.2529	0.6356	0.97
T-stat	(3.02)***	(30.83)***	1.04	(-3.55)***	(0.56)	

Notes: Table 8A reports descriptive statistics for the benchmark and portfolios used in the study over the 180-month period of 01/1995-12/2009. RET is the monthly raw return and SD is the standard deviation of the monthly returns, and SHP is the Sharpe ratio. MKT is the value-weighted CRSP/Ziman REIT market index. VWCOI is the value-weighted cointegration-based portfolio with five essential real estate type REITs. EWCOI is the equally-weighted weighted cointegration-based portfolio with five essential real estate type REITs. Table 8B presents the results of a four-factor model regression for VWCOI and Table 8C presents the results of a four-factor model regression for EWCOI:

$$R_t - R_{ft} = a + b_1(ReitR_{MKT} - R_{ft}) + b_2 ReitSMB_t + b_3 ReitHML_t + b_4 ReitUMD_t + e_{pt},$$

where R_{pt} is the monthly portfolio return, R_{ft} is the monthly Citigroup 3-month Treasury bill return, $ReitR_{MKT}$ is the value-weighted CRSP/Ziman REIT market index, ReitSMB (small minus big) is the REIT size factor, ReitHML (high minus low) is the REIT style factor and ReitUMD (up minus down) is the REIT momentum factor. T-stats are reported beneath each parameter estimate.***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

each of the two COI portfolios have a significant beta loading which shows that REIT market premium significantly explains for the return variations of the two portfolios. Notably, both cointegration-based portfolios VWCOI ($b_{VWCOI,1}$=0.9699) and EWCOI ($b_{EWCOI,1}$=0.9215) have a beta smaller than one, a proxy for a less systematic market risk than the all-sector MKT portfolio. Both COI portfolios have an insignificant size loading ($b_{VWCOI,2}$=0.0329, t-stat=1.02; $b_{EWCOI,2}$=0.0995, t-stat=-1.63) and momentum loading ($b_{VWCOI,4}$=0.0820, t-stat=0.13; $b_{EWCOI,4}$=0.7479, t-stat=0.69) which imply that COI portfolios consist of different sized REITs and both winner and loser REITs of last year. Both COI portfolios also have a significant style loading ($b_{VWCOI,3}$=-0.1581, t-stat=-4.68; $b_{EWCOI,3}$=-0.1831, t-stat=-2.67) which implies that these portfolios tend to hold REITs with low Book-to-Market ratios. In sum, any rational investor would prefer VWCOI and EWCOI over the MKT.

From Table 5, we understand that cointegrated real estate sectors become more cointegrated in down market conditions. Segmented real estate sectors are expected to extend better performance (protection) for investors during down markets. To further determine the extent to which general market conditions impact on portfolio performance, we run performance tests separated by up and down market conditions. The results are reported in Table 8B and 8C. Interestingly, no significant abnormal return is evident for both COI portfolios during up markets which implies that these portfolios performed indifferently from the MKT portfolio under good market conditions. This finding also suggests that the overall outperformance of the COI portfolios came from their superior performance under down markets. Importantly, both the VWCOI and EWCOI portfolios significantly outperformed the four-factor model during down markets ($a_{VWCOI,Down}$=0.23%, t-stat=1.69; $a_{EWCOI,Down}$=0.39%, t-stat=3.02). Therefore, the COI portfolios performed in line with the four-factor benchmark in up markets, but significantly outperformed it by 23 basis points per month for the VWCOI and 39 basis points for the EWCOI during down markets, thus effectively providing a downside risk protection. Interestingly, we find that an equal-weighted strategy, more easily implemented, provides better performance than that of a more active value-weighted strategy. The results in Table 8 provide an important finding that the better diversified cointegration-based portfolios, VWCOI and EWCOI, which eliminated redundant diversifiers, provided better protection to investors than an MKT during unfavorable market conditions, in which investors most need diversification benefits.

Finally, similar to Chaudhry, Myer, and Webb (1999) and Yunus, Hansz, and Kennedy (2012), we present the model specification for price discovery and cointegration forecasting purposes based on the entire sample period from 1984 to 2009.

Table 9A Cointegration Rank Tests (from 03/1984 to 12/2009)

I(1)-Analysis (n=5, lag=1)	G(r)	p-r	r	Eig.Value	Trace	Bartlett Trace	P-Value	Bartlett P-Value
	1.50	5	0	0.13	65.73	65.21	0.00***	0.00***
	P-value=0.22	4	1	0.04	23.57	23.42	0.47	0.47

Table 9B Exclusion Tests (from 03/1984 to 12/2009)

REIT Index (n=5)	r	DF	5% C.V.	HEAL	INDU	LODG	RETL	SELF
L-R statistic	1	1	3.84	12.39	28.16	11.89	22.67	8.11
P-value				0.00***	0.00***	0.00***	0.00***	0.00***

Table 9C BETA (transposed)

Log of Price	HEAL	INDU	LODG	RETL	SELF
Beta(1)	-4.954	-8.171	2.246	3.794	5.988
Beta(2)	4.383	0.475	0.903	-0.656	-3.38
Beta(3)	-0.602	2.014	-2.991	-1.243	0.918
Beta(4)	-4.734	-4.208	0.449	7.185	-0.03
Beta(5)	1.911	-1.588	0.804	-2.094	0.173

Notes: In Table 9A, the G(r) statistic is distributed chi-square with r degrees of freedom and used to detect common linear trends among any of the real estate sector indices, p is the number of dimensional vectors of real estate sector indices, r is the number of cointegrated vectors, Eig Value is the eigenvalue obtained from maximum likelihood estimation of the ECM. Trace is Johansen trace statistic for cointegration rank test; Bartlett Trace is the Bartlett-small-sample-corrected trace statistics, P-value is the probability value for Johansen trace statistics. The final column reports the p-value for the Bartlett trace statistic. Table 9B presents the exclusion test results based on the rank tests in Table 9A. Insignificant likelihood-ratio (L-R) statistics affirm the null hypothesis that the index is independent from cointegrating relations. Corresponding p-values are shown under the L-R test statistics. DF is the degree of freedom, while 5% C.V. represents the critical value at a 5% level. Table 9C reports the beta matrices which display the long-term relationship among the five cointegrated sectors in the system. ***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

First, as expected, the results of the cointegration rank testing, as presented in Table 9A, show that a significant CIV exists among the five aforementioned cointegrated sectors(Bartlett λ_{trace}=65.21, p-value=0.00) over the 1984-2009 period. Insignificant G(r) statistics (p-value=0.22) suggests that a linear trend is not present in the cointegrative relation. Table 9B confirms that none of the five cointegrated sectors are independent from their bonded long-term cointegrative relation as suggested by significant L-R statistics. Table 9C lays out the beta matrices that display the long-term relations among the five cointegrated sectors in the system. Since one significant CIV is identified in Table 9A, only the Beta (1) relationship is meaningful. Therefore, the long-term cointegrative relation based on the full sample period (03/1984 to 12/2009) is as follows:

$$- 4.954HEAL_t - 8.171INDU_t + 2.246LODG_t + 3.794RETL_t + 5.988SELF_t = 0 \tag{7}$$

where $HEAL_t$, $INDU_t$, $LODG_t$, $RETL_t$, $SELF_t$, are logs of price at time t. The coefficient of each sector means the extent to which one sector reacts to other sectors in the system. Equation (7) can be normalized with respect to any variable in the model. For instance, the model can be rewritten as $HEAL_t = -1.649INDU_t + 0.453LODG_t + 0.766RETL_t +1.209SELF_t$, so that in the long run, a unit change in the INDU price index is expected to decrease the value of the HEAL index by 1.649 while a unit change in the LODG, RETL, and SELF price index is expected to increase the value of the HEAL index by 0.453, 0.766, and 1.209, respectively. [18] Equation (7) discloses a price discovery among the five cointegrated sectors for developing forecasts of real estate prices.

6. Conclusion

This study extends the work of Chaudhry, Myer, and Webb (1999) by investigating the long-term cointegrative relation among four real estate sectors. Unlike Chaudhry, Myer, and Webb (1999), this study focuses on both the long-term cointegrative and short-term causal relations among seven U.S. real estate sector REITs from March 1984 to December 2009. Real estate sector price indices are found to be non-stationary with unit root. By using cointegration approaches, we find a significant cointegrative relation among five real estate sectors (retail, industrial/office, lodging/resort, healthcare and self storage) and the residential and unclassified sectors are segmented from

[18] Chaudhry, Myer, and Webb (1999) report a positive long-term cointegrative relation between retail and office. In particular, in the long run on a quarterly basis from 1978 to 1996, a unit change in the office price index increases the value of retail by 2.87 (retail = 2.87 office - 4.31 RD office + 1.45 ware). Our study is on a monthly basis from 1984 to 2009 and we find that a unit change in the industrial/office price index increases the value of retail by 2.154 ($RETL_t=1.306HEAL_t+2.154INDU_t-0.592LODG_t-1.578SELF_t$).

the cointegrative space. Among the five cointegrated sectors, the industrial/office and lodging/resort sectors are found to be Granger-caused by the retail, healthcare, and self-storage sectors so that these sectors are considered as the leading indices while the industrial/office and lodging/resort sectors are considered as the followers, which are redundant diversifiers. We provide evidence that it is necessary to include leading cointegrated sectors that sufficiently represent the cointegrating vector, but exclude redundant cointegrated ones in diversified real estate sector portfolios.

Importantly, this study investigates the causes of cointegrative disequilibrium among five cointegrated real estate sectors. Based on our proposed vector autoregressive model, we find that cointegrated sectors have become more cointegrated under down real estate markets. In addition, we conduct a multivariate sensitivity regression model with seven macroeconomic explanatory variables which show that unexpected inflation significantly and negatively influences the cointegrative disequilibrium. These two proposed models will help scholars to study the causes of disequilibrium of a cointegrative relation and can be extended to international securitized markets for future study.

Lastly, we provide evidence that cointegration-based portfolios have higher raw returns as well as smaller standard deviations, as compared to a traditional all-sector benchmark MKT portfolio. We argue that the inferior performance of the MKT stems from two redundant cointegrated sectors: industrial/office and lodging/resort. Moreover, by using a REIT-based four-factor asset-pricing model with different market conditions, the cointergration-based strategy provides good protection to investors under down markets, in which investors most need diversification.

Since many REITs exclusively operate in the U.S. and institutional investment portfolios often have specific U.S. real estate allocation mandates, our results help to clarify real estate portfolio selection and allocation policy important to institutional investors and real estate portfolio managers. The implications of this study are limited to U.S. domiciled investors. It is of interest to consider how these results hold for global real estate sector diversification, especially when emerging real estate markets are considered. A better understanding of the integration- convergence process of the real estate sector could be also beneficial to academicians and practitioners. These proposed extensions remain promising areas of future research.

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