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Long-Term Benefits from Investing in International Securitized Real Estate

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This paper analyzes long- and short-term co-movements between 14 international real estate stock markets based on cointegration and correlation analyses. The results indicate that there exist strong long-term relationships within economic and geographical regions, but less long-run linkages between real estate markets in different continents. Thus, investors would benefit from broadening their investment horizon from their domestic continent to Australia, Europe, and North America. Furthermore, it is shown that within each region, there are one or two key markets that influence neighboring markets, such as Australia in the Asia-Pacific region, the US in the Anglo-Saxon countries, and France and the Netherlands in the European Monetary Union (EMU). Therefore, from an investor's point of view, it is implied that it should be sufficient to only focus on these central markets. With respect to the efficient market hypothesis, the findings by the cointegration analysis further question its validity for securitized real estate markets.

Keywords

Cointegration; Correlation Analysis; Diversification; Securitized Real Estate Markets

1. Introduction

Over the last two decades, real estate has attracted an increasing amount of investors worldwide and has become a fast growing asset class, with securitized real estate in particular. This trend is accompanied by the introduction of real estate investment trust (REIT) legislation in several countries worldwide, such as Belgium (1995), France (2003), Germany (2007), Hong Kong (2003), Italy (2007), Japan (2000), Singapore (1999), and Great Britain (GB) (2007).[1] Other countries, such as the US and Australia, have had this type of legislation or its equivalent for a long time, and represent the leading securitized real estate markets according to the market capitalization that is related to their GDP.

While the stock and bond markets have become more and more integrated over the last decades, benefits from diversification across international stock and bond markets have decreased, both in the long- and short-term. These stronger linkages between international stock markets have prompted investors to search for different opportunities to diversify their portfolio. Aside from investments in raw materials, such as oil, and precious and industrial metals, real estate investments exhibit a low correlation with stocks and bonds, and therefore, have appropriate characteristics that contribute to portfolio optimization. A summary of former research on the benefits of investing in real estate is presented by Sirmans and Worzala (2003) and Worzala and Sirmans (2003). Further analyses have been conducted by Steinert and Crowe (2001), Conover et al. (2002), Bond et al. (2003), Brounen and Eichholtz (2003), Lee (2005), Lee and Stevenson (2005), Waggle and Agrrawal (2006), Cheng and Roulac (2007), Idzorek et al. (2007), Jin et al. (2007), Fugazza et al. (2008), Yat-Hung et al. (2008), and Sebastian and Sturm (2009). However, the vast majority of research on mixed-asset-portfolio analyses is concentrated more on the characteristics of real estate as an asset class and less on the linkages between national real estate markets and the optimal composition of a real estate portfolio.

In the relevant literature, it is well documented that available diversification benefits are eliminated because asset allocation is home biased by investors. This argument is even more relevant for real estate investments since property companies mainly operate in their own domestic markets. Since these companies are exposed to domestic economic and political shocks, thus, their business is more influenced by local shocks than the business of internationally operating companies in other sectors, e.g. automobile or pharmaceuticals. On the other hand, these considerations raise the question of how investors in domestic real estate can benefit from broadening their investment horizon to neighboring markets and other continents. Second, international investors are interested in the opportunities offered by the long-

[1] See EPRA (2008) and Ooi et al. (2006).

and short-term co-movements between their domestic real estate markets and the foreign ones. These two major concerns present the main points of this study. In previous research, the main focus with regards to benefits from diversification across real estate markets mainly rests on the US market and is based on different types of real estate (e.g. office, residential, industrial, and retail) and geographical regions within the US Even though European and Asia-Pacific real estate markets have experienced fast growth and the number of listed property companies has increased rapidly in the last decade, there have been very few studies which show their contribution to diversification benefits. From the perspective of international investors, investment opportunities in Europe and the Asia-Pacific real estate markets in particular, have dramatically increased. Furthermore, the institutional framework supports this tendency with fewer trade barriers, open markets, and by introducing the REIT legislation according to the US REIT framework. Bardhan et al. (2008) explicitly analyze the impact of a country's economic openness on returns of publicly traded real estate companies. They conclude that economic openness has a significant impact on excess returns of a country's real estate firms and find a negative relationship between excess returns and openness, as well as between excess returns and international financial integration. Furthermore, Liow and Webb (2009) investigate the presence of common factors in the securitized real estate markets of Hong Kong, Singapore, GB, and the US, and show that the linkages across these four markets are much weaker than the linkages across these four economies. This finding suggests that there are pervasive benefits from international diversification across publicly traded real estate markets. Focusing on direct real estate markets and applying a correlation analysis, Lim et al. (2008) find mixed results related to the benefits from international diversification and emphasize that the benefits highly depend on the selected national markets.

In the relevant literature, the primary examinations of benefits from diversification and portfolio optimization are based on correlation analyses. However, this concept is associated with some crucial points which results in strong equivocality as to its meaning. First, from a technical point of view, the returns have to be normally distributed when applying correlation analyses and portfolio optimization based on the mean-variance-approach by Markowitz (1952). However, as shown by Brounen et al. (2008), Liow and Sim (2006), and Liow (2007), this assumption does not hold for real estate returns. Thus, the concept of portfolio optimization based on the first two moments of a return distribution is not sufficient and preferences of investors towards skewness and kurtosis have to be taken into consideration or a different concept must be applied. Second, correlation coefficients capture only the short-term dependence between asset returns, even though investors are usually interested in long-term interrelation and linkages between prices, which are the focus of cointegration analyses. Third, correlation analyses are combined with a loss of valuable information contained in a time series, since correlation coefficients have to be based on stationary variables and price indices are generally not stationary. Hence, first differences or logarithmic returns, respectively, have to be used together with information on the level of the price series, as this is valuable information for long-term oriented investors. Thus, it is more appropriate to investigate the cointegration of prices rather than the correlation of returns with regard to the long-term oriented investor.

Due to these shortcomings of correlation analyses, this paper concentrates on the long-term benefits from diversification across international real estate markets by applying a cointegration methodology as suggested by Engle and Granger (1987) and thus, contributes further evidence on the long-run co-movements between international real estate stock markets to the existing literature. The implications of a cointegration analysis on portfolio diversification depend on the type of investor assumed. Long-run oriented investors with a passive investment strategy realize their highest utility by diversifying across non-cointegrated markets as these markets share no common price trend and do not have a significant linkage between each other. Contrary to this investor type, investors who follow an active investment approach focus on cointegrated markets and the modelling of a short-term error correction model (ECM), to exploit these adjustment processes for additional return. Thus, the concept of cointegration is relevant for different types of investors. Two recent studies on cointegration between international real estate stock markets have been conducted by Yunus (2009), and Gallo and Zhang (2009). However, both studies focus on a time period before the outbreak of the still ongoing financial crisis. When comparing correlation analyses and cointegration methodologies, it is worth emphasizing that these two concepts are not mutually exclusive, but complementary and supportive of each other.

The remainder of the paper is laid out as follows. Section 2 briefly discusses the methodology of testing long- and short-term real estate market interdependence. After discussing the data, the empirical findings are presented in Section 4 (bivariate analysis) and Section 5 (multivariate analysis), while Section 6 summarizes the central results and draws some concluding remarks.

2. Econometric Methodology

A two-stage cointegration methodology presented by Engle and Granger (1987) is employed in the analysis of long-term co-movements between international real estate stock markets. An analysis of each individual long-term relationship between the two markets not only enables us to draw some conclusions on building up real estate portfolios, but also retains the analogy with the concept of bivariate correlation coefficients. As a robustness check on the results from the bivariate Engle-Granger cointegration methodology, as well as to analyze inter-continental and multivariate cointegration relationships, the cointegration test developed by Johansen (1988) is applied.

2.1 Augmented Dickey-Fuller Tests

In the first step, the order of integration of each time series Y is tested before applying the cointegration analysis, or in other words, it is necessary to test whether each time series requires the same degree of differencing to achieve stationarity. In this paper, the order of integration of a time series is determined by applying different approaches of the augmented Dickey-Fuller (ADF) t-tests (Dickey and Fuller (1981) and Said and Dickey (1984)). The ADF values are calculated by estimating regression equations for a random walk, a random walk with drift, and a random walk with drift and trend, respectively:

$$\Delta Y_t = \gamma Y_{t-1} + \sum_{i=1}^{p-1} \beta_i \Delta Y_{t-i} + \varepsilon_t , \tag{1}$$

$$\Delta Y_t = \mu + \gamma Y_{t-1} + \sum_{i=1}^{p-1} \beta_i \Delta Y_{t-i} + \varepsilon_t , \tag{2}$$

$$\Delta Y_t = \mu + \lambda t + \gamma Y_{t-1} + \sum_{i=1}^{p-1} \beta_i \Delta Y_{t-i} + \varepsilon_t , \tag{3}$$

where μ and λ are coefficients of the constant, and the time trend, respectively, β_i are coefficients of the ith order lagged differenced series ($\Delta\ Y_{t-i}$) and the error term ε_t ~ i.i.d. $(0,\sigma^2)$. If γ is equal to zero, the time series Y_t is said to have a unit root and is nonstationary, whereas the time series of $\Delta\ Y_t$ is stationary, I(1). The time series Y_t is stationary and integrated of order zero, I (0), if the null-hypothesis, in which γ equals zero, is rejected.

In contrast to the Dickey-Fuller test (Dickey and Fuller (1979)), the ADF test solves the problem of autocorrelation in the residuals by incorporating a sufficient number of lagged changes of the dependent variable in a regression equation.

2.2 Engle-Granger Test for Cointegration

While the concept of correlation refers to the co-movement in asset returns, cointegration is related to asset prices and their linkages. Two time series are said to be cointegrated and are characterized by mean-reversion if they share a common stochastic trend. The procedure by Engle and Granger (1987), which tests the null hypothesis of no cointegration against the alternative of cointegration, consists of two steps. First, the two nonstationary time series Y_{1t} and Y_{2t} are regressed on each other to obtain the residuals from an ordinary least squares (OLS) regression:

$$Y_{2t} = \alpha + \beta\ Y_{1t} + \varepsilon_t \tag{4}$$

In the second step, these residuals ε_t are tested for unit root characteristics by again, employing the ADF test. Since the residuals do not consist of observed values, but are estimated from the OLS regression, the estimated critical values K for the test statistic according to MacKinnon (1991) are applied:

$$K = \beta_\infty + \beta_1 Z^{-1} + \beta_2 Z^{-2} \tag{5}$$

where Z denotes the sample size and *βs* are the parameters to be estimated and tabulated in MacKinnon (1991), depending on the level of significance and the ADF test specification.

Technically, the two time series are said to be cointegrated if they are integrated of the same order and the residuals from the OLS regression are stationary in levels and integrated of the order zero, respectively.

2.3 Error Correction Model

Furthermore, if two time series share a common stochastic trend and are said to be cointegrated, an ECM can be estimated (Granger representation theorem) and specified, which delivers further insight into the linkage between the two time series and their co-movement over time. Since the estimation is based on a stationary time series, the logarithmic return series is used:

$$\Delta Y_{1t} = \gamma_1 + \lambda_1 \cdot \varepsilon_{t-1} + \sum_{i=1}^{m} \alpha_{11}(i) \cdot \Delta Y_{1t-i} + \sum_{j=1}^{n} \alpha_{12}(j) \cdot \Delta Y_{2t-j} + u_{1t}, \tag{6}$$

$$\Delta Y_{2t} = \gamma_2 + \lambda_2 \cdot \varepsilon_{t-1} + \sum_{i=1}^{m} \alpha_{21}(i) \cdot \Delta Y_{1t-i} + \sum_{j=1}^{n} \alpha_{22}(j) \cdot \Delta Y_{2t-j} + u_{2t}, \tag{7}$$

where γ_1 and γ_2 are coefficients of the constant, $\varepsilon_{t-1} = Y_{2t-1} - \alpha - \beta Y_{1t-1}$ from Equation (4), and α_{11}, α_{12}, α_{21}, and α_{22} represent the coefficients that measure the impact of the lagged returns on the current return of series Y_{1t} and Y_{2t}, respectively. The coefficients λ_1 and λ_2 mainly describe the error correction process.

By implementing lagged returns in the ECM, which is also estimated by using an OLS regression, the short-term relationship and linkages between time series are detected (e.g. analyzing whether the lagged returns from series Y_{1t} influence the returns of series Y_{2t} and/or vice versa). Furthermore, by adding the stationary residuals from the cointegration equation, the adjustment process to the common stochastic trend is analyzed. While ε_{t-1} indicates how far the system has drifted apart from the common long-term path of equilibrium, the sign and the magnitude of the coefficients λ_1 and λ_2 from the regression indicate which time series adjusts to the common trend and how fast the adjustment process takes place. If $\lambda_1 > 0$ ($\lambda_2 < 0$) and is significant, then a deviation from the common stochastic trend is at least partially corrected by the series Y_{1t} (Y_{2t}). A higher absolute coefficient value means that the adjustment process takes place faster.

2.4 Johansen Methodology for Testing Multivariate Cointegration

In the multivariate case, nonstationary time series are said to be cointegrated, and thus share a common long-run relationship, if they are integrated of the same order, I(1), and if their r linear combinations are stationary. To test for the existence of long-run equilibrium relationships among the nonstationary indices of the real estate security prices, we use the maximum-likelihood-based testing procedure suggested by Johansen (1988) and Johansen and Juselius (1990). This methodology is briefly described below.

The analysis starts by formulating an n variable vector autoregression(VAR) process with the lag length k given by:

$$x_t = A_1 x_{t-1} + \ldots + A_k x_{t-k} + \mu + \varepsilon_t, \tag{8}$$

where x_t is an n-dimensional vector of real estate stock price indices, A_i is n x n coefficient matrix, and the matrix μ contains all the deterministic components. The white noise error term is defined by ε_t. It is well documented in the literature on the topic of cointegration that the results of the Johansen test procedure are sensitive to the selection of the lag length k (Boswijk and Frances (1992), Cheung and Lai (1993)). Although there are several different procedures to compute k, we have determined the optimal lag length in the VAR system by using the Akaike information criterion (AIC).

By first differencing Equation (8), the VAR can be transformed into an ECM:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu + \varepsilon_t \tag{9}$$

where $\Gamma_i = -(I - A_1 - \ldots - A_i)$ with $i = 1, \ldots, k-1)$

and $\Pi = -(I - A_1 - \ldots - A_k)$.

While the n x n coefficient matrix Γ_i represents the short-run dynamics, the n x n coefficient matrix Π contains information about the long-run relationships between the variables and its rank r determines the number of cointegration vectors. However, there are three different possibilities:

(i) matrix Π has full rank which means that $r = n$ and indicates that the vector x_t is stationary. Thus, cointegration is not defined and standard VAR in levels can be applied,

(ii) matrix Π is the null matrix r which means that $n - r = n$ and indicates that Equation (2) corresponds to a traditional differenced vector time series model, and

(iii) matrix Π is of a reduced rank r which means that $0 < r < n$ and indicates that there exist r linear combinations of x_t that are stationary or cointegrated. Thus, although x_t itself is non-stationary, the cointegration vectors β have a property in which $\beta' x_t$ is stationary. If this is the case, matrix Π can be decomposed into n x r matrices such that $\Pi = \alpha\beta'$. While α is the matrix of the error correction coefficients that measures the average speed of adjustment towards the cointegrating relationship, matrix β describes the matrix of the cointegration vectors.

The cointegration rank r of matrix Π or the number of common stochastic trends in a multivariate system of nonstationary variables is determined by two tests: the trace and the maximum eigenvalue tests. Both tests examine the number of eigenvalues that are significantly different from zero. Based on the results of Monte Carlo simulations, Cheung and Lai (1995) suggest that the trace test is more robust to skewness and excess kurtosis in the residuals than the maximum eigenvalue test. Therefore, in the empirical section, we have relied on the trace test to determine the cointegration rank. The test is based on the null hypothesis of r cointegration relationships against the alternative hypothesis of n cointegration relationships. The test statistic is given as follows:

$$\lambda_{\text{trace}}(r) = -T \sum_{i=r+1}^{n} \log(1 - \hat{\lambda}_i), \tag{10}$$

where $r = 0, 1, 2, \ldots, n-2, n-1$; $\hat{\lambda}_i$ represents the estimated ith eigenvalue from the eigenvalue problem:

$$\left| \lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k} \right| = 0. \tag{11}$$

The critical values for the trace statistic have been tabulated by MacKinnon et al. (1999). It is well known and documented in the literature that the asymptotic distribution of $\lambda_{\text{trace}}(r)$ and thus the number of identified cointegration vectors is heavily dependent on the specification of the deterministic components of the VAR (Maddala and Kim (1998), Juselius (2007)). To identify the deterministic components of the model, we rely on the selection approach put forth by Juselius (2007).

Following the maximum likelihood estimation technique and identifying the cointegration vector(s), exclusion tests and tests of weak exogeneity are conducted to analyze the significance of each real estate stock market in the cointegration relationship and weak exogeneity of each market. While the latter hypothesis can be tested by setting the relevant row of matrix α to zero, the exclusion test from the cointegration relationship is conducted by restricting the corresponding row of matrix β to zero.

3. Data and Descriptive Statistics

The empirical analysis in this paper is based on the monthly total return indices from the European Public Real Estate Association (EPRA) and the National Association of Real Estate Investment Trusts (NAREIT) between January 1990 to December 2008. The study covers the following 14 national real estate stock markets: Australia (AU), Belgium (BE), Canada (CA), France (FR), Germany (DE), Great Britain (GB), Hong Kong (HK), Italy (IT), Japan (JP), the Netherlands (NL), Singapore (SG), Sweden (SE), Switzerland (CH), and the United States (US). The time series contains 228 monthly data for each market. However, due to the lack of data, the analysis of the Canadian market is only based on 144 monthly returns between 1997 and 2008. To our knowledge, it is the most comprehensive analysis of international cointegration in securitized real estate markets. Sample statistics are calculated in market values based on local currency to focus on real estate factors and avoid distortions caused by changes in exchange rates. The real estate indices are calculated in natural logarithms, whereas the monthly rates of return are calculated based on the first differences of the logarithmic monthly index levels. The national real estate indices are delivered by the same index provider (EPRA/NAREIT) with respect to potential differences between index construction and criteria when using different index providers. The time span from 1990 to 2008 is used given the availability of data.

Figures 1 and 2 present the logarithms of the level of the indices. The Anglo-Saxon real estate markets (AU, CA, GB, and the US) show a continuous upward trend from the beginning of the 1990s until mid 2007 as depicted in Figure 1. In contrast, the Asian markets are characterized by a much more volatile performance, but they seem to have a common trend and move together, which support the notion of applying a cointegration analysis. The performance of the continental European real estate markets is mixed as well. While the markets moved within a certain range in the 1990s with the exception of the small Swedish market, this pattern changed in the second half of the period investigated. From a graphical point of view, the markets can be divided into two groups with one outlier (DE). First, the Belgian and Swiss market moved close in line with each other, where the economic motivation is not obvious. The second group consisted of the securitized real estate markets in FR, IT, NL, and SE. These markets show a strong common upward trend until the first half of 2007 followed by a downward movement in the second half of 2007 and 2008. These markets are members of the European Union and are subject to the monetary policy of the European Central Bank with the exception of SE. Due to the performance of the real estate markets, a different story applies to the German market, which is characterized by high volatility and poor performance. Out of all the European markets, the German market suffered the most from the burst of the high-tech-bubble at the beginning of the 21st century. Afterwards, it was accompanied by a huge upward movement until January 2007, followed by a period dominated by a downward movement, which suggests a close link to the common stock markets. The reasons are manifold. First, the German securitized real estate stock market is small compared to the Dutch and French ones and might thus be more closely related to influences from the common stock market. Second, only a few listed property companies dominate the market. Therefore, company specific events have greater consequences for index changes and trends. Lastly, the German direct (residential) real estate market did not take part in the tremendous international growth and appreciation from the last decade like the markets in Ireland, Spain, GB, and the US This last point is relevant, as real estate companies mainly invest in their domestic market and less in foreign markets. Thus, their performance is highly related to the performance of the national real estate market in the long-run (Fuess et al. (2008)).

From Figures 1 and 2, it is also evident that the Asian markets follow a common downward trend in the aftermath of the Asian and Russian crises in 1997 and 1998, which was more extensive than those of the non-Asian markets. A more common development in the international real estate stock markets is shown in the aftermath of the turmoil in the international financial markets starting in June 2007, when Bear Stearns announced serious problems with regards to their hedge funds. Thirteen out of the 14 securitized real estate markets recorded their highest index level between December 2006 and June 2007, while the markets in GB and the US reached their turning point prior to the other markets. This finding is the first indication for the potential leading function of these two markets, with the US market in particular. The exception of this trend is the market in HK, which did not reach its highest index level until November 2007.

Table 1 provides an overview of the return and risk characteristics of the 14 national real estate stock indices for the period under consideration. As evident, the performance of the securitized real estate markets is very heterogeneous and differs substantially between national markets. While the US has an average monthly return of 0.90 %, and AU, CA, FR, and HK around 0.75 %, respectively, the Japanese and the Swedish markets only have a slightly negative monthly average return of around -0.24 % and -0.05 %, respectively. The three countries with the highest average returns; AU, CA, and the US, have the longest history in REIT legislation. Furthermore, with the exception of HK, the well-performing countries are characterized by relatively low standard deviations which result in the highest Sharpe ratios for the real estate markets in AU, CA, FR, and the US However, there is one important point in defence of the high volatility in the Asian markets. The Asian securitized real estate markets are dominated by property developers and construction activities. Therefore, the cash flows of their business and consequently the equity returns are more volatile in contrast to REITs and other property companies which are dominated by rental investments (Newell and Chau (1996), Liow (1997).

Figure 1 Price Series of the Non-Continental European Country Indices

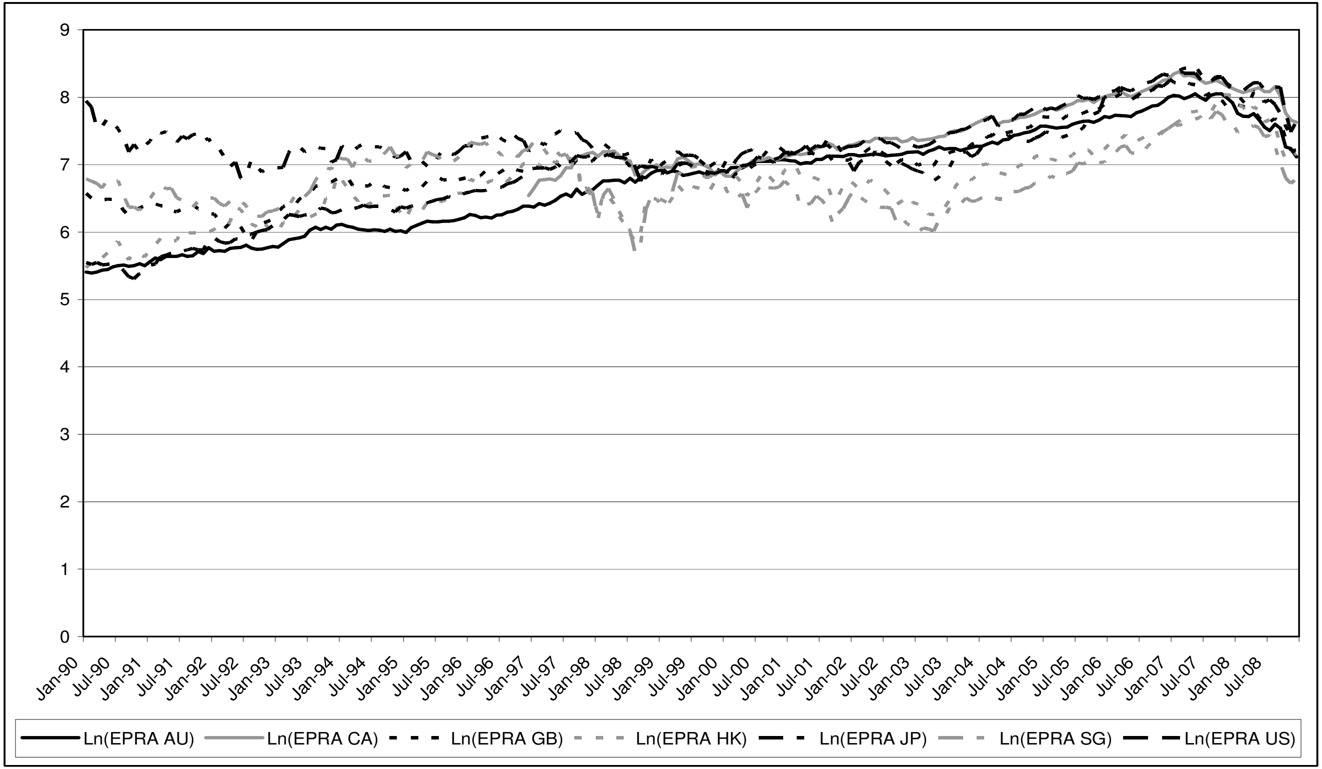


Figure 2 Price Series of the Continental European Country Indices

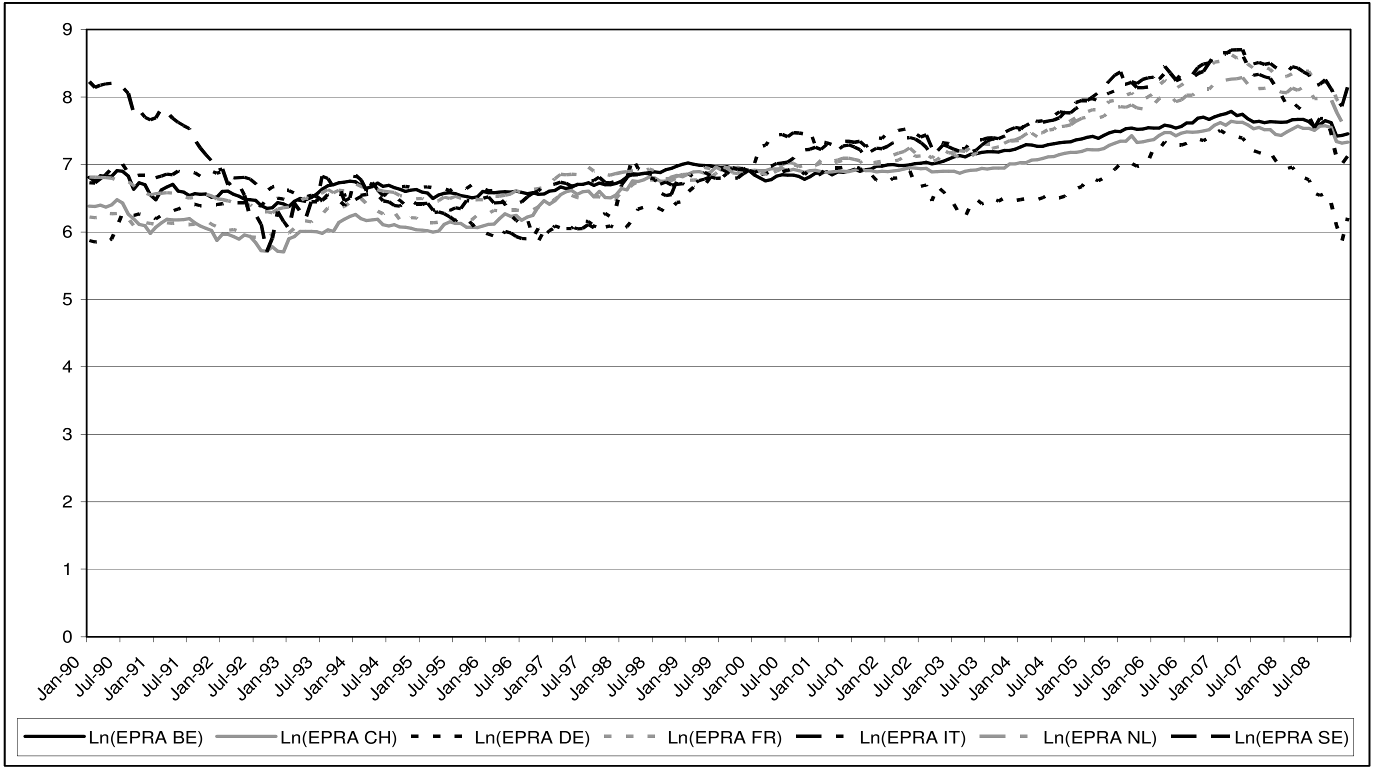


Table 1 Descriptive Statistics of the EPRA Country Indices

Index	Mean	Min.	Max.	S.D.	Skewness (z-stat.)	Kurtosis (z-stat.)	J.-B.
AU	0.0076	-0.2751	0.1060	0.0414	-1.8048(11.2724)	12.7786(31.2278)	1,032.1810***
BE	0.0026	-0.1985	0.1464	0.0408	-1.0535(6.5799)	8.0235(16.0829)	281.9085***
CA	0.0076	-0.2525	0.1685	0.0543	-1.1289(5.6463)	7.5947(11.9599)	157.2501***
CH	0.0040	-0.2112	0.1951	0.0482	-0.3264(2.0387)	6.1851(10.2278)	100.4243***
DE	0.0013	-0.3451	0.3452	0.0751	-0.2665(1.6647)	7.7571(15.2345)	217.6834***
FR	0.0074	-0.2308	0.1298	0.0468	-0.8630(5.3900)	5.7446(8.8247)	99.8594***
GB	0.0025	-0.2517	0.1498	0.0565	-0.6440(4.0222)	4.3218(4.2933)	32.3567***
HK	0.0073	-0.4406	0.4498	0.1023	-0.0439(0.2745)	6.0533(9.8082)	88.6413***
IT	0.0018	-0.3712	0.3420	0.0807	-0.0287(0.1793)	6.7254(11.9487)	131.8795***
JP	-0.0024	-0.3174	0.2299	0.0891	-0.1737(1.0849)	3.3727(1.2706)	2.4664
NL	0.0036	-0.1808	0.0967	0.0405	-0.7516(4.6941)	5.3142(7.4542)	72.3430***
SE	-0.0005	-0.4064	0.3978	0.0964	0.2732(1.7062)	7.9066(15.7105)	231.5416***
SG	0.0007	-0.3899	0.4844	0.1090	-0.2339(1.4611)	5.8490(9.1572)	79.1868***
US	0.0090	-0.3886	0.1581	0.0548	-2.3493(14.6732)	17.1341(45.0995)	2,107.5510***

Notes:
Min. and Max. are the minimum and maximum monthly returns, whereas S.D. is the standard deviation of the return distribution of the national real estate stock indices. ***, ** and * indicate the rejection of the null hypothesis of the Jarque-Bera test statistic (J.-B.) for normality at the 1 %-, 5 %-, and 10 %-levels of significance, respectively. The test results of statistical significance from zero, for skewness coefficients, and from three, for the kurtosis coefficients, are reported in parentheses. The critical values for the coefficient test at the 1 %-, 5 %-, and 10 %-levels of significance are 2.58, 1.96, and 1.65, respectively.

However, it is reasonable to evaluate the attractiveness of the markets by their Sharpe ratios solely based on the first and second moments of the return distribution, when the observed returns are normally distributed or the utility functions of investors are quadratic. However, according to the test statistics of the Jarque-Bera normality test, the null hypothesis of normally distributed returns is rejected for 13 out of 14 national indices at the 1%-level of significance. This is in line with the findings from previous research conducted by Brounen et al. (2008), Liow (2007), Liow and Sim (2006), and Schindler (2009), among others. Only the Japanese real estate market has normally distributed returns. The third and fourth moments emphasize these findings. With the exception of the Japanese market, the return distributions are leptokurtic and negative skewness dominates. Due to the results above, the use of standard deviation as a measure of risk may result in distortions of the true performance. The z-values, in parentheses in Table 1, specify whether the deviation from normality is attributed to the third and/or the fourth moment of the return distribution. Using the testing method suggested by Urzúa (1996), the findings indicate that for AU, BE, CA, FR, GB, NL, and the US, both of the higher moments are responsible for the significant non-normality. For all other non-normally distributed indices, kurtosis alone determines the rejection of normality. Thus, low correlation coefficients can support pervasive diversification benefits, but portfolio optimization and investment decisions that are based on them have restricted relevance. Furthermore, the findings show that non-normally distributed returns are not only characteristic for low-capitalized and developing securitized real estate markets, such as the Belgian, German or Italian markets, but also for high-capitalized markets with a long history, such as the Anglo-Saxon markets, where the Australian and US markets show extremely high negative skewness and leptokurtosis.

4. Empirical Results from a Bivariate Analysis

The presentation of the empirical findings is divided into two parts. First, the correlation structure is considered despite its mentioned limitations. In the second part, the examination focuses on the long-term relationships between the securitized real estate markets and their implications for diversification and investment opportunities for investors.

4.1 Correlation Analysis

The return correlation coefficients between the 14 securitized real estate market indices are displayed in Table 2. All correlation coefficients are positive and range between 0.16 (HK/IT) and 0.55 (AU/CA), with four exceptions. Thus, they are very low (compared to correlations between common stock markets) and indicate pervasive benefits from diversification across national borders and continents, even if the correlation coefficients have increased over the past two years as a consequence of the international financial crisis. The relatively low correlation between real estate markets could be due to the national orientation of the vast majority of publicly listed property companies. Thus, this sector is not subjected to the global events and shocks as much as banking companies or mainly exporting firms. From the perspective of an investor, the focus on international diversification is even more important when investing in real estate stocks than broad stock markets.

The highest dependencies exist between markets strongly connected both geographically and economically. In the Asia-Pacific region, these are the real estate markets in HK and SG with the highest pairwise correlation coefficient of 0.73. The two well-integrated markets in North America show the second highest correlation of 0.69, whereas the largest securitized real estate markets in Europe show a relatively high correlation of 0.66 and 0.58, between the two continental European markets of FR and NL, and FR and GB, respectively. The lowest correlation can be found when considering the relationship of real estate markets located in different continents. Therefore, the findings suggest that real estate investors gain more benefits from diversification when branching out across continents rather than across markets within one continent.

However, the validity of the suggested results has certain limitations. Since a correlation analysis is only valid for stationary variables, the prices have to be de-trended by calculating first differences. However, this procedure causes valuable information with regards to the detection of common trends in prices to vanish. While correlation is an appropriate and highly used measure of short-term co-movements, low correlation coefficients do not assure that there are also low long-term co-movements and vice versa. Thus, further examinations in this paper focus on long-term linkages between the price series of the 14 real estate indices and the dynamic interactions between these markets.

4.2 Unit Root Test of Prices and Returns

As described above, stationarity tests are conducted by applying an ADF unit root test to levels and first differences. ADF values are calculated by estimating regression equations of three types of specification: a random walk (ADF), a random walk with drift (ADF_C), and a random walk with drift and trend (ADF_T), respectively. The relevant literature suggests different procedures to determine the lag length and the ADF test.

Table 2 Return Correlation Coefficients between the EPRA Country Indices

	AU	BE	CA	CH	DE	FR	GB	HK	IT	JP	NL	SE	SG	US
AU	1.0000													
BE	0.3722	1.0000												
CA	0.5548	0.4684	1.0000											
CH	0.2661	0.4057	0.4029	1.0000										
DE	0.2206	0.2730	0.3176	0.2453	1.0000									
FR	0.4264	0.4997	0.5027	0.4655	0.4218	1.0000								
GB	0.3465	0.4106	0.4888	0.4053	0.3853	0.5821	1.0000							
HK	0.2881	0.2080	0.3477	0.2524	0.2058	0.2527	0.2814	1.0000						
IT	0.3950	0.3204	0.4326	0.4180	0.4556	0.4639	0.3867	0.1567	1.0000					
JP	0.2363	0.1640	0.4134	0.1771	0.1447	0.3512	0.2385	0.1658	0.2188	1.0000				
NL	0.4406	0.4432	0.5002	0.4609	0.4283	0.6579	0.5021	0.2862	0.4487	0.3153	1.0000			
SE	0.2604	0.4305	0.3979	0.2387	0.2190	0.4857	0.3688	0.2078	0.3028	0.2566	0.3983	1.0000		
SG	0.2939	0.2697	0.4387	0.2410	0.2612	0.3113	0.3193	0.7287	0.2654	0.2330	0.3658	0.2623	1.0000	
US	0.4315	0.3302	0.6860	0.3147	0.3684	0.4580	0.5024	0.2777	0.3728	0.2880	0.4775	0.2754	0.3320	1.0000

Table 3 Unit Root Test of Prices and Returns

Indices	Unit Root Test in ln (prices)			Unit Root Test in Δ ln (prices)			Integration Level
	ADF_T	ADF_C	ADF	ADF_T	ADF_C	ADF	
AU		-1.7704 (10)				-2.7579*** (4)	I (1)
BE	-2.7454 (10)					-3.4940*** (9)	I (1)
CA			0.3144 (7)			-3.1270*** (6)	I (1)
CH			1.3795 (10)			-3.9615*** (9)	I (1)
DE		-2.5159 (6)				-12.7766*** (0)	I (1)
FR	-2.7378 (6)					-2,8584*** (10)	I (1)
GB		-1.6022 (5)				-4.2536*** (4)	I (1)
HK	-3.0220 (7)					-4.0434*** (10)	I (1)
IT			-0.0967 (9)			-3.5122*** (8)	I (1)
JP	-2.9232 (0)					-14.3145*** (0)	I (1)
NL	-2.6736 (5)					-4.6103*** (4)	I (1)
SE	3.5185** (3)			-7.2598*** (2)			I (0)
SG		-2.7960* (8)				-4.5023*** (10)	I (1)
US		-1.7722 (2)				-2.2300** (10)	I (1)

Notes: ***, ** and * indicate the rejection of the null hypothesis of a unit root at the 1 %-, 5 %-, and 10 %-levels of significance. The lag lengths for unit root tests of prices and returns are given in parentheses.

In principle, there exist two ways to determine the adequate lag length. In one procedure, the optimal lag length is found by successively adding an additional lag until a significant lag is found. Monte Carlo studies have shown that this procedure is biased in its specification selection. Alternatively, the determination process can be started with a relatively long lag length and the model is pared down until a significant lag is identified as proposed by Ng and Perron (1995), and Enders (2004). In this study, the latter approach is used by starting with a lag length of 10 as the initial value. If the t-statistics are insignificant for all lags at the 10 percent level of significance, the equations are re-estimated and the results are tested at the 20 percent level. The correct ADF test is chosen by minimizing the AIC or the Schwarz criterion. Additionally, the testing procedure by Phillips and Perron (1988) is conducted, which confirms the stationarity of the first differences of logarithmic prices in 13 out of the 14 indices.

As displayed in Table 3,the findings of the unit root tests are consistent for all 14 real estate indices with the exception of the Swedish real estate stock market. The null hypothesis of a unit root cannot be rejected for the logarithmic prices. Thus,the indices are not I(0) at the minimum significance level of at least 5 percent and not stationary in levels, respectively.However,the first differences do not exhibit a unit root at the 5 percent level and are stationary. The preferred specification of the ADF test is the model without a constant and trend. A different picture emerges for the Swedish securitized real estate market index. The ADF test in logarithmic prices rejects the null hypothesis of a unit root at the 5 percent level of significance and thus, indicates stationarity in levels and not in first differences. Hence, since the degree of integration differs from the other national real estate indices, the Swedish market is excluded from further examination and also from the cointegration analysis.

4.3 Unit Root Test for Cointegration Residuals

Following the results of the unit root tests, all securitized real estate markets – with the exception of the Swedish real estate market –are integrated of the same order, which is essential for estimating the cointegration vectors. As described above, the first step of the pairwise cointegration test proposed by Engle and Granger (1987) implies the estimation of the OLS regression of logarithmic real estate market indices. In the second step of the two-stage procedure, the residuals from the OLS regression are subjected to the unit root test. From a theoretical point of view, there should not be any differences in the testimony on cointegration when Y_{2t} is regressed on Y_{1t} instead of the regression of Y_{1t} on Y_{2t}. However, it is documented in the relevant literature that differences emerge when using empirical data. Therefore, 156 regressions are estimated instead of only 78.

The methodology chosen for the unit root test of the residuals from the OLS regression is equivalent to the one described above with one exception. Instead of using the critical values of MacKinnon (1996), the critical values of MacKinnon (1991) are applied. The rejection of the null hypothesis of a unit root of the residuals indicates that the two time series are cointegrated.

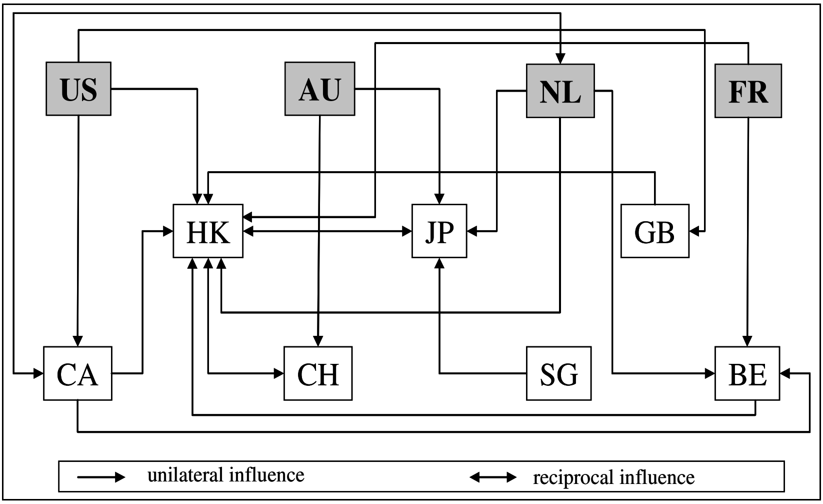
For 30 out of the 156 residual series, the null hypothesis of a unit root is rejected by the ADF test at least at the 10 percent level of significance and thus, these real estate markets share a common stochastic trend and are said to be cointegrated. While for nine relationships, this result is independent of the endogenous and exogenous variables, the modelling matters for twelve pairs of real estate markets. Table 4 summarizes the unit root tests for the cointegrated real estate market indices. A more graphic illustration of the linkages between the different markets is given in Figure 3.

Table 4 Results for Bivariate Cointegration between Real Estate Markets

Indices		Unit Root Tests in Regression Residuals	
Endogenous Variable	Exogenous Variable	ADF_T	ADF_C
CH	AU	-4.0540** (6)	
AU	CH	-3.7619* (6)	
FR	BE		-3.7980** (10)
BE	FR		-3.7160** (10)
HK	BE		-3.9727*** (10)
BE	HK	-3.8274** (10)	
NL	BE		-4.0560*** (10)
BE	NL		-4.0231*** (10)
NL	CA	-4.2359** (7)	
CA	NL		-4.2126*** (7)
US	CA		-4.7553*** (5)
CA	US	-5.1299*** (5)	
US	CH	-3.6360* (6)	
CH	US	-3.8746** (6)	
HK	GB		-3.7181** (8)
GB	HK	-3.5773* (8)	
US	GB		-4.0159*** (9)
GB	US		-3.9220** (9)
HK	AU		-3.1274* (8)
CA	BE		-3.1620* (3)
HK	CA		-3.2648* (6)
HK	CH		-3.1457* (7)
HK	FR		-3.2724* (7)
JP	FR		-3.2991* (0)
AU	JP		-3.1354* (0)
HK	JP		-3.3769** (5)
HK	NL		-3.3658** (7)
JP	NL		-3.2596* (0)
JP	SG	-3.9869** (0)	
HK	US		-3.4560** (8)

Notes: Approximate critical values for ADF tests are based on MacKinnon (1991). ***, **, and * indicate the rejection of the null hypothesis of a unit root at the 1%-, 5%-, and 10%-levels of significance. The lag lengths for the unit root test of the regression residuals are given in parentheses. For brevity, Table 4 presents the results for cointegrated relationships only. The results for all bivariate cointegration tests (156 in total) are available from the author upon request.

Figure 3 Linkages between Real Estate Markets



Notes: The German and Italian real estate stock markets do not share a cointegration relationship with other markets and are thus not mentioned in Figure 3.

While correlation analyses indicate pervasive benefits from diversification across securitized national real estate markets, cointegration analyses offer a different conclusion. During the period investigated, there are long-term interdependences between eleven national real estate markets, which narrow the benefits of international diversification. The German and Italian real estate stock markets are unique, sharing no common stochastic trend with any other market and thus enabling investors from these markets to gain substantial benefits from broadening their investment horizon to other markets, which are both intra- and inter-continental. On the other hand, international investors who are looking for long-run diversification opportunities might be attracted to these markets. However, there is a substantial shortcoming of these two markets. Both are low capitalized and dominated by a few listed property companies, which extensively limit the attractiveness and investability. With respect to the eleven remaining securitized real estate markets under consideration, the Swiss and Belgian markets are very small, low capitalized, dominated by a small number of listed property companies, and characterized by a thin trading volume. Thus, both markets are afflicted by the same limitations as the German and the Italian markets. However, as shown below, neither the Belgian nor the Swiss market plays a key role or has any influence on the market. Rather, they mainly adjust to the changes and trends of large and well-functioning markets, such as FR, NL and the US.

In consideration of the Asia-Pacific markets, the Japanese market shows cointegration relationships with all three other markets within this region; namely, AU, HK, and SG. However, there is only one further long-run relationship between the Australian market and the one in HK. It is clear from Table 4 that the cointegration relationship between AU and the Asian markets is weaker than that between the Asian markets. This result from the ADF test is in line with the economic motivation that the Australian economy in total and the securitized real estate market in particular are more developed, were not affected by the Asian and Russian crises in the late 1990s as much as HK and SG, and have shown a more stable performance during the last 20 years. In taking the Japanese market into consideration, one must bear in mind that the Japanese economy and its stock market suffered from deflation, decreasing house and stock prices, and slow economic growth. These factors have resulted in a huge budget deficit and an unstable banking system with highly indebted banking and insurance companies during the last 20 years. Hence, the Japanese market cannot be compared to the Australian situation in the last 20 years. In spite of these facts, the long-term benefits from diversification across Asian real estate markets are limited for long-term oriented investors with passive investment strategies. The results of these strong long-run equilibrium relationships among the Asian real estate markets contradicts with the findings by Liow et al. (2005), which do not find any cointegrating relationships among the four Asian property stock indices of HK, JP, Malaysia, and SG. Using the Engle-Granger-test for cointegration, Garvey et al. (2001) identify only one long-term relationship between the real estate markets of AU and SG during the period of 1993 to 2001, but no further cointegrating relationship between AU, HK, JP, and SG.

Focusing on inter-continental relationships between Asia-Pacific and European or North American real estate markets, there exist only three weak long-term relationships with the exception of the real estate market in HK. The Japanese market is cointegrated with the two largest real estate markets in continental Europe: FR and NL. For AU's real estate market, a cointegration relationship with the Swiss market is identified. However, the linkage to the Swiss market is almost negligible for investors due to the reasons mentioned above. Only the market in HK is characterized by several long-term relationships with both the European and the North American real estate markets. However, as will be discussed later on, the relationship, on the whole, is not dominated by HK.

In summary, we find that in line with the findings of the correlation analysis (with some limitations with regards to HK), investors located in Asia can benefit from broadening their investment horizon to AU, Europe, and North America. In consideration of real estate investments in the Asia-Pacific region, the Australian market is probably the most attractive for international investors due to its low risk, low correlation, and because there are no strong long-term relationships with the international markets. Through HK's close link to the Chinese market, investors could benefit from China's fast growing economy, with its booming construction sector and large infrastructure projects, as well as from investments in HK's real estate stock market.

In contrast to the findings in the Asia-Pacific area, much stronger long-term relationships can be identified among the Anglo-Saxon markets; namely, CA, GB, and the US, with the latter being cointegrated with the two former markets. Additionally, the US market shows a cointegration relationship with the Swiss real estate market, which adds an interesting feature to the Swiss market. While there is no cointegration with any other market in Europe, the Swiss real estate market is linked with the market in AU, HK, and the US, which reflects the distance of the European Monetary Union (EMU). The GB real estate market shows a similar picture. By contrast, CA is more closely connected to the EMU via long-term linkages to the Dutch and Belgian securitized real estate markets, additionally supported by a correlation coefficient of around 0.50 between the Canadian market and the markets in BE, FR, and NL.

The Dutch market could be counted as part of the Anglo-Saxon oriented markets as well, even if the categorization is not that clear-cut. On the one hand, NL is a member of the EMU, being historically and geographically linked to continental Europe. On the other hand, the Dutch financial market is affected by the Anglo-Saxon system, to which it is also quite similar, being based more on financial markets and thus belonging to the so-called market-oriented systems. Due to its economic size, the Dutch stock market is highly capitalized, and securitized real estate markets have a longer history than those in DE or FR, where the financial system is built on a bank-oriented system. Therefore, NL is somewhere between the typical Anglo-Saxon and continental European markets, which also becomes apparent when considering correlation coefficients and the findings of the cointegration analysis. With the exception of the three Asian and the Swedish real estate markets, correlation coefficients for the Dutch market are higher than 0.40. The results from the cointegration analysis reveal long-term relationships with CA and the neighboring market in BE, as mentioned above. The results for the French market show the same tendency. Pairwise correlations are higher than 0.40 with the exception of the Asian markets again, and we also find cointegration with the Belgian market. On the other hand, no such linkage is found to the Anglo-Saxon market, which does not come as a surprise when considering the historical background of the financial system in FR.

In summary, the results of the correlation and the cointegration analyses show a mixed compatibility and are to some extent, supportive of each other, even when it is worth emphasizing that both analyses are not redundant, as they focus on different time horizons, and are based on different assumptions. The results suggest that intra-continental diversification is less beneficial for investors than inter-continental diversification. However, the previous examinations are flawed in that they fail to regard causality between the national real estate stock markets.

As a further robustness check of the results above, cointegration analysis is conducted in a common currency, the US dollar. The results do not differ much from the ones for local currencies when the same method is applied. There are only two cointegration relationships that cannot be confirmed. However, both of these relationships, the Belgian-Canadian as well as the Australian-Japanese relationships, are only weakly significant when using local currencies and become insignificant when using the US dollar as a common currency. Thus, the results can be seen as quite stable with respect to the applied currency denomination.

4.4 Short-Term Relationships According to the Error Correction Model

While cointegration methodologies present a concept of modeling long-term relationships, nothing has been said about the short-term behavior of cointegrated markets until this point. In general, cointegrated markets share a common stochastic trend, but both types of markets fluctuate around this common trend and do not adhere to their long-term path at each point in time. From an investor's perspective, it is of interest how and by which market the adjustment takes place, when one or both markets move away from the long-term path of equilibrium. This procedure is often modeled by an ECM, which indicates the direction and rate of adjustment. In this paper, the analysis is conducted by the ECM-framework presented above. The ECM is estimated by an OLS regression with stationary variables, including an intercept term, the lagged residuals from the cointegration equation and the lagged returns of both cointegrated markets up to six months as exogenous variables and the actual return as an endogenous variable. The model is re-specified until only the significant coefficients for the lagged returns are left.

The magnitude and the sign of the regression coefficient from the cointegration equation residuals are of special interest and indicate the rate and direction of adjustment as presented in Table 5.[2] The results are not uniform, but mixed. For one-sided cointegration relationships (between CA and NL), the deviation from the common long-term stochastic trend is revised through the impact of both markets. The coefficients have the "correct" sign and are significant. It is also shown that the magnitude of the coefficients is almost identical, regardless of how the regression is run. Furthermore, the adjustment process takes place very quickly compared to the other adjustment processes specified. Additionally, there are five further stable long-term relationships, in which the estimation of the ECM results in two significant adjustment coefficients. While the sign of the coefficient is "correct" in the

[2] To maintain a clear layout, we only present the adjustment coefficient. The model specification is available from the author upon request.

sense that both markets contribute to stabilizing the process between HK and CH, and between HK and JP, respectively, the other three ECMs (AU/HK, FR/JP, and CH/US) do not indicate any stabilizing processes. However, one must bear in mind that the coefficients that are responsible for destabilizing are not highly significant and that the value of the coefficients is small in magnitude, compared to the adjustment coefficient of the other market. For all the other cointegrated securitized real estate markets, the adjustment process is driven by one market only (see Figure 3), but the rate of adjustment tremendously varies between the individual pairwise ECMs. While the adjustment coefficient is estimated to be 0.23 for the ECM between CA and the US, this coefficient is reduced to only one tenth between AU and JP. The relatively high value of 0.23 means that almost one fourth of the deviation from the long-term common stochastic trend is adjusted within one period. An effect of a similar magnitude is observed for the relationship between the Canadian and the Dutch real estate markets, by adding up the absolute values of the two coefficients (0.1214 and 0.1039). For all the other linkages, the adjustment process works much slower. In reference to Table 5, the average adjustment process takes place faster for these cointegration relationships, in which the cointegration residuals are stationary, regardless of the regression specification. This finding confirms the empirical evidence mentioned above with regards to the properties of stationarity of the cointegration residuals.

From an economic point of view, the findings from cointegration analyses and the ECM(s) are in line with the assumption and the empirical evidence for common stock markets, according to which transmission and causality move from the most developed and highly capitalized markets to the smaller markets. This issue is well documented for the relationship between the two neighboring markets of FR and BE, and CA and the US, respectively, but also for the majority of the other relationships, e.g. AU and CH or GB and the US Furthermore, it is shown by Figure 3 that there are one or two leading markets for each region, such as AU for Asia-Pacific, the US for the Anglo-Saxon region, and FR and NL for the continental European countries which share one common currency. To find further evidence on the direction of causation, Ganger causality tests are conducted. The findings support the results from the cointegration analysis and the error correction modelling from above.

Table 5 Direction and Rate of Short-Term Adjustments between Cointegrated Markets

Indices		Adjustment coefficient of the ECM for	
Endogenous variable	Exogenous variable	Endogenous variable	Exogenous variable
CH	AU	-0.1020***	-0.0177
AU	CH	0.0035	0.0819***
FR	BE	-0.0020	0.0741***
BE	FR	-0.1490***	-0.0016
HK	BE	-0.0596***	0.0057
BE	HK	-0.0101	0.0578**
NL	BE	-0.0136	0.0890***
BE	NL	-0.1362***	0.0051
NL	CA	-0.1035**	0.1054**
CA	NL	-0.1214**	0.1039**
US	CA	-0.1180	0.2061***
CA	US	-0.2265***	0.1106
US	CH	0.0116	0.0531***
CH	US	-0.0672***	-0.0304*
HK	GB	-0.0961***	-0.0006
GB	HK	-0.0098	0.0778***
US	GB	-0.0082	0.0309*
GB	US	-0.0513**	-0.0006
HK	AU	-0.0715***	-0.0157*
CA	BE	-0.0402	0.0493**
HK	CA	-0.0721**	0.0184
HK	CH	-0.0548***	0.0204*
HK	FR	-0.0800***	-0.0015
JP	FR	-0.0557***	-0.0247**
AU	JP	-0.0047	0.0230***
HK	JP	-0.0419***	0.0418***
HK	NL	-0.0661***	0.0103
JP	NL	-0.0748***	-0.0170
JP	SG	-0.0554**	0.0247
HK	US	-0.0874***	0.0000

Notes:
***, ** and * indicate significance of the coefficient from the OLS regression at the 99%-, 95%-, and 90%-confidence levels, respectively.

The estimation results from the ECM bear implications for the hypothesis of market efficiency and feasible trading strategies. With respect to the definition of the weak form of market efficiency by Fama (1970 and 1991), the existence of cointegration relationships and Granger causality rejects the hypothesis of market efficiency, because there are lagged linkages between markets. At the same time, the findings raise the question if and how investors can benefit from this type of inefficiency. For the investor who uses active trading strategies, the deviations from the stable common long-term trend can be exploited in two ways, depending on the market situation. First, when the responding market is above its correct level according to the cointegration relationship, it is attractive to sell this market. On the other hand, when the responding market is below its theoretically expected level, this market should be bought. Analogous thoughts apply when both markets are responding. In this case, one market should be bought and the other should be sold to exploit the deviations from the common equilibrium. Subject to the estimated adjustment coefficients, these effects are highly pronounced for the cointegration relationships and the corresponding ECMs between the real estate markets in CA and the US, and the markets in BE and FR, respectively. With respect to the extension of the adjustment process, similar effects are exploitable based on the markets of CA and NL, where the sum of the two significant coefficients in absolute terms (0.1214 and 0.1039 respectively) adds up to 0.2253 and thus, the effect is equivalent to the one between CA and the US This raises the question of whether these effects are exploitable with regards to trading strategies and after trading cost. However, this is not the focus of the paper and will therefore be left for further research.

5. Empirical Results from Multivariate Cointegration Analysis

In the previous section, the bivariate, two-step cointegration methodology suggested by Engle and Granger (1987) is considered. This approach shows deficiencies in at least two points. First, it is not possible to test for multivariate cointegration, which is especially important when benefits from diversification are discussed and more than two markets are considered. Second, the long-term cointegration relationship and the short-term ECM are estimated in two stages, not simultaneously. Thus, to overcome these limitations, the analysis in Section 4 is extended by applying the multivariate cointegration methodology suggested by Johansen (1988) which is briefly described in Section 2.4. This approach enables us to combine short-term and long-term analyses and becomes much more concise and precise on the linkages between more than two markets, which are particularly important when analyzing relationships between several markets from one continent and at least one market from another continent (inter-continental linkages).

The analysis of inter-continental long-term relationships is conducted as follows. First, the multivariate cointegration methodology is applied to markets from the same economic and geographical region (Asia-Pacific, Europe, and North America). Second, the most developed and highly capitalized markets of each region (AU and HK, FR and GB, as well as the

US) are added to the regional markets and the procedure of step one is repeated. Third, inter-continental relationships are challenged by conducting exclusion tests. This means that we test whether the added market can be excluded from both the cointegration relationship and/or the vector ECM. As mentioned above, the appropriate lag length is selected according to the AIC and the significance of the cointegration relationship is measured by the trace statistic.
The results from the multivariate cointegration analysis are presented in Table 6. In consideration of the Asia-Pacific markets and their long-term relationships to the leading markets of the other continents, it can be summarized that neither the two European markets nor the US market can be excluded from the long-term cointegration relationship. However, all three markets can be excluded from the vector ECM and thus do not adjust to deviations from the long-term relationship and are weakly exogenous. This is in line with the findings from Section 4, at least for the French and US markets. In the case of GB, it should be emphasized that the number of cointegration equations increases from one to two and that GB can be excluded from one of these two relationships. This can be seen in support of the results from the Engle-Granger approach in which GB shows a common long-term relationship with the market in HK, but not with any of the other three Asia-Pacific markets. For FR and the US, the relationship might also be driven by the linkage to the market in HK as shown in Figure 3.

Two cointegration equations are identified in the analysis of the linkages between the seven European markets. The results obtained by adding Asia-Pacific markets or the US market are mixed. While both Asia-Pacific markets share a common long-term relationship with at least one European market, the US market can be excluded from the long-term as well as the short-term relationship with the European markets. However, these findings are not too surprising when considering Figure 3. It can be seen that the number of bivariate cointegration relationships is much higher between Europe and HK than between Europe and the US market. However, HK can be excluded from the vector ECM and is weakly exogenous. In the case of AU, the number of cointegration equations increases to three and a more detailed analysis shows that AU is excludable from all cointegration equations from which CH is excluded as well. This suggests that the third cointegration equation in the multivariate analysis is driven by the relationship between AU and CH, which is identified as the unique relationship between AU and any European market in the bivariate framework.

For North America, it is shown in Table 6 that neither the Asia-Pacific nor the European markets show any significant linkage to the Canadian and US markets with the exception of HK. However, even this relationship is only weakly significant.

Table 6 Results from Multivariate Cointegration Analysis

Regional Markets	Added Market i	Number of Lags	Number of Cointegration Equations	Exclusion Tests (Test Statistic) $\beta_i = 0$	$\alpha_i = 0$	$\alpha_i = \beta_i = 0$
AU, HK, JP, SG		0	1^{***}			
AU, HK, JP, SG	FR	1	1^{***}	O	X	O
AU, HK, JP, SG	GB	0	2^{***}	O	X	O
AU, HK, JP, SG	US	1	1^{***}	O	X	O
BE, CH, DE, FR, GB, IT, NL		0	2^{**}			
BE, CH, DE, FR, GB, IT, NL	AU	0	3^{**}	O	O	O
BE, CH, DE, FR, GB, IT, NL	HK	0	2^{***}	O	X	O
BE, CH, DE, FR, GB, IT, NL	US	0	2^{***}	X	X	X
CA, US		1	1^{***}			
CA, US	AU	2	1^{**}	X	X	X
CA, US	HK	1	1^{***}	O	X	X
CA, US	FR	1	1^{***}	X	X	X
CA, US	GB	1	1^{**}	X	X	X

Notes: The coefficients α_i and β_i have the same meaning as those in Section 2.3. In column 4, ***, ** and * indicate significance of the trace statistic at the 99%-, 95%-, and 90%-confidence levels, respectively, according to MacKinnon et al. (1999). In columns 5, 6, and 7, X indicates that the null hypothesis of $\alpha_i = 0$ and / or $\beta_i = 0$ is not rejected at the 5% significance level, otherwise O.

Summarizing the results from the multivariate cointegration analysis and considering inter-continental linkages, it can be stated that the results are mainly in line with the findings from the bivariate cointegration analysis and that markets in other continents are at least weakly exogenous and should provide diversification opportunities. However, the long-term linkages between the Asia-Pacific and European markets seem to be a little bit stronger than those to the North American markets. Furthermore, the results are qualitatively quite similar when the estimations are conducted with US dollar denominated indices.

6. Conclusion

In the relevant literature, the authors have often argued that diversification benefits are driven by country factors, thus broadening the investment horizon from a domestic to a more global perspective. This improves the mean-variance-characteristics of a portfolio by an upward shift of the efficient frontier. The achievement of these beneficial return-risk-characteristics is often based on a concept in which risk reduction is measured by correlation and covariance structures between the returns of different assets or markets. However, correlation analyses are accompanied by some essential limitations, which were discussed above in more detail. First, from a technical point of view, the returns have to be normally distributed when applying portfolio optimization based on correlation analyses. However, as also shown above, this assumption does not hold, at least not for real estate returns. Second, correlation coefficients only capture the short-term dependence between these assets and investors, who are usually interested in long-term linkages between prices, the focus of cointegration analyses. Third, correlation analyses are combined with a loss of valuable information contained in the time series, since correlation coefficients have to be based on stationary variables and price indices are not commonly stationary. Hence, first differences or logarithmic returns have to be used combined with information on the level of the price series, which is important information for long-run oriented investors. Thus, the investigation of the cointegration of prices rather than the correlation of returns is a more appropriate approach with regard to a long-run oriented investor type.

By using 14 securitized real estate markets in total; 4 from the Asia-Pacific region, 8 from Europe, and 2 from North America, the findings, based on the approach suggested by Engle and Granger (1987), provide the following main conclusions:

First, there exist several cointegration relationships between national real estate markets, between continents and within continents. Second, it is shown that within each region, there are one or two key markets that influence neighboring markets, such as AU in the Asia-Pacific region, the US in the Anglo-Saxon area, and FR and NL in the EMU. This implies that focusing on these central markets is sufficient from an investor's point of view and reduces the efforts of analyzing the international real estate markets. Third, the finding of stable long-term relationships across real estate markets challenges the implications given by low correlation among national securitized real estate markets. The weaker long-term linkages between national real estate markets across continents suggest that long-term oriented investors benefit from extending their investment horizon beyond domestic markets, while long-term benefits from diversification across markets within a single continent are limited. These findings are quite stable with respect to the currency denomination and supported by the results from multivariate cointegration analyses. Fourth, the findings from cointegration analyses, the modelling of ECMs, and Granger causality tests raise further questions on the validity of the efficient market hypothesis for securitized real estate markets. The question that arises, from the perspective of an investor who is using active trading strategies, is whether these effects are exploitable by means of trading strategies and after trading cost. However, this question is not the focus of the present paper and will be left to further research.

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