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Linkages between International Securitized Real Estate Markets: Further Evidence from Time-Varying and Stochastic Cointegration

Felix Schindler and Svitlana Voronkova

ZEW

Zentrum für Europäische
Wirtschaftsforschung GmbH

Centre for European
Economic Research

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Non-technical Summary

This paper analyzes long-run co-movements between 14 international real estate stock markets and between three economic and geographic regions based on bivariate and multivariate tests for cointegration. While the topic has been analyzed by previous studies such as Gallo and Zhang (2009) and Yunus (2009) among others, this paper is of significant contribution to existing studies since we compare results from different cointegration methodologies. To our knowledge, it is the first study that explicitly controls for structural breaks in the cointegration relationships and consider time-varying cointegration as well as stochastic cointegration when analyzing long-run co-movements between international securitized real estate markets. Furthermore, using data from 1990 to 2009 the covered period is not only characterized by fast growing and upward moving real estate stock markets as many previous studies but also by the period of the current and still ongoing financial crisis that started in 2007.

In line with previous studies, the empirical results indicate several cointegration relationships between national real estate stock markets. However, it is shown that most cointegration relationships are unstable and time-varying and that the results from cointegration methodologies suggested by Engle and Granger (1987) and Johansen (1988) might be misleading in that common long-run co-movements are time-varying and are much stronger when structural breaks are considered. Additionally, the detected cointegration relationships are much stronger between national markets within one economic and geographic region than between national markets located in different regions.

Thus, from an investor's point of view, the results indicate that broadening the investment horizon from the domestic continent to others regional markets might be more beneficial than diversifying within one region. This conclusion applies particularly to the European real estate stock markets and thus to investors holding European real estate securities.

Das Wichtigste in Kürze

Dieser Aufsatz analysiert die langfristigen Zusammenhänge zwischen 14 internationalen Immobilienaktienmärkten und zwischen drei ökonomischen und geographischen Regionen. Dabei werden sowohl bivariate als auch multivariate Kointegrationsverfahren eingesetzt. Während sich zahlreiche, bisherige Untersuchungen wie z. B. von Gallo und Zhang (2009) und Yunus (2009) mit dieser Thematik befassen, ist der wissenschaftliche Beitrag dieser Studie vor allem im Vergleich verschiedener Kointegrationsverfahren zu sehen. Nach unserem Kenntnisstand handelt es sich bei der Untersuchung um die erste Analyse für internationale Immobilienaktienmärkte, die explizit Strukturbrüche in der Kointegrationsbeziehung und zeitlich variierende Kointegrationsbeziehungen berücksichtigt sowie stochastische Kointegrationsprozesse betrachtet. Außerdem handelt es sich um einen vergleichsweise langen Untersuchungszeitraum von 1990 bis 2009, der nicht nur nicht nur von einem stark wachsenden und steigenden Immobilien(-aktien-)marktumfeld geprägt war, sondern auch die Auswirkungen in Folge der seit dem Jahr 2007 anhaltenden Finanzmarktkrise umfasst.

Im Einklang mit bisherigen Untersuchungen deuten die empirischen Ergebnisse auf zahlreiche Kointegrationsbeziehungen zwischen den nationalen Immobilienaktienmärkten hin. Allerdings zeigt sich auch, dass sich die Kointegrationsbeziehungen über die Zeit als instabil erweisen, die Ergebnisse – basierend auf den von Engle und Granger (1987) und Johansen (1988) vorgeschlagenen Verfahren – unzureichend sind und die Kointegrationsbeziehungen zunehmen, wenn Strukturbrücke berücksichtigt werden. Des Weiteren zeigt sich, dass zwischen den nationalen Märkten eines Kontinents mehr Kointegrationsbeziehungen bestehen als zwischen nationalen Märkten verschiedener ökonomischer und geographischer Regionen.

Aus Investorensicht deuten die Ergebnisse darauf hin, dass die langfristigen Diversifikationspotentiale für Investoren eher in interkontinentalen als in intrakontinentalen Immobilienaktienanlagen zu realisieren sind. Dies trifft vor allem auf den europäischen Markt zu. Des Weiteren wird durch die Analyse deutlich, dass durch Kointegrationsverfahren, die Strukturbrüche unberücksichtigt lassen, langfristige Zusammenhänge und Gleichläufe zwischen den einzelnen nationalen Immobilienaktienmärkten nur unzureichend identifiziert werden und Fehlspezifikationen vorliegen können.

Linkages between International Securitized Real Estate Markets: Further Evidence from Time-Varying and Stochastic Cointegration

Felix Schindler* and Svitlana Voronkova**

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Abstract

This paper analyzes long-run co-movements between international real estate stock markets and between regions based on bivariate and multivariate tests for cointegration. While the topic has been analyzed in previous studies such as Gallo and Zhang (2009) and Yunus (2009) among others, this paper is of significant contribution to existing studies since we compare results from different cointegration methodologies and explicitly control for instability in cointegration relationships and deviations from normality. Furthermore, the analyzed time period is longer than in previous studies and ranges from 1990 to 2009 covering 20 years. In line with previous studies, the empirical results indicate several cointegration relationships between national real estate stock markets. However, it is also shown that most cointegration relationships are unstable and that the results from cointegration methodologies suggested by Engle and Granger (1987) and Johansen (1988) might be misleading in that common long-run co-movements appear to be stronger when structural breaks are considered. Thus, the results indicate that investors would benefit from broadening their investment horizon from their domestic continent to international markets. This particularly applies for the European securitized real estate markets.

Keywords: International Securitized Real Estate Markets, Diversification, Time-Varying Cointegration, Stochastic Cointegration

JEL Classifications: C22; C11; G14

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* Corresponding author, Centre for European Economic Research (ZEW), P.O. Box 10 34 43, D-68034 Mannheim, Germany; Phone: +49-621-1235-378; Fax: +49-621-1235-223; E-mail: schindler@zew.de.

** Centre for European Economic Research (ZEW), P.O. Box 10 34 43, D-68034 Mannheim, Germany; Phone: +49-621-1235-379; Fax: +49-621-1235-223; E-mail: voronkova@zew.de.

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1 Introduction

The issue of diversification and the question about what assets can provide investors with diversification benefits have been at the forefront of financial literature for at least last three decades (Longin and Solnik (1995), Goetzmann et al. (2005)). There exists substantial body of research which studies diversification potential of various asset types when added to a stock portfolio. Given the growing importance of securitized real estate markets and their perceived segmentation from the non-real estate securities, numerous studies have analyzed potential diversification benefits from adding domestic securitized real estate to a portfolio of traditional assets, such as equity, bonds and cash (Okunev and Wilson (1997), Ling and Naranjo (1999), Chaudhry et al. (1999), Glascock et al. (2000), Conover et al. (2002), Liow and Yang (2005), Bond and Glascock (2006) and Westerheide (2006), among others). The consensus emerging from these studies is that over time the extent of such benefits is likely to have declined due to increased integration of domestic real estate and stock markets.

Despite some studies have highlighted the advantages of including *foreign* securitized real estate such to achieve more efficient portfolios (see studies mentioned in Worzala and Sirmans (2003)), only a limited number of papers studied potential diversification benefits that may arise due to investing in international real estate securities (Eichholtz (1996a), Eichholtz et al. (1998), Lizieri et al. (2003), Liow et al. (2005), Yang et al. (2005), Schindler (2010)). This dearth of studies is surprising provided the remarkable growth in real estate markets around the globe in the past two decades and the introduction of real estate investment trusts (REITs) in a number of countries, an event that was expected to give significant impetus to the development of these markets (Schindler (2010)).[1]

Since the seminal paper by Grubel (1968) analyses of correlations between international asset markets has become a major tool for making inferences regarding presence of diversification benefits. However, while correlation coefficients capture short-term co-movements among asset returns they fail to provide information about the long-term diversification benefits, which might be of interest to investors with long investment horizons, such as, for example, pension funds. When correlation analysis is used to

[1] REIT(s) were introduced in Germany, Italy and the UK in 2007, in Hong Kong in 2003, Japan in 2000, and in Singapore in 1999 (Schindler (2010)).

measure long-term diversification benefits it has to be considered that correlation and covariance structures are time-varying. In contrast to other studies indicating time-varying correlation coefficients, Eichholtz (1996a) and Schindler (2009a) show that the hypothesis of stable correlation matrices over time is statistically rejected for international securitized real estate markets. These findings put further doubt on the application of correlation analysis and the mean-variance framework for assessing long-term diversification benefits. Additionally, correlation analysis may result in a loss of valuable information contained in asset price series, since correlation coefficients have to be calculated using stationary variables and asset prices are, as a rule, non-stationary. This requires taking first differences of the prices before correlation coefficients may be calculated (Schindler (2010)).[2] Cointegration analysis developed by Engle and Granger, Johansen (1988) and Johansen and Juselius (1990) is devoid of both disadvantages. Kasa (1992) points out that it is a preferred technique for establishing presence of common long-term trends and establishing presence of long-term diversification benefits. Provided that real estate is commonly viewed as a suitable asset for long-term investment, this is a particularly apt methodology for studying real estate prices (NAREIT (2009)).

Up to our knowledge, there are only few studies of *long-term diversification benefits* from investment in international real estate securities. The aim of this paper is to provide new information on long-term diversification benefits of investment in international securitized real estate and obtain a more in-depth understanding of linkages among international real estate markets by studying them in a time-varying framework. The studies of long-term diversification benefits mentioned above either assumed stable long-term relationships or supposed that a change in long-term relationships occurred around an assumed date. They generally find no or limited long-run co-movements in international real estate security prices and conclude that investing in international real estate securities may result in a better diversified portfolio. However, a number of studies have underscored the time-varying nature of inter-market relations for common stock markets (Longin (1995), Bekaert and Harvey (1995), Gelos and Sahay (2000), and Ang and Bekaert (2002)). They point out that the violation of the stability assumption is especially likely to occur over long periods. This suggests that we may obtain a better

[2] Additionally, most of the existing studies assume normally distributed asset returns and static correlation coefficients. Both of these assumptions have been challenged in several studies (Mandelbrot (1963), Fama (1965), Longin and Solnik (1995), Goetzmann et al. (2005). The assumption of normality in real estate returns has been challenged by Liow (2007) and Brounen et al. (2008) among others.

understanding of the integration processes in the international real estate securities' markets by studying them in a time-varying framework. In other words, evidence of lack of long-term relationships and resulting from it conclusion about the presence of diversification benefits may not hold when static methodology is applied to data where instability and structural breaks are present and long-term relationships, subject to breaks, exist.

Campos et al. (1996) and Gregory and Hansen (1996) show that neglecting structural breaks leads to the underrejection of the null hypothesis of no cointegration. This has important implications for long-term asset allocations in that one may overestimate the extent of diversification benefits when static long-run relationships are assumed. Therefore, in this paper we propose to use time-varying cointegration tests by Gregory and Hansen (1996) and Hansen and Johansen (1999). The key advantage of the Gregory and Hansen (1996) test over conventional cointegration tests is that in case of instability in a long-run relationship, the test allows to estimate the date of the structural break endogenously. Additionally, such a date can vary across the markets. Thus, there is no need in imposing potentially incorrect structural change date on the data. Moreover, in the wake of the recent financial crises, such an approach will allow us to make inferences about the effect it might have had on the long-term relationships among real estate markets. The Hansen and Johansen (1999) recursive cointegration test allows testing for cointegration in a multivariate framework. To accommodate the observation that the distribution of the real estate securities' returns tend to deviate from normality (Seiler et al. (1999), Lizieri et al. (2003), Brounen et al. (2008)), we also apply a recently developed stochastic cointegration test by McCabe et al. (2003).

While there is a long tradition of using asset price interrelationships to measure market co-movements (Longin and Solnik (1995), Arshanapalli and Doukas (1993), Bachman (1996), Koedijk et al. (2002), Garcia Pascual (2003)), another illuminating way of studying asset market integration is to study the equality of risk premia across the markets as implied by the law of one price (Kasa (1992), Bekaert and Harvey (1995), Baele et al. (2004)). Specifying appropriate asset pricing model for estimation of the risk premia is a challenge posed by this approach. We leave analysis of integration by means of establishing the equality of risk premia across international real estate markets for future research.

In this paper we focus on securitized rather than direct real estate. The reason behind this is that the former asset is much more liquid, frequently traded and transparent and

therefore is open to a much broader spectrum of investors. Additionally, Oikarinen et al. (2009) and Morawski et al. (2008) show for the U.S. market, that securitized real estate is cointegrated with and leading direct real estate in both short and long run. Additionally, while securitized real estate prices behave more like stock prices in the short term, in the long term they show more affinity with the direct real estate market.
The remainder of the paper is laid out as follows. Section 2 provides a review of relevant literature. This is followed by the review of econometric methodology in Section 3. After describing the data in Section 4, the empirical findings are presented in Section 5, while Section 6 summarizes the results and draws some concluding remarks.

2 Literature Review

Up to our knowledge, most studies on diversification benefits from investment in the international real estate securities mainly focus on short-term analysis and studies of long-term diversification benefits are quite scarce. By comparing correlation structures on real estate stock markets, bond markets and common stock markets, the results by Eichholtz (1996b) indicate that real estate stock markets are less strongly internationally correlated than common stock and bond markets. This implies that international diversification can reduce the risk of securitized real estate portfolios even more than it can reduce the risk of common stock and bond portfolios at least in the short run. Applying a multi-factor and multi-country model, Bond et al. (2003) find strong local market risk factors attesting the adequacy of international portfolio diversification for U.S. real estate investors. Eichholtz et al. (1998) find evidence of a strong European factor, which appear to strengthen since the early 1990s after the European single market has been launched. Lizieri et al. (2003) study the transmission of monetary integration within the European Union into equity and real estate markets. Using a number of methods, such as analysis of correlations, principal component analysis, Granger causality tests and analysis of the impulse response functions from the vector autoregressive (VAR) models, they find only weak evidence of increased integration across the eight Euro-zone countries. These findings contrast sharply with the results for equity markets which demonstrate increased integration since 1997. The obvious asymmetry in the results between the securitized real estate and equity markets is attributed to the size of real estate companies and domestic composition of their real estate portfolios. The study by Lizieri et al. (2003) highlights the need to consider the long-run relationship in a time-varying framework to gain a more complete understanding of the integration process in the real estate sector.

A study by Garvey et al. (2001) focusing on the four Asia-Pacific markets Australia, Hong Kong, Japan, and Singapore documents limited evidence of cointegration between the four markets applying cointegration methodology suggested by Engle and Ganger (1987) and by Johansen (1988). Additionally, Garvey et al. (2001) show statistically significant performance improvements from extending national real estate stock portfolios into other Asia-Pacific markets. These findings indicate that significant long-term diversification gains can be realized by diversifying real estate portfolios throughout Asia-Pacific markets. However, the study considers four markets only and does not include Non-Asia-Pacific markets like the U.S. or the U.K. The analyzed period from 1975 to 2001 is also a period where securitized real estate markets were mostly undeveloped or at an early stage of market development.

Yunus and Swanson (2007) extend the number of covered markets by adding the U.S. market to the four Asia-Pacific markets analyzed by Garvey et al. (2001) and consider the period between 2000 and 2006. Furthermore, Schindler (2009b) focuses on seven Asia-Pacific real estate stock markets and the markets in the U.K. and the U.S. for the period from 1992 to 2008. Both Yunus and Swanson (2007) and Schindler (2009b) conclude from their results of cointegration analysis, that U.S. investors can derive diversification benefits from investing in Asia-Pacific real estate stock markets. According to Schindler (2009b), this result holds for investors from the U.K. as well.

Yang et al. (2005) study real estate market integration over the period from 1994 to 2002 in a sample consisting of six European Monetary Union (EMU) and three non-EMU members. They use VAR and forecast error variance decomposition methodology. Yang et al. (2005) find one cointegration relationship in the period before and after the EMU launch. However, their forecast error variable decompositions indicate that only a subset of the countries (Germany, France and the Netherlands) show increased integration with other countries in the sample. For rest of the countries, both members and non-members of the EMU, the evidence of integration is mixed in that either no change or less integration after the EMU introduction was found.

A study by Liow et al. (2005) focuses on the long-run relationships among four Asian real estate stock markets (Japan, Hong Kong, Malaysia, and Singapore) and four European real estate stock markets (Germany, France, Italy, and the UK). By applying the cointegration methodology suggested by Johansen (1988), Liow et al. (2005) find weak cointegration for the markets within one continent. The results from global cointegration analysis reveal that there is no cointegrating relationship between the Asian and European real estate stock markets. Thus, the findings indicate, that significant benefits from diversification may be obtained by diversifying a real estate portfolio across continents.

More recent studies are conducted by Gallo and Zhang (2009), Schindler (2010), and Yunus (2009). These studies cover a longer time period and more markets from around the world, from regions such as Asia-Pacific, Europe, and North America.

Gallo and Zhang (2009) analyze fourteen developed markets from North America, Europe, and Asia-Pacific region between 1992 and 2007. They find no long-term relationships among the three regional indices. However, when they analyze relations within the continents, they find one cointegration relation for Asia-Pacific and North American regions and two cointegration relations for the group of the eight European markets. Gallo and Zhang (2009) find only 20 out of possible 98 bivariate cointegration relations, mostly among the European markets. This confirms their notion that weak inter-country relationships suggest strong country enhancements to regional diversification gains. Gallo and Zhang (2009) show that a portfolio consisting of independent markets outperforms a portfolio consisting of cointegrated markets, albeit the former portfolio fails to achieve lower risk and thus seems to be under-diversified.

Yunus (2009) study real estate security returns for seven international markets (Australia, France, Hong Kong, Japan, the Netherlands, the U.K., and the U.S.) over the period from 1990 to 2007. She finds that these markets share three equilibrium relations. However, France and the Netherlands are subsequently excluded from the equilibrium in the result of the exclusion tests. Yunus is the first to conduct a formal time-varying cointegration test of Hansen and Johansen (1999). She finds fist evidence of cointegration after 1998 which has been strengthening afterwards, especially since 2003. Furthermore, she finds a second cointegration relationship emerging towards the end of her sample. Based on this, she concludes that securitized real estate markets follow the path of international stock and bond markets towards more integration.

Schindler (2010) studies fourteen international real estate markets from Asia-Pacific region, Europe, and North America over the period from 1990 to 2008. He uses monthly data on EPRA/NAREIT national real estate market indices. Using univariate test for cointegration by Engle and Granger (1987), error correction models (ECM) and Granger causality tests, he finds a number of univariate long-term relationships among the real estate markets in his sample. He finds that the markets both within and across the continents seem to co-move around a common level in the long-run, although the evidence in favour of such co-movements is much stronger among the markets from the same continents, rather than for the markets from different continents. These findings support the results by Liow et al. (2005) and extend the perspective to a more global one by integrating the markets of Australia and North America to the analysis. ECMs and Granger causality tests help to identify key markets within each of the regions to which other markets within the respective region adjust. These results suggest that investors have to re-consider evidence from correlation analysis on low level of short-term co-movements across international real estate markets. Namely, the short-term diversification benefits from investing in international real estate securities may not realize over the long-term periods. These results are particularly important for investors with long-term horizons, such as pension funds.

The studies of long-term diversification benefits mentioned above either assumed stable long-term relationships or supposed that a change in long-term relationships occurred around an assumed date. They generally find no or only limited long-run co-movements in international real estate security prices and conclude that investing in international real estate securities may result in a better diversified portfolio. However, a number of studies have underscored the time-varying nature of inter-market relations for common stock markets (Longin and Solnik (1995), Bekaert and Harvey (1995), Gelos and Sahay (2000), and Ang and Bekaert (2002)). They point out that the violation of the stability assumption is especially likely to occur over long periods. This suggests that we may obtain a better understanding of the integration processes in international real estate securities' markets by studying them in a time-varying framework.

Thus, the aim of this paper is to focus on this gap in existing research by testing the results of the existing studies regarding long-term diversification benefits of investment in international securitized real estate for longer time periods which may include periods of substantial instability including crises, and obtaining a more in-depth understanding of linkages among international real estate markets by studying them in a time-varying framework.

3 Empirical Methodology

The analysis of long-run co-movements between international securitized real estate markets is conducted using different methodologies of cointegration analysis that are briefly presented in this section. First, the traditional concepts of bivariate and multivariate cointegration analysis suggested by Engle and Granger (1987) and Johansen (1988) and Johansen and Juselius (1990) are described. Second, recursive cointegration analysis based on Hansen and Johansen (1999) is introduced. Thereafter, more recent cointegration tests are considered. While the approach by Gregory and Hansen (1996) explicitly controls for time-varying cointegration, McCabe et al. (2003) develops a testing procedure for stochastic cointegration.

3.1 Engle and Granger (1987) 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 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 using the ordinary least squares regression (OLS) to obtain the residuals:

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

In the second step, the residuals ε_t are tested for unit root by employing the augmented Dickey-Fuller (ADF) test. Since the residuals are not observed values, but are estimated from the OLS regression, MacKinnon (1991) critical values are applied. The critical values K are estimated as follows:

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

where Z denotes the sample size and the 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, I(1), and the residuals from the OLS regression are stationary in levels.

3.2 Johansen and Juselius (1990) Test for Cointegration

Two or more 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 securities 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 process (VAR) with the lag length k given by:

$$x_t = A_1 x_{t-1} + ... + A_k x_{t-k} + \mu + \varepsilon_t, \quad (3)$$

Where x_t is an n-dimensional vector of real estate stock price indices, A_i is a 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 cointegration literature, 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 for computing *k*, we determine the optimal lag length in the VAR system using the Akaike information criterion (AIC). As the subsequent estimations show (not presented in the paper, but available upon request), in case of our sample the lag length chosen using the AIC is the same as the one suggested by other tests.

By first differencing of equation (1), the VAR can be transformed into an error correction model:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + ... + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu + \varepsilon_t, \quad (4)$$

where

$$\Gamma_i = -(I - A_1 - ... - A_i) \text{ with } i = 1, ..., k-1)$$

and

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

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) The 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) The matrix Π is the null matrix (r = 0) which means that n – r = n and indicates that equation (2) corresponds to a traditional differenced vector time series model.

(iii) The matrix Π is of 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 the property that $\beta' x_t$ is stationary. If this is the case, the 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, the 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 test and the maximum eigenvalue test. 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 rely 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 following:

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^{n} \log(1-\hat{\lambda}_i), \tag{5}$$

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

$$\left|\lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k}\right| = 0. \tag{6}$$

The critical values for the trace statistic have been tabulated by MacKinnon et al. (1999). It is well documented in literature that the asymptotic distribution of $\lambda_{trace}(r)$ 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 forward in 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 securitized real estate 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.

3.3 Gregory and Hansen (1996) Time-Varying Cointegration Test

Results of Monte Carlo experiments (Campos et al. (1996), Gregory and Hansen (1996) (1996)) show that in the presence of structural change standard tests for cointegration (like that of Engle and Granger (1987)) may lose power and falsely signal the absence of equilibrium in the system. A number of tests of unit roots under structural stability have been developed (see Maddala and Kim (1998) for an overview). In this paper we use the Gregory and Hansen (1996) test. The Gregory and Hansen (1996) test tests the null hypothesis of no cointegration against the alternative hypothesis of cointegration with a single structural break of unknown timing. That is, this test does not assume stable parameters, but allows the parameters of cointegrating vectors to shift at an unknown time point. The advantage of the Gregory and Hansen (1996) approach is that the timing of the structural change under the alternative hypothesis is estimated endogenously. Gregory and Hansen (1996) suggest three alternative models accommodating changes in parameters of the cointegration vector under the alternative hypothesis of cointegration with structural change. A level shift model allows for change in the intercept only (C):

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha y_{2t} + \varepsilon_t, \ t = 1,...,n \qquad (7)$$

The second model accommodating a trend in the data also restricts shifts to changes in level with trend (C/T):

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha y_{2t} + \varepsilon_t, \ t = 1,...,n \qquad (8)$$

The most general specification allows for changes in both the intercept and slope of the cointegration vector (R/S):

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1 y_{2t} + \alpha_2 y_{2t} \varphi_{t\tau} + \varepsilon_t, t = 1,...,n \qquad (9)$$

The dummy variable, which captures the structural change, is represented as

$$\varphi_{t\tau} = \begin{cases} 0, & t \le [n\tau] \\ 1 & t > [n\tau] \end{cases} \qquad (10)$$

where $\tau \in (0,1)$ is relative timing of the change point. The trimming interval is usually taken to be (0.15n, 0.85n), as recommended in Andrews (1993). The three models presented in equations (7)-(9) are estimated sequentially with break point changing over the interval $\tau \in$ (0.15n, 0.85n). Non-stationarity of the residuals, expected under the null hypothesis, is checked by the ADF- and Phillips-Perron (PP-) tests. Setting the test statistic, denoted as ADF^*, Z_a^*, and Z_t^*, at the smallest values of estimated ADF, Z_a and Z_t statistic in the sequence, we select the value that constitutes the strongest evidence against the null hypothesis of no cointegration.

3.4 Recursive Hansen and Johansen (1999) Test

Hansen and Johansen (1999) and Juselius (2007) applied recursive approach Johansen and Juselius (1990) cointegration tests and developed several tests to analyze the stability of parameters in cointegrated VAR models. Given the multitude of the stability tests and given that this is not the purpose of this paper to exhaust all available stability tests, we focus on the one that in our view is best fitting the purpose of the paper: to explore the presence and the extent of the cointegration relationships among the studied markets without imposing specific restrictions on the individual cointegration vectors. For this reason, we focus on the recursive trace test which we briefly describe below.

Hansen and Johansen (1999) suggest recursively calculating trace test statistic to get an insight into constancy or non-constancy of the individual cointegration relations. This is possible because eigenvalues are shown to be a quadratic function of the vectors α and β. Thus when α and β are reasonably constant, then λ_i will also be constant (Juselius (2007)). Recursive analysis is performed for an initial period and thereafter updated as new data are added to the initial sample. The trace test statistic given by equation (11) is calculated over the base sample, t_0 to t_n. This sample is then extended by j periods and the statistic is re-estimated for the period from t_0 to $t_{n+j.}$. Eventually, the estimation procedure reaches the end of the data, producing the test statistic results equivalent to the standard static Johansen and Juselius (1990) estimation over the entire time period. The statistic are calculated for the **X**-form, or the so-called full model, and for the **R**-form, or the so-called concentrated model. For ease of interpretation, the calculated trace statistic is rescaled by the 95% quintile of the asymptotic distribution for a model without exogenous or dummy variables. The recursive trace test statistic divided by the 95% quintile is given by the following equation:

$$\tau_r = \left\{ -t_1 \sum_{i=r+1}^{n} \ln(1-\hat{\lambda}_i) \right\} / C^*_{0.95}(r), \qquad r = 0,\ldots,n-1, \qquad t_1 = t_n,\ldots,T, \tag{11}$$

where r is the number of cointegration relations and $C^*_{0.95}(r)$ is the 95% critical value for the corresponding null hypothesis.

The recursively calculated trace statistic is then plotted against time and examined for instability. The plotted test statistic provides a visual impression of whether the cointegration relations are reasonably constant. If the equilibrium relationship is constant, then the trace test statistic will also be constant and the graphs will grow linearly with the slope coefficient $\ln(1-\hat{\lambda}_i)$ (Juselius (2007)). Rescaled values above 1 of the trace statistic for the null hypothesis of r cointegration relationships against the hypothesis of n cointegration relationships indicate that the null is rejected.

3.5 McCabe et al. (2003) Stochastic Cointegration Test

It has been noted that some economic variables, including stock prices, tend to be more volatile than assumed for an I(1) process. The recent approach of Harris et al. (2002) suggests considering cointegration in a sense wider than that of Engle and Granger (1987) by loosening the strict requirement of stationarity of first differences of the series and requiring only the absence of stochastic I(1) trends.[3] Their process allows for the presence of a non-linear form of heteroscedasticity that gives rise to volatile behaviour of the first differences of the series. The process in a regression form may be written as

$$
\begin{aligned}
y_t &= \alpha + kt + x_t^{'}\beta + u_t \\
u_t &= e_t + q'w_t + \nu_t^{'}w_t ,
\end{aligned}
\tag{12}
$$

where y_t is a scalar, x_t is a *m x 1* vector, and w_t is a vector integrated process. The regression error term, u_t, is composed of the stationary term, e_t, the integrated term, $q'w_t$, and the heteroscedastic component, $\nu_t^{'}w_t$.

McCabe et al. (2003) suggest that the null hypothesis of stochastic cointegration against the alternative of no cointegration can be expressed as

$H_0 :\ q = 0$ and $H_1 :\ q \neq 0$.

Within H_0, the null hypothesis of stationary cointegration against the heteroscedastic alternative is:

$H_0^0 :\ E(\nu'\nu) = 0$ and $H_1^0 :\ E(\nu'\nu) > 0$.

For deriving the test statistic, McCabe et al. (2003) adopt a semi-parametric approach that does not rely on distributional assumptions. They utilise an asymptotic instrumental variable estimator (AIV) of Harris et al. (2002), which is consistent under heteroscedastic cointegration. The test statistic for the null hypothesis of stochastic cointegration is given by

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[3] It should be noted that the term 'stochastic cointegration' has been previously used (see Campbell and Perron (1991) and Ogaki and Park (1997)) in the sense of a presence of non-zero deterministic trends in an I(0) combination of the I(1) variables. Here however we refer to stochastic cointegration as it is defined by McCabe et al. (2003).

$$S_{nc} = \frac{T^{-1/2} \sum_{t=k+1}^{T} \hat{u}_t \hat{u}_{t-k}}{\hat{\varpi}(\hat{u}_t \hat{u}_{t-k})}, \tag{13}$$

where $k = k(T)$.

Under the cointegrating null hypothesis the test statistic is asymptotically normal distributed with N (0,1*)*. The test statistic for the null hypothesis of stationary cointegration is

$$S_{hc} = \left(\frac{1}{12}\right)^{1/2} \frac{\sum_{t=1}^{T} t(\hat{u}_t^2 - \hat{\sigma}_u^2)}{\hat{\omega}(\hat{u}_t^2 - \hat{\sigma}_u^2)}. \tag{14}$$

Harris et al. (2002) show that this statistic is *N (0,1)* under weak regularity conditions.[4]

4 Data

The empirical analysis in this paper relies on the monthly price indices from the European Public Real Estate Association (EPRA) and the National Association of Real Estate Investment Trusts (NAREIT) between January 1990 and June 2009. The study covers the following 14 national real estate stock markets: Australia (AU), Belgium (BE), Canada (CA), France (FR), Germany (DE), Hong Kong (HK), Italy (IT), Japan (JP), the Netherlands (NL), Singapore (SG), Sweden (SE), Switzerland (CH), United Kingdom (UK), and the United States (US). Furthermore, the study also includes three continental indices for Asia-Pacific (AP), Europe (EU), and Northern America (NA). The time series contains 234 monthly data for each market. Due to the lack of data, the analysis of the Canadian market is based on 150 observations between January 1997 and June 2009 only. The data are obtained from DataStream Thomson Reuters. To our knowledge, it is the most comprehensive analysis of international cointegration analysis in securitized real estate markets. Taking a perspective of an US-investor, sample statistic is calculated using values based in US-dollars. All analyses are conducted in local currency terms as well but the results are almost identical irrespective of the currency used. Therefore and for the sake of brevity, we report the results in US-dollars only.[5] The real estate indices

[4] GAUSS code for calculation of the test statistics was kindly provided by Brendan McCabe.
[5] The results in local currency terms are available from the authors upon request.

are calculated in natural logarithms, whereas the monthly rates of return are calculated as the first differences of the logarithmic monthly index levels. The national real estate indices are delivered by the same index provider (EPRA/NAREIT) to eliminate potential differences in index construction and index criteria which may arise when using data from different index providers.

As pointed out by Bond et al. (2003), Yang et al. (2005), and Serrano and Hoesli (2009), the EPRA/NAREIT indices cover the largest and most heavily traded real estate stocks for each national market and thus they represent suitable benchmarks for the respective national real estate markets. From an investor's point of view, it is worth mentioning that exchange traded funds exist for at least some EPRA/NAREIT indices. Therefore, an easy-to-trade and low-cost product tracking the corresponding EPRA/NAREIT indices is available to investors. A detailed analysis of different real estate stock market indices is conducted by Serrano and Hoesli (2009). They also conclude that the EPRA/NAREIT indices are well suited for the analysis of the real estate stock markets.

The choice of the sample period from 1990 to mid of 2009 is stipulated by the availability of the data. With the exception of Italy instead of Spain, the same countries are examined as by Bond et al. (2003) and Gallo and Zhang (2009). However, Yunus (2009) and Gallo and Zhang (2009) focus on a period ending in September 2007 and August 2007, respectively. Thus, both studies do not analyze the aftermath of the financial crises which influences their empirical results at least to some extent as it is shown in this paper.

Figure 1 and Figure 2 present the logarithms of the level of the country indices whereas Figure 3 focuses on the level of the continental indices. Depicted in Figure 1, the Anglo-Saxon real estate markets (Australia, Canada, the U.K., and the U.S.) show an almost continuous upward trend from the beginning of the 1990s until mid of 2007. 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 lends preliminary support for applying cointegration analysis. The performance of the continental European real estate markets is mixed as well. While the markets moved within a range in the 1990s with the exception of the small Swedish, Italian, and the German market, this pattern changed in the second half of the period investigated. All markets are members of the European Union and are subject to the monetary policy of the European Central Bank with the exception of Sweden.

During the sample period, all European markets show a strong common upward trend and an increase in their index levels until the first half of 2007 but the extent of growth differs among the countries. The last two years of observation are characterized by a sharp decline on the securitized real estate markets and a gradual recovery since the end of 2008. This is similar to the developments on the common stock markets and suggests a close link between real estate stock markets and common stock markets at least in the short and medium run. One exception is the German real estate market. From all the European real estate stock markets, the German market suffered the most from the burst of the high-tech-bubble at the beginning of the 21st century. Furthermore, the German direct real estate market did not take part in the tremendous growth and appreciation of the last decade that took place in Ireland, Spain, the U.K., and the U.S. This last point is relevant, as real estate companies invest mainly in their domestic market and less in foreign markets. Thus, their performance is closely related to the performance of the national real estate market in the long-run.[6]

From Figure 1, Figure 2, and Figure 3 it is also evident, that the Asian markets followed a common downward trend in the aftermath of the Asian and Russian crisis in 1997 and 1998, which was more extended than for the non-Asian markets. As can be seen from Figure 3, another common development on the international real estate stock markets can be observed at the time of the international financial crisis starting in June 2007, when Bear Stearns announced serious problems with their hedge funds.

[6] See Morawski et al. (2008) and Oikarinen et al. (2009).

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

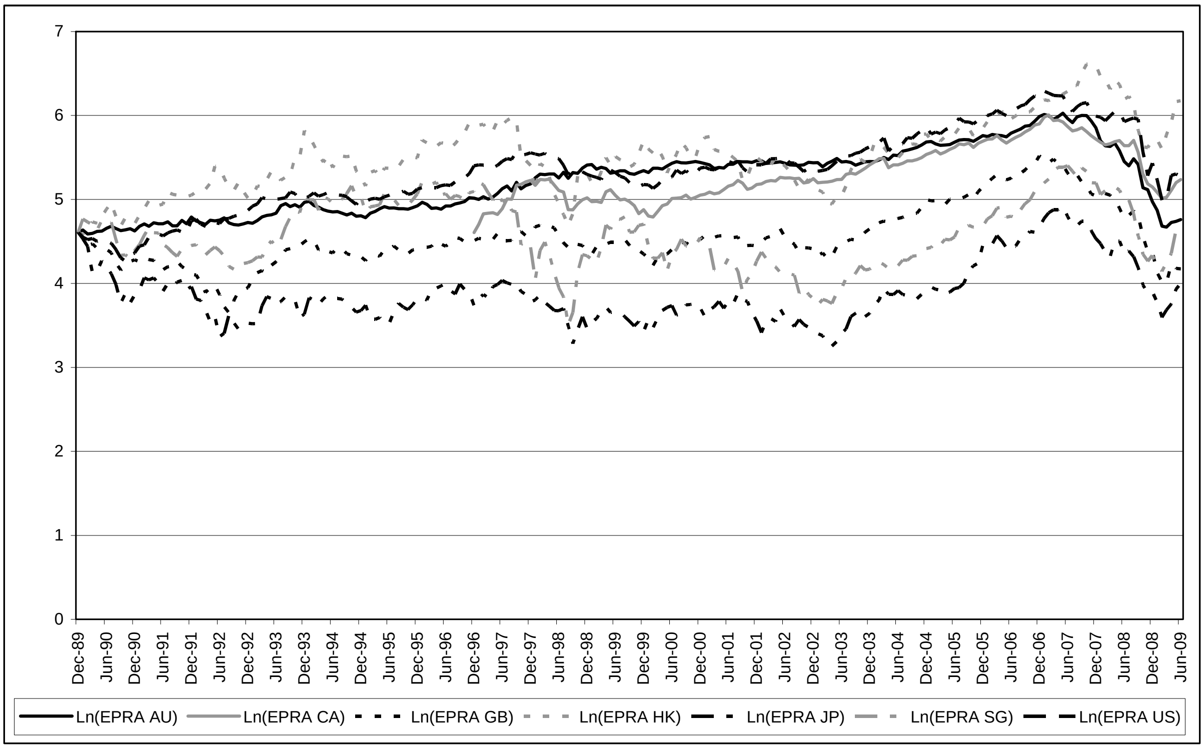

Figure 2: Price Series of the Continental European Country Indices

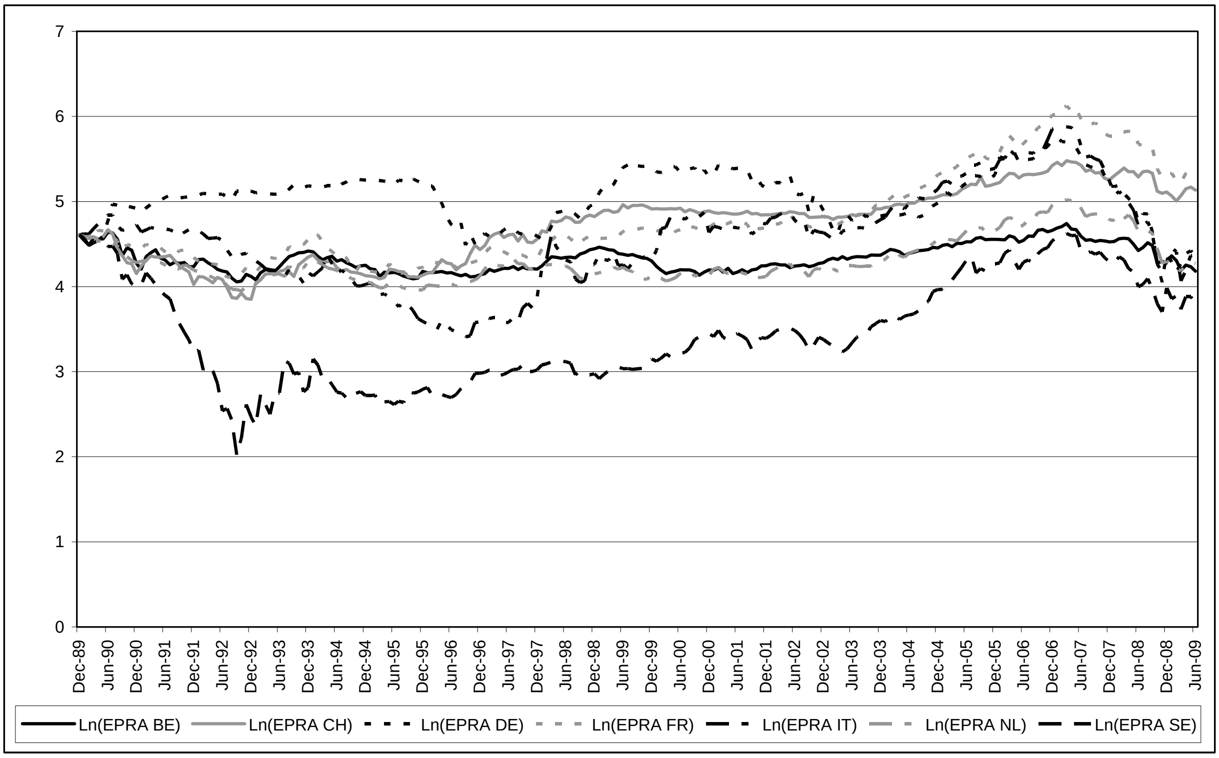

Figure 3: Price Series of the Continental Indices

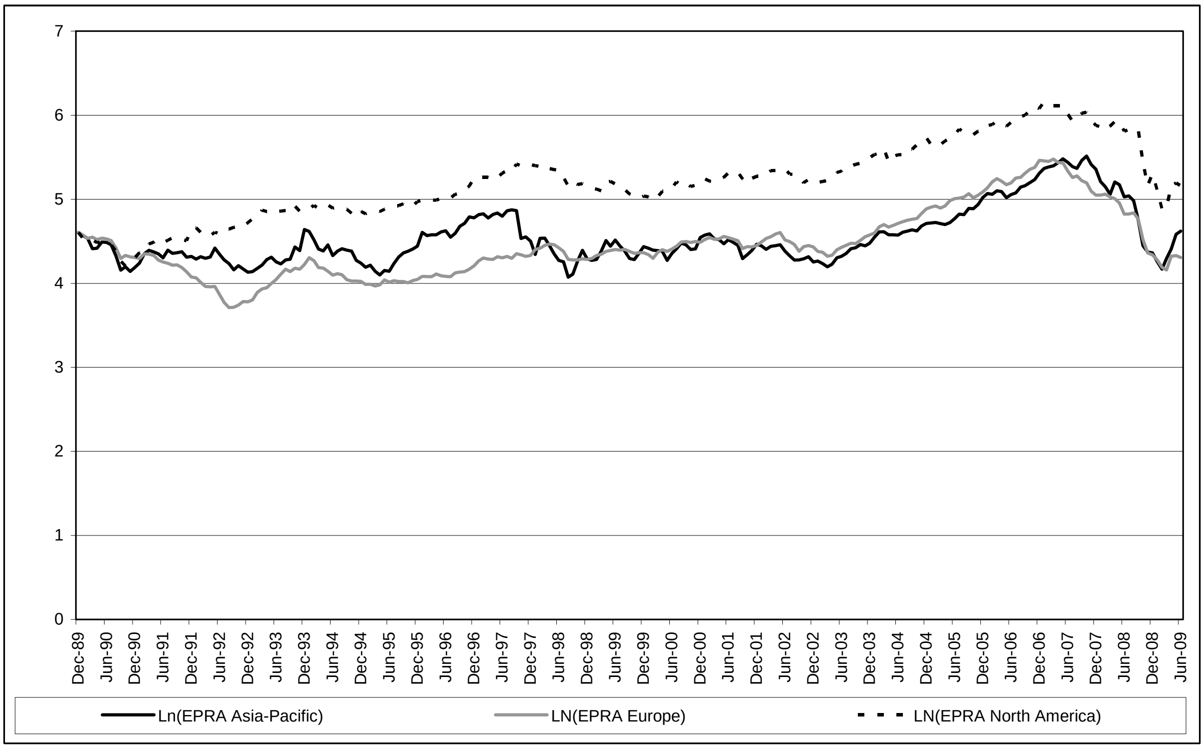

Table 1 gives an overview of the average return, risk, and distributional characteristics of the 14 national real estate stock indices and the three continental indices as well. As it can be seen, the performance of the countries' securitized real estate markets is very heterogeneous and differs substantially among national markets. While the market in Hong Kong has an average monthly return of 0.67 % and Canada, France, Switzerland, and the U.S. of around 0.30 % respectively, the Swedish market has a highly negative monthly average return of around -0.49 %. The Asian countries and the poorest performing market of Sweden are the countries with the highest standard deviation. But there has to be made one 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, where rental investments dominate.[7] The European markets of Belgium, France and the Netherlands show lower risk. But it is also evident that these European markets are the worst performing ones and have a negative average return which is mainly caused by the

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7 See Newell and Chau (1996), Liow (1997), and Serrano and Hoesli (2010) as well.

market crash in the last two years. On the other hand, the markets in Northern America show both the relative high average return and relative low risk.

According to the test statistic of the Jarque-Bera normality test, the null hypothesis of normally distributed returns is rejected for all 14 national indices at the 1 %-level of significance.[8] By applying the test suggested by Urzúa (1996), the third and fourth moment emphasize these findings. 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. Without any exception, the return distributions are significantly leptokurtic and negative skewness dominates. The German, Swedish, and the three Asian markets are the exceptions and show no significant negative skewness. However, they are also slightly negative skewed. The findings indicate that the characteristic of non-normally distributed returns is not only typical for low-capitalized and developing securitized real estate markets like the Belgian, Italian or Swedish market, but also for the high-capitalized markets with a long history like the Anglo-Saxon markets, where the Australian and US-market show extremely high negative skewness and leptokurtosis. Furthermore, the continental indices are non-normally distributed and are characterized by negative skewness and significant leptokurtosis as well. Due to the results above, the use of standard deviation as a measure of risk may result in distortions of the true performance.

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8 See Brounen et al. (2008), Liow (2007), and Liow and Sim (2006) as well.

Table 1: Descriptive Statistics of the EPRA Country and Region Indices

Index	Mean	Min.	Max.	S.D.	Skewness (z-stat.)	Kurtosis (z-stat.)	J.-B.
AU	0.0003	-0.4555	0.1223	0.0594	-2.3386 (14.7927)	17.3606 (46.3791)	2,224.0122***
BE	-0.0011	-0.3033	0.1232	0.0486	-1.2932 (8.1802)	9.6358 (21.4753)	494.5550***
CA	0.0052	-0.3935	0.1723	0.0671	-1.6857 (8.5981)	11.6653 (22.8728)	540.3380***
CH	0.0031	-0.2521	0.1792	0.0517	-0.3080 (1.9484)	5.4856 (8.0955)	63.9375***
DE	-0.0003	-0.4474	0.4288	0.0839	-0.1439 (0.9100)	9.5124 (21.0775)	414.3199***
FR	0.0037	-0.3422	0.1364	0.0569	-0.9586 (6.0636)	7.9322 (15.9831)	273.0233***
GB	-0.0020	-0.3528	0.2554	0.0647	-0.8849 (5.5973)	7.6181 (14.9703)	238.4707***
HK	0.0067	-0.4423	0.4481	0.1035	-0.0060 (0.0378)	5.7929 (9.0862)	76.0528***
IT	-0.0003	-0.4735	0.3236	0.0880	-0.3867 (2.4461)	7.5307 (14.6888)	205.9748***
JP	-0.0017	-0.3828	0.3171	0.0962	-0.0479 (0.3030)	4.0447 (3.4504)	10.7316***
NL	-0.0007	-0.2848	0.1114	0.0473	-1.0644 (6.7329)	8.2088 (16.8749)	308.7215***
SE	-0.0049	-0.4417	0.3953	0.0977	-0.1090 (0.6897)	6.7059 (12.0296)	134.3659***
SG	0.0009	-0.4193	0.5218	0.1160	-0.1327 (0.8397)	6.0658 (9.9659)	92.3260***
US	0.0028	-0.3922	0.2714	0.0603	-1.7578 (11.1188)	14.5916 (37.4522)	1,430.5714***
AP	0.0018	-0.3323	0.3199	0.0750	-0.1842 (1.1653)	6.0515 (9.9201)	92.1149***
EU	-0.0006	-0.3484	0.1620	0.0514	-1.5000 (9.4878)	11.6764 (28.0539)	821.7228***
NA	0.0024	-0.3923	0.2603	0.0593	-1.8138 (11.4732)	14.5666 (37.3716)	1,432.7299***

Notes: Min. and Max. are the minimum and maximum monthly return, 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 %-level of significance. 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 1 %-, 5 %-, and 10 %-level of significance are 2.58, 1.96, and 1.65.

5 Empirical Results

However, there are limitations in the validity of the results presented above. First, as shown above, the returns of securitized real estate markets exhibit negative skewness as well as excess kurtosis and thus they are not normally distributed. Therefore, low correlation coefficients that can be interpreted as supportive of pervasive diversification benefits may offer misleading results when applied for the purposes of portfolio optimization and investment decisions. Second, since correlation analysis is only valid for stationary variables, the prices have to be de-trended by calculating first differences. However, this procedure reduces valuable information regarding the presence of common trends in prices. While correlation is an appropriate and widely used measure of short-term co-movements, low correlation coefficients themselves do not assure that there are low long-term co-movements as well. Thus, the further examinations of this paper focus on long-term linkages between the price series of the 14 real estate indices and the dynamic interactions between these markets. Additionally, the focus is put on the inter-continental long-run dependencies as well.

5.1 Unit Root Tests

Since cointegration methodology is based on the assumption that at least two of the time series contain a unit root and that they are integrated of the same order, the analysis is started with unit root tests on the 14 national and on the 3 regional securitized real estate indices. We conduct four different unit root tests: Dickey-Fuller-Generalized-Least-Squares (DF-GLS) (Elliot et al. (1996)), Phillips-Perron (PP) (Phillips and Perron 1988), Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (Kwiatkowski et al. (1992)), and the Zivot-Andrews (ZA) (Zivot and Andrews (1992)) test. We use DF-GLS test instead of the conventional augmented Dickey-Fuller unit root test due to its superior power properties (Maddala and Kim (1998)). While the DF-GLS, PP, and ZA test statistic test the null hypothesis of a unit root, the KPSS test has a reversed null hypothesis of stationarity. We use the KPSS test for the purposes of confirmatory analysis (Maddala and Kim (1998)). By contrast to the three other well known tests, the procedure suggested by Zivot and Andrews (1992) additionally controls for a single structural break of unknown timing. It tests the null hypothesis of a unit root against an alternative hypothesis of stationarity with a single structural break of unknown timing in the parameters of the data-generating process. Two specifications are implemented. First, a model with a shift in the mean is analyzed and second, a regime shift is considered. The test statistic for the four tests is presented in Table 2.

Table 2: Unit Root Tests for Levels

Index	DF-GLS$_C$	PP$_C$	ZA$_M$	ZA$_R$	KPSS$_C$
AU	-1.3005	1.5218	-2.59	-3.63	3.0177***
BE	-1.0978	-1.7884	-3.7	-2.59	1.1999***
CA	-0.6229	-1.7459	-2.75	-2.96	2.3661***
CH	-0.2779	-0.4324	-3.71	-3.42	4.0270***
DE	-1.6294	-2.1240	-3.08	-3.57	0.2862
FR	-0.3778	-0.4675	-3.62	-2.23	3.3638***
HK	-0.2817	-2.3123	-4.93**	-4.99*	2.1792***
IT	-1.3645	-1.2584	-2.4	-2.21	2.4072***
JP	-1.1173	-2.4971	-4.23	-4.09	2.1846***
NL	-1.0592	-1.3982	-3.66	-2.31	0.9524***
SE	-0.5110	-2.1279	-3.79	-3.4	1.7232***
SG	-2.0278**	-2.1183	-3.81	-3.92	1.3874***
UK	-1.9015*	-1.4302	-2.9	-3.42	0.4334*
US	-0.6914	-1.6583	-2.89	-3.86	3.7501***
AP	-1.5540	-1.7405	-3.94	-3.97	2.2532***
EU	-1.4730	-1.2119	-3.53	-2.83	2.7142***
NA	-0.8304	-1.5148	-2.89	-4.01	3.7937***

Notes: For the DF-GLS, the PP, and the KPSS tests specifications include constants only. ZA$_M$ indicates the Zivot and Andrews (1992) test with an endogenous change in the level of the series. ZA$_R$ indicates the Zivot and Andrews test (1992) with an endogenous change in both level and trend. ***, ** and * indicate the rejection of the null hypothesis at the 1 %-, 5 %-, and 10 %-level of significance.

As we can see, the null hypothesis of a unit root is not rejected by the DF-GLS, PP, and ZA tests for the majority of the series. Hong Kong and Singapore are the exceptions: for these series the null hypothesis of a unit root is rejected at the 5 %-level of significance by the DF-GLS and ZA test procedure respectively. The KPSS test complements this finding by rejecting the null hypothesis of stationarity for all indices with the exception of Germany. When analyzing first differences, the null hypothesis is rejected for all indices by DF-GLS, PP, and ZA tests (Table 3).

Shown in Table 3 are the results for the first differences of the securitized real estate indices. The KPSS-test indicates non-stationarity in the first differences of the Swedish index only. For all the other indices the hypothesis of stationary first differences in the time series is accepted. Based on the results presented above, we conclude that the time series are integrated of order one. These results are similar to those by Yunus and Swanson (2009), Yunus (2009), and Gallo and Zhang (2009) for international securitized real estate indices. Thus, from a statistical perspective, the results from unit root tests support the implementation of cointegration methodology.

Table 3: Unit Root Tests for First Differences

Index	DF-GLS$_C$	PP$_C$	ZA$_M$	ZA$_R$	KPSS$_C$
AU	-5.0409***	-13.0880***	-14.10***	-14.05***	0.2405
BE	-2.8963***	-13.7647***	-12.55***	-12.67***	0.1778
CA	-1.9099*	-9.1627***	-9.47***	-9.85***	0.1847
CH	-6.7144***	-13.7240***	-8.64***	-8.63***	0.1978
DE	-4.7406***	-14.3032***	-14.85***	-14.89***	0.1764
FR	-5.2619***	-12.1759***	-12.43***	-12.41***	0.2467
HK	-6.3007***	-12.8436***	-13.05***	-13.04***	0.0474
IT	-4.5421***	-12.4900***	-10.36***	-6.30***	0.1669
JP	-3.1942***	-14.4946***	-14.91***	-14.94***	0.1628
NL	-5.4681***	-11.5740***	-12.13***	-12.10***	0.2464
SE	-5.0722***	-15.4491***	-9.16***	-13.97***	0.6268**
SG	-2.2734**	-13.0575***	-13.33***	-13.38***	0.0597
UK	-3.9030***	-11.3862***	-12.40***	-12.36***	0.1713
US	-2.7756***	-12.9200***	-6.84***	-6.78***	0.2402
AP	-4.0779***	-12.8844***	-13.23***	-13.23***	0.0549
EU	-4.3117***	-11.0115***	-12.13***	-12.09***	0.2084
NA	-2.2999**	-12.6427***	-6.90***	-6.85***	0.1967

Notes: For the DF-GLS, the PP, and the KPSS tests specifications include constants only. ZA$_M$ indicates the Zivot and Andrews (1992) test with an endogenous change in the level of the series. ZA$_R$ indicates the Zivot and Andrews test (1992) with an endogenous change in both level and trend. ***, ** and * indicate the rejection of the null hypothesis at the 1 %-, 5 %-, and 10 %-level of significance

5.2 Cointegration without Structural Breaks

As follows from the results of the unit root tests, all securitized real estate markets are integrated of the same order which is essential for estimating the cointegration vectors. Thus, the prerequisites for applying cointegration methodology are fulfilled. Before testing for time-varying cointegration and structural breaks in the cointegration vectors, we apply the bivariate cointegration methodology suggested by Engle and Granger (1987), the multivariate cointegration framework proposed by Johansen (1988). The results on these tests can be seen as a starting point for analyzing time-varying cointegration and can be compared with those of the more recent tests suggested by Gregory and Hansen (1996) and McCabe et al. (2003) as well as recursive cointegration methodology (see section 5.3). The tests have been conducted using both domestic currencies and US-dollar. However, since the results are qualitatively the same for both currencies, we report the results of the test in US-dollars only.

5.2.1 Engle and Granger (1987) Test for Bivariate Cointegration

The first step of the pair-wise cointegration test proposed by Engle and Granger (1987) implies the estimation of the OLS regression of logarithmic securitized 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, test results should the same both when variable x_{2t} is regressed on x_{1t} and when variable x_{1t} is regressed on x_{2t}. However, since the literature reports that differing results do emerge when using empirical data are used, we estimate each regression is run in both directions.

The methodology chosen for the unit root test of the residuals from the OLS regression is equivalent to the ADF-test with one exception. Instead of using the critical values of MacKinnon (1996), the critical values of MacKinnon (1991) are applied. The unit root tests are conducted for specifications including a constant, as well as both a constant and a trend component. The rejection of the null hypothesis of a unit root of the residuals indicates that the two time series are cointegrated.

The results are presented for the following regional groups of markets: North-American markets (Canada and the US); European markets (Germany, France, the Netherlands, Belgium, Italy, Sweden, Switzerland, and the UK); and Asia-Pacific markets (Australia, Japan, Hong Kong, and Singapore). For the purpose of multivariate tests, European markets are further sub-divided into so-called core European markets (Germany, France, the Netherlands, and the UK) and periphery European markets (Belgium, Italy, Sweden and Switzerland). This classification of the European markets is based on considerations of market size and findings related to the importance of the core markets for the periphery European markets in previous studies such as Schindler (2010). Grouping markets by region allows us to investigate diversification benefits available to investors who seek diversification benefits within a given region. To take into account the perspective of the US-based investor, we also consider portfolios consisting of the US market index and indices of the other regions (i.e., US and European indices, US and Asia-Pacific indices). Finally, we present results from the stand point of internationally diversified investors by considering a group of three regional indices: Asia-Pacific, European, and North-American regional indices available from EPRA/NAREIT. All in all, we consider six major groups of markets. This grouping of markets is maintained across all the following sub-sections. Such grouping of markets also enables us to draw comparison with the earlier literature in this area.

North-American markets

Table 4 presents the test statistic from the Engle and Granger (1987) bivariate cointegration tests for the US and other markets. Considering the sample period, we find no cointegration between the two neighboring real estate stock markets in North America, even though, the economy of the two countries is strongly linked and previous studies conducted by Gallo and Zhang (2009) and Schindler (2010) detected common stochastic trends and stable long-run relationships. However, the studies differ by sample period, by methodology, and by the real estate indices. Thus, the results are not directly comparable to the ones in this paper.

US and European markets

As presented in Table 4, the evidence of cointegration between national European markets and the US market is quite limited. Cointegration relationships are identified for the French, Italian, Swedish, and Swiss securitized real estate market. With the exception of the French market, the markets and economies are relatively small, probably not covered by and interesting for international investors and unable to develop their own driving forces as it might be the case for the UK. Furthermore, the test statistic is only significant at the 5% and the 10% level, respectively. The weak linkages between these two regions on a national market level are also documented by Schindler (2010) and Yunus (2009).

US and Asia-Pacific markets

We find no cointegration between the US securitized real estate market and the markets in Asia. However, the Asian financial crisis, the openness of the markets, and the fast growing markets in the Asia-Pacific region during the last 20 twenty years might be responsible for finding no stable long-run relationship between these markets and the world's largest securitized real estate in the US (Table 4). The finding is confirmed by Yunus and Swanson (2007). They do not identify any cointegration between Asia-Pacific markets and the US between 2000 and 2006, but between the relatively short period between 2003 and 2006. This is further evidence for the hypothesis that the dramatic changes in the Asia-Pacific region during the last two decades might have resulted in structural breaks in the cointegration relations which are not detected by the Engle and Granger (1987) methodology.

Table 4: Results from Engle and Granger (1987) Bivariate Cointegration Tests between National Securitized Real Estate Markets and the US

Indices		Unit root tests in regression residuals	
Endogenous variable	Exogenous variable	ADF_C	ADF_T
US and Canada			
CA	US	-0.8471 (2)	-2.0673 (4)
US	CA	-1.2028 (4)	-1.6713 (4)
US and European Markets			
BE	US	-2.1226 (1)	-2.1204 (1)
US	BE	-2.9719 (12)	-2.4120 (12)
CH	US	-1.8790 (13)	-3.4257 (13)
US	CH	-3.5010 (12)**	-3.1295 (12)
DE	US	-2.8018 (6)	-2.8836 (6)
US	DE	-1.6491 (3)	-2.5890 (3)
FR	US	-1.2042 (3)	-1.1839 (0)
US	FR	-3.2570 (12)*	-2.1604 (12)
IT	US	-1.6863 (1)	-1.7708 (1)
US	IT	-3.7444 (12)**	-2.8590 (12)
NL	US	-2.1209 (0)	-2.1413 (1)
US	NL	-2.9396 (6)	-0.8984 (6)
SE	US	-2.7383 (3)	-3.3072 (3)
US	SE	-3.3390 (11)*	-2.2776 (11)
UK	US	-2.4051 (1)	-2.3522 (1)
US	UK	-2.7436 (0)	-2.6496 (0)
US and Asia-Pacific Markets			
AU	US	-2.3812 (3)	-2.3942 (3)
US	AU	-1.9687 (1)	-2.2299 (1)
HK	US	-2.9394 (1)	-2.9323 (1)
US	HK	-1.9625 (0)	-1.8744 (0)
JP	US	-3.0013 (0)	-2.8825 (0)
US	JP	-2.5486 (0)	-1.3387 (0)
SG	US	-2.1691 (1)	-2.1413 (1)
US	SG	-2.4037 (11)	-1.9282 (11)

Notes: ADF_C and ADF_T denote the values of the ADF test with constant and with constant and trend respectively. The lag lengths for unit root test of the regression residuals are given in parentheses. Approximate critical values for ADF-tests are taken from MacKinnon (1991). ***, **, and * indicate the rejection of the null hypothesis of a unit root at the 1 %-, 5 %-, and 10 %-level of significance. Cells shaded in grey indicate the presence of a cointegration relationship at at least 5% level between the given two markets.

European markets

For the European markets we find weak evidence of cointegration among national securitized real estate markets (Table 5). With the exception of the linkages between the Belgian and Dutch market, all further cointegration relationships are related to the Swedish market. This market shows long-run co-movements with the markets in Germany, France, and Italy in the Euro area as well as with the Swiss market. However, the results should be treated carefully according to the unit root test results and the results from the KPSS test above since there is weak evidence that the Swedish securitized real estate market index is stationary in levels. Schindler (2010) finds similar evidence and thus, excluded Sweden from further cointegration analysis. A strong long-run relationship exists between the two neighboring markets of Belgium and the Netherlands which is also well documented in the relevant literature such as Gallo and Zhang (2009) and Schindler (2010) and expected from an economic point of view. Notably, there are no further long-run linkages between the European markets in the background of its economic relationships, their geographical vicinity, and there common currency, at least in five of the eight analyzed markets. Yang et al. (2005) conclude that real estate market integration increased slightly among members of the European Monetary Union with advanced industrial structures after its establishment. This finding can be indicative of a structural break and therefore provides another incentive for applying time-varying cointegration analysis and tests for structural breaks.

Table 5: Results from Engle and Granger (1987) Bivariate Cointegration Tests between European Securitized Real Estate Markets

Indices		**Unit root tests in regression residuals**	
Endogenous variable	Exogenous variable	ADF_C	ADF_T
BE	CH	-2.0312 (0)	-2.0161 (0)
CH	BE	-0.6400 (1)	-2.2690 (1)
BE	DE	-1.9701 (0)	-2.5117 (0)
DE	BE	-3.0114 (7)	-2.0164 (0)
BE	FR	-2.3578 (14)	-2.6709 (14)
FR	BE	-0.6193 (13)	-3.1320 (14)
BE	IT	-2.2253 (1)	-2.1817 (1)
IT	BE	-1.8451 (2)	-1.9338 (2)
BE	NL	-3.8042 (0)**	-4.0827 (10)**
NL	BE	-3.5660 (0)**	-4.6750 (12)***
BE	SE	-2.5527 (1)	-2.5046 (1)
SE	BE	-2.4734 (0)	-2.6460 (0)
BE	UK	-2.2191 (1)	-2.0735 (1)
UK	BE	-2.2253 (1)	-1.6810 (1)
CH	DE	0.1413 (0)	-3.3346 (0)
DE	CH	-2.8891 (5)	-2.9163 (5)
CH	FR	-1.5243 (2)	-1.8783 (2)
FR	CH	-1.5506 (2)	-1.5750 (2)
CH	IT	-0.6421 (0)	-2.0786 (1)
IT	CH	-1.4600 (1)	-1.5291 (1)
CH	NL	-0.6969 (3)	-2.0001 (0)
NL	CH	-1.7935 (1)	-1.7339 (1)
CH	SE	-1.6737 (8)	-2.0796 (8)
SE	CH	-3.2934 (8)*	-3.0748 (8)
CH	UK	0.4619 (0)	-0.7880 (0)
UK	CH	-1.2256 (2)	-1.3385 (2)
DE	FR	-2.7923 (6)	-2.9699 (6)
FR	DE	-0.4219 (6)	-2.9412 (6)
DE	IT	-2.0683 (0)	-2.3680 (0)
IT	DE	-1.3618 (0)	-2.5121 (0)
DE	NL	-3.0070 (6)	-3.0348 (6)

Table 5 continues on the next page

Indices		Unit root tests in regression residuals	
Endogenous variable	Exogenous variable	ADFC	ADFT
NL	DE	-1.9538 (12)	-2.7835 (1)
DE	SE	-3.1444 (5)*	-3.0364 (5)
SE	DE	-2.2573 (8)	-3.8317 (3)**
DE	UK	-2.2303 (0)	-2.5201 (0)
UK	DE	-2.1325 (8)	-2.6056 (8)
FR	IT	-0.8339 (0)	-1.6740 (0)
IT	FR	-1.5622 (1)	-1.6868 (1)
FR	NL	-1.9381 (0)	-2.8814 (0)
NL	FR	-2.3784 (0)	-2.5044 (1)
FR	SE	-2.8712 (5)	-2.9914 (12)
SE	FR	-3.9024 (3)**	-3.4733 (3)
FR	UK	0.1557 (1)	-0.6298 (1)
UK	FR	-1.9066 (9)	-0.7969 (1)
IT	NL	-1.8861 (1)	-1.7503 (1)
NL	IT	-1.9745 (1)	-2.0363 (1)
IT	SE	-2.8640 (0)	-2.4282 (0)
SE	IT	-3.3886 (3)**	-3.1757 (3)
IT	UK	-1.7900 (1)	-1.8604 (1)
UK	IT	-1.8604 (1)	-1.7597 (1)
NL	SE	-2.3962 (0)	-2.2838 (0)
SE	NL	-3.0349 (8)	-2.8985 (5)
NL	UK	-1.2390 (0)	-0.4279 (6)
UK	NL	-1.1317 (0)	-0.4389 (0)
SE	UK	-2.4095 (9)	-1.3644 (9)
UK	SE	-2.4142 (9)	-2.6878 (6)

Notes: ADF_C and ADF_T denote the values of the ADF test with constant and with constant and trend respectively. The lag lengths for unit root test of the regression residuals are given in parentheses. Approximate critical values for ADF-tests are taken from MacKinnon (1991). ***, **, and * indicate the rejection of the null hypothesis of a unit root at the 1 %-, 5 %-, and 10 %-level of significance. Cells shaded in grey indicate the presence of a cointegration relationship at at least 5% level between the given two markets.

Asia-Pacific markets

Considering the Asia-Pacific markets, the Japanese market shows cointegration relationships with all three markets within this region, namely Australia, Hong Kong, and Singapore. However, there is no further long-run relationship between the other markets. While the cointegration relationships between the Japan and Hong Kong as well as between Japan and Singapore are highly significant the third identified long-run co-movement between Japan and Australia is statistically much weaker as can be seen from Table 6. This result 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 Hong Kong and Singapore, and have shown a more stable performance for the last 20 years. Furthermore, the Australian economy is less focused on the banking and industrial sector than Hong Kong, Singapore, and Japan and more related to and dependent on the influence by the mining sector. The findings from previous analysis of long-run co-movement between Asian real estate markets are mixed. While Garvey et al. (2001) find limited cointegration between Asia securitized real estate markets, more recent studies conducted by Gallo and Zhang (2009) as well as Schindler (2009b, 2010) find much stronger evidence of long-run co-movement between Asian real estate markets which is in support of the hypothesis that the markets have become more developed for the last two decades and thus, more integrated.

Table 6: Results from Engle and Granger (1987) Bivariate Cointegration Tests between Asia-Pacific Securitized Real Estate Markets

Indices		Unit root tests in regression residuals	
Endogenous variable	Exogenous variable	ADF_C	ADF_T
AU	HK	-0.9413 (0)	-05845 (0)
HK	AU	-2.2672 (1)	-2.3837 (1)
AU	JP	-2.2556 (0)	-0.9035 (0)
JP	AU	-3.0725 (0)*	-2.8488 (0)
AU	SG	-1.7609 (4)	-1.4859 (4)
SG	AU	-2.1131 (1)	-2.0688 (1)
HK	JP	-4.4757 (1)***	-5.3775 (1)***
JP	HK	-4.3023 (0)***	-4.1467 (0)**
HK	SG	-2.0616 (3)	-5.1912 (1)***
SG	HK	-2.2462 (1)	-2.3542 (1)
JP	SG	-3.7768 (0)**	-4.8542 (0)***
SG	JP	-1.9647 (9)	-3.9706 (0)**

Notes: ADF_C and ADF_T denote the values of the ADF test with constant and with constant and trend respectively. The lag lengths for unit root test of the regression residuals are given in parentheses. Approximate critical values for ADF-tests are taken from MacKinnon (1991). ***, **, and * indicate the rejection of the null hypothesis of a unit root at the 1 %-, 5 %-, and 10 %-level of significance. Cells shaded in grey indicate the presence of a cointegration relationship at at least 5% level between the given two markets.

Regional indices

The weak evidence of cointegrated national securitized real estate markets across geographical and economic regions from above is confirmed by the cointegration analysis based on the three regional indices for the Asia-Pacific region, Europe, and North America (Table 7). We only find one statistically significant cointegration relationship between Europe and North America while the Asia-Pacific region is neither integrated with the European securitized real estate market nor with the securitized real estate market in North America. The general finding that cointegration exists within economic regions but much weaker or not at all across international real estate markets is in line with the literature such as Gallo and Zhang (2009), Schindler (2010), and Yunus (2009).

Thus, investors in the securitized real estate market are advised to invest in international real estate and to complement their investments into local or regional assets by more global ones.

Table 7: Results from Engle and Granger (1987) Bivariate Cointegration Tests between Regional Securitized Real Estate Markets

Indices		**Unit root tests in regression residuals**	
Endogenous variable	Exogenous variable	ADF_C	ADF_T
AP	EU	-2.3750 (0)	-2.4041 (0)
EU	AP	-1.9930 (0)	-2.1024 (0)
AP	NA	-2.4391 (0)	-2.1342 (4)
EU	AP	-2.3070 (0)	-2.0219 (0)
EU	NA	-2.5894 (4)	-2.6763 (4)
NA	EU	-5.0850 (12)***	-4.3967 (12)***

Notes: ADF_C and ADF_T denote the values of the ADF test with constant and with constant and trend respectively. The lag lengths for unit root test of the regression residuals are given in parentheses. Approximate critical values for ADF-tests are taken from MacKinnon (1991). ***, **, and * indicate the rejection of the null hypothesis of a unit root at the 1 %-, 5 %-, and 10 %-level of significance. Cells shaded in grey indicate the presence of a cointegration relationship at at least 5% level between the given two markets.

5.2.2 Johansen and Juselius (1990) Test for Multivariate Cointegration

To investigate the presence of cointegration relations among several securitized real estate markets, we start with the multivariate cointegration test suggested by Johansen (1988) and Johansen and Juselius (1990) tests, traditionally applied for measuring long-term market co-movements, before proceeding with the use of more recent cointegration tests. The results are presented for the same groups of markets as described in the previous section but in a multivariate framework. For comparison with the previous literature, we also perform Johansen and Juselius (1990) cointegration test for all the market groups for the period finishing in June 2007. The results are not provided here for the sake of space but are available from authors upon request.

As presented in the methodology section, to identify the deterministic component of the VAR models we follow the approach proposed in Juselius (2007) and to determine the optimal lag length by the Akaike information criterion (AIC). When determining cointegration rank of the system we place more weight to the results of the trace test. This test has been shown to be more robust to skewness and excess kurtosis in the residuals than the maximum eigenvalue test (Cheung and Lai (1995)).

North-American markets

We find no cointegration relations between the US and the Canadian securitized real estate indices. This result contradicts that of Gallo and Zhang (2009) who find a cointegration relation between these two markets over the period from 1992 to 2007. However, it should be pointed out that Gallo and Zhang (2009) study a shorter period and this may be a reason behind the difference. As the results in Table 8 show, this is indeed the case. The results of the cointegration test for these two markets are therefore not robust with regard to the sample period. Among the possible reasons that may explain the unstable pattern of this cointegration relation may be the financial crisis that started in August 2007. The crisis events that started with the instability in the sub-prime sector of the US real estate market might have disrupted the long-run relationship between these two markets. In the later sections we will see whether the detected instability in this relationship is driven by this single event only or whether it is suggestive of multiple incidences of instability present in this relationship.

Table 8: Results from Johansen and Juselius (1990) Multivariate Cointegration Tests between Canada and the US

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.0645	12.5625	0	1	9.7441
1	1	0.0191	2.8211	1	2	9.1645

Notes: * indicates significance at the 5% level.

US and European markets

The US real estate securities and securities of the European markets, both core and periphery markets show no cointegration relations (Tables 9-11). However, when all European markets are included in the group along with the US market, there is an evidence of two cointegration relations (Table 11). These results hold for the whole sample and for the pre-crisis period finishing in June 2007. Gallo and Zhang (2009) also find limited evidence of cointegration between the US and EU markets in that they report a single relationship between the US and seven EU markets within a bivariate set-up. As presented in the previous section, we find four cointegration relations when we apply Engle and Granger (1987) test to the group consisting of the US and all European markets. Since the markets share long-run stochastic trends, it suggests limited diversification potential.

Table 9: France, Germany, Netherlands, U.K., US: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.0941	58.1099	0	1	23.0283
1	1	0.0685	35.0816	1	2	16.5425
2	2	0.0361	18.5391	2	3	8.5721
3	3	0.0261	9.9670	3	4	6.1645
4	4	0.0162	3.8026	4	5	3.8026

Notes: * indicates significant at the 5% level.

Table 10: Belgium, Italy, Sweden, Switzerland, US: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.1220	69.4575	0	1	30.0535
1	1	0.0789	39.4040	1	2	18.9882
2	2	0.0412	20.4158	2	3	9.7667
3	3	0.0348	10.6492	3	4	8.1803
4	4	0.0106	2.4689	4	5	2.4689

Notes: * indicates significant at the 5% level.

Table 11: All European Markets and US: December 1996 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.2677	251.0076*	0	1	72.8989*
1	1	0.2235	178.1087*	1	2	59.1859*
2	2	0.1600	118.9228	2	3	40.8011
3	3	0.0931	78.1218	3	4	22.8610
4	4	0.0703	55.2608	4	5	17.0677
5	5	0.0622	38.1931	5	6	15.0272
6	6	0.0449	23.1660	6	7	10.7486
7	7	0.0332	12.4173	7	8	7.8886
8	8	0.0192	4.5288	8	9	4.5288

Notes: * indicates significant at the 5% level.

US and Asia-Pacific markets

In case of the US and Asian markets the results of the trace and maximum eigenvalue test contradict each other. While maximum eigenvalue test indicates presence of one cointegration relationship among these markets during the whole sample period, the trace test does not suggest that the relationship is present. This result holds in the pre-crises sample as well as for the whole sample. As was mentioned at the start of the section, we give more weight to the trace test, which seems to be more robust to the presence of nonnormality. We therefore conclude that these five markets are not cointegrated, at least when cointegration is measured by means of the Johansen cointegration test. (Table 12). Long-run co-movements between the US and Asian-pacific markets have been previously investigated by Garvey et al. (2001), Yunus and Swanson (2007) and Schindler (2009b). These studies arrive at the same result in that they find limited evidence of cointegration. These findings are suggestive of the diversification benefits for the US investors seeking to diversify into Asian markets. We will see whether this conclusion holds when the assumption of time-invariability of cointegration relations is relaxed.

Table 12: Australia, Hong Kong, Japan, Singapore, US: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.1560	68.8552	0	1	39.5043*
1	1	0.0635	29.3509	1	2	15.2971
2	2	0.0325	14.0538	2	3	7.6906
3	3	0.0214	6.3631	3	4	5.0336
4	4	0.0057	1.3295	4	5	1.3295

Notes: * indicates significant at the 5% level.

European markets

There is an extensive literature on increasing integration among the Western European stock markets (Rangvid (2001), Aggarwal et al. (2005), Hardouvelis et al. (2006)). Using a battery of statistical tests, Lizieri et al. (2003) find that European real estate markets show markedly less integration than the European equity markets. They explain it by the strong domestic focus of the real estate companies' portfolios, resulting in informational asymmetries and high information costs, and the relatively smaller size of the real estate securities markets. Liow et al. (2005) find weak evidence in favour of one cointegration relationship among France, Italy, Germany and the UK over the period from 1993 to 2003. Our longer sample may help to identify whether the integration processes have accelerated over the last decade. We find that European core real estate securities markets (Germany, France, the UK and the Netherlands) show no equilibrium relationships either before the last financial crisis or in the whole sample period (Table 13). Therefore we find less evidence of integration than Liow et al. (2005). The same result holds for the group including the periphery European markets (Belgium, Italy, Sweden and Switzerland, (Table 14)). These findings suggest presence of intra-regional diversification benefits within the two groups of European markets.

Table 13: France, Germany, Netherlands, U.K.: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.0569	31.8319	0	1	13.6430
1	1	0.0464	18.1889	1	2	11.0680
2	2	0.0185	7.1209	2	3	4.3441
3	3	0.0118	2.7768	3	4	2.7768

Notes: * indicates significant at the 5% level.

Table 14: Belgium, Italy, Sweden, Switzerland: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.0766	38.3702	0	1	18.4055
1	1	0.0491	19.9647	1	2	11.6182
2	2	0.0249	8.3465	2	3	5.8283
3	3	0.0108	2.5173	3	4	2.5183

Notes: * indicates significant at the 5% level.

When all European markets are considered as a group, we find evidence in favour of a single cointegration relation at the 5% level of significance (Table 15). The relationship is stable as it also existed in the pre-crisis period. In fact, during the pre-crisis period we find evidence of two cointegration relations among the eight European markets, albeit at 10% level of significance only. Yang et al. (2005) also find one cointegration relation in a sample of European markets. Schindler (2010) finds that European markets share two cointegration relations.

Table 15: All European Markets: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.2242	174.7961*	0	1	59.3928*
1	1	0.1675	115.4033	1	2	42.9021
2	2	0.0787	72.5012	2	3	19.1821
3	3	0.0703	53.3192	3	4	17.0515
4	4	0.0584	36.2677	4	5	14.0794
5	5	0.0455	22.1883	5	6	10.8945
6	6	0.0306	11.2937	6	7	7.2664
7	7	0.0171	4.0273	7	8	4.0273

Notes: * indicates significant at the 5% level.

Asia-Pacific markets

Johansen cointegration test fails to find any relationships for the group of the four Asia-pacific markets in both whole (Table 16). This finding also holds for the pre-crisis period. The lack of cointegration among the Asia-Pacific markets is in line with the earlier results in Garvey et al. (2001). Garvey et al. (2001) studied the Asia-Pacific markets over the period from 1975 through 2001. Our results suggest that over the decade that followed the end of sample used by Garvey et al. (2001), the integration process in these markets has not accelerated, at least when measured by the Johansen cointegration test. The later study by Schindler (2009b) investigates bivariate cointegration relations among the Asia-Pacific markets by means of Engle and Granger (1987) cointegration tests. Their results are therefore not directly comparable to the results presented in this section due to differences in the methodology.

Table 16: Australia, Hong Kong, Japan, Singapore: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.1147	42.8420	0	1	28.5029
1	1	0.0364	14.3392	1	2	8.6882
2	2	0.0184	5.6509	2	3	4.3485
3	3	0.0056	1.3025	3	4	1.3025

Notes: * indicates significant at the 5% level.

All markets

Finally, we consider all markets in one group, taking the point of view of an investor interested in broad international diversification. For the purposes of the estimations, we exclude Canada from the group since the data for the Canadian index is only available starting from December 1996. We find that in the group of all markets excluding Canada, the diversification benefits are less likely to be expected, as we find five cointegration relations at the 5% significance level during the whole period (Table 17). For the pre-crisis period, we find six cointegration relations at the 5% significance level. However, it should be pointed out, that the results of the trace and maximum eigenvalue statistic contract each other for this group of markets. Namely, the maximum eigenvalue test fails to reject the null of no cointegration relations for this group of markets. Yunus (2009) studied a much smaller group of international markets from Asia-Pacific, Europe and North America (seven markets in total). She finds one cointegration relation in her sample (January 1990 - August 2007), with the second relation emerging after 2003.

Based on this finding, Yunus (2009) concluded that real estate securities markets move towards more integration. However, her sample period does not include the recent financial crisis period. When the recent financial crisis period is included, we observe a slight decline in the number of cointegration relations from six to five when we use trace test and we find no relationships at all when maximum eigenvalue test if used. We therefore cannot state that we obtained unequivocal evidence in favour of increasing integration for this group of markets

Table 17: All Markets excluding Canada: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.3117	496.6798	0	1	87.3940
1	1	0.2648	409.2858*	1	2	71.9676
2	2	0.2391	337.3182*	2	3	63.9337
3	3	0.2331	273.3845*	3	4	62.1152
4	4	0.1836	211.2693*	4	5	47.4708
5	5	0.1743	163.7985	5	6	44.8083
6	6	0.1587	118.9902	6	7	40.4425
7	7	0.1110	78.5478	7	8	27.5386
8	8	0.0796	51.0092	8	9	19.4104
9	9	0.0461	31.5988	9	10	11.0490
10	10	0.0390	20.5498	10	11	9.3035
11	11	0.0257	11.2463	11	12	6.0904
12	12	0.0218	5.1560	12	13	5.1560

Notes: * indicates significant at the 5% level.

Regional indices

Previous studies (Gallo and Zhang (2009)) have considered inter-regional integration using the region-wide real estate securities indices. For the purposes of comparison with this study, we also conduct the tests for the set of the three regional indices: Asia-Pacific, Europe and North America. When the three indices are considered in a group, we find no evidence of cointegration relations (Table 18), which is in line with the results of Gallo and Zhang (2009).

Table 18: Asia-Pacific, Europe, North America: December 1989 – June 2009

H0: Rank=r	H1: Rank>r	Eigenvalue	Trace	H0: Rank=r	H1: Rank=r+1	Max Eigenvalue
0	0	0.0765	29.6392	0	1	18.5326
1	1	0.0316	11.1067	1	2	7.4919
2	2	0.0154	3.6147	2	3	3.6147

Notes: * indicates significant at the 5% level.

In conclusion, we would like to point out several issues. When we apply conventional Johnasen cointegration test we find no evidence of integration within the three considered regions, with only some limited integration in the group of the eight EU markets and the group of all EU markets and the US market. We check the robustness of the results with regard to the sample period and also estimate all the tests for the pre-crisis period, which we define as December 1989 - June 2007. The only exception is that we find some evidence of lower number of cointegration relations for the eight EU markets and the group of all EU and US market. This result suggests that the recent financial turmoil might have had a disrupting effect on the integration process of some of these markets. However, in turn this may imply increased diversification benefits for these markets that have been affected. Based on the results of the Johansen cointegration test we conclude that the real estate securities markets have not been following the path of increased integration as seems is shown to be the case with the international non-real-estate equity and bonds. Therefore, these markets may still offer diversification benefits to the international investors. In the next sections we explore whether these findings hold when we relax some of the assumptions of the conventional cointegration tests, such as time-invariability and homoscedasticity.

5.3 Cointegration Tests: Structural Breaks and Stochastic Cointegration

5.3.1 Gregory and Hansen (1996) Test

To investigate the hypothesis that the results of the Engle and Granger (1987) tests may be affected by the presence of a structural break and may falsely indicate lack of equilibrium relationships when in fact such relationships are present, we use the cointegration test by Gregory and Hansen (1996). The latter test accounts for a presence of a single structural break of unknown timing in an equilibrium relationship. It tests the null hypothesis of no cointegration against an alternative hypothesis of cointegration with structural break in the parameters. The advantage of the Gregory and Hansen (1996) test is that the timing of the structural change under the alternative hypothesis is estimated from the model rather than arbitrarily set by the researcher. In the latter case the date of the break point often tied to certain economic or historical event. While the latter method has the advantage that the break point defined in this way can be easily interpreted, it may not necessarily coincide with that estimated endogenously and thus may not correctly reflect the true data process.

The Gregory and Hansen (1996) test allows for the following alternative specifications in the parameters of the cointegration relation: (1) *level shift:* change in intercept only; (2) *trend shift:* change in the slope coefficient only; (3) *regime shift:* change in both intercept and slope.[9] For each pair of market indices all three model specifications are estimated. Within each of the three model specifications we consider evidence provided by the three unit root tests: Augmented Dickey-Fuller and two Phillips-Perron tests, denoted as *ADF*, *PP Z(a)* and *PP Z(t)* in the tables. The tests are described in more detail earlier in Section 3.2.3. Like in case of Engle and Granger test described above, when testing for cointegration between two markets, we address a problem of cointegration vector normalization that arises within the bivariate approach by testing the hypothesis of no cointegration in both directions. That is, for each pair of markets two relations are considered: with each of the two variables being dependent variable in the first relation and independent variable in the second relation. To conclude whether the cointegration relation is present, we rely on the significance level of the test statistic. In case of test statistic being significant at 5% or 1% level, we conclude in favour of the presence of the cointegration relation. A feature of the Gregory and Hansen (1996) test is that the dates of the estimated break may depend on the choice of the dependent and independent variable, on the specification model and on the unit root test applied to the residuals of the cointegration regression. However, in our view identifying the break date endogenously is still preferable to running into risk of imposing wrong break date. We select the number of lags using Hall (1994) general-to-specific approach with the maximum number of lags defined using Schwert (1989) method as recommended in Maddala and Kim (1998).

North-American markets

The null hypothesis of no cointegration is rejected by the Gregory and Hansen (1996) test for the US and Canadian real estate markets in both directions. As the Table 19 shows, the null hypothesis is rejected for all three model specifications in case of US market as independent variable at at least 5% significance level and it is rejected at 1% significance level for trend change specification when Canadian market is independent variable. For the Canada – US relationship, with Canada being dependent variable, the estimated break date for the most general specification, that is, regime change, the break date is estimated

[9] The GAUSS code for Gregory and Hansen (1996) test is available from the web-site of Bruce E. Hansen at http://www.ssc.wisc.edu/~bhansen/. The estimations were performed using GAUSS 9.0 software.

to be May 2005 (according to ADF test) and August 2007 (according to PP test). For the US – Canada relationship the strongest evidence against the null is provided by the trend change model.

The estimated break dates are August 2000 (according to ADF test) or June 2000 (according to PP tests). It is worth pointing out that the results of Engle and Granger (1987) test do not support existence of the equilibrium relationship between these two markets. We suggest that it may be to the presence of the structural break in the relation. Finding cointegration relation between the US and Canada is in line with the result of Gallo and Zhang (2009) who find a cointegration relation between these two markets over the period 1992-2007 using Johansen and Juselius cointegration test.

US and European markets

We find only very weak evidence in favour of bivariate cointegration between the US and European markets in our sample (Table 20). The null of no cointegration between the UK and the US is rejected at 10% level in case of one test for trend change and regime change model. The estimated break dates are March 1994 and July 1998, respectively. In case of Italy and Sweden the evidence is even weaker: the null is rejected at the 10% by a single unit root test in case of level change model. This points out to a lack of bivariate equilibrium relationships among the considered markets and indicates a presence of potential diversification benefits for the US investors seeking to invest in considered markets. While our sample is somewhat longer than that used in Gallo and Zhang (2009) our finding of weak cointegration relation among the US and eight EU markets within a bivariate set-up is similar to theirs.

US and Asia-Pacific markets

Table 21 presents results of the Gregory and Hansen (1996) tests for the US and Asian securitized real estate markets. Like in the case with the European markets, we find rather weak evidence in favour of cointegration. Namely, according to the test results, only in case of US and Australia we find that the null hypothesis of no cointegration is rejected, but only by the ADF test for level and for trend change models. The dates of the estimated breaks are March 1994 and December 1992, respectively. The absence of stable long-run relationships suggests that markets tend to deviate from long-run equilibrium and implies presence of diversification benefits among these markets. The latter result is in line with the findings by Yunus and Swanson (2007) and Schindler (2009b).

Table 19: Gregory and Hansen (1996) Test Results: North-American Bivariate Relationships

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
Canada – US	-4.74**	-4.71**	-30.49	-6.53***	-6.50***	-55.61**	-5.07**	-5.26**	-30.62
	10/2003	08/2007	08/2007	03/2000	01/2000	01/2000	05/2005	08/2007	
US – Canada	-4.03	-4.50*	-30.50	-6.47***	-6.42***	-55.59**	-3.31	-5.07*	-33.55
		08/2007	08/2007	08/2000	06/2000	06/2000	08/2007	08/2007	

Notes: The numbers in the first row indicate values of the corresponding test statistic. The numbers in the second row indicate the estimated date of the structural break in case of significant test statistic. ***, **, * indicate significance at the 1%-, 5%- and 10%-level, respectively.

Table 20: Gregory and Hansen (1996) Test Results: Bivariate Relationships between US and European Markets

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
US – Belgium	-3.92	-2.65	-13.52	-3.85	-1.80	-10.11	-4.21	-3.70	-25.02
US – France	-4.14	-2.39	-14.07	-3.81	-1.87	-10.87	-3.99	-2.78	-18.81
US – Germany	-3.51	-2.78	-15.03	-3.63	-2.34	-12.64	-3.81	-2.86	-15.84
US – Italy	-4.49*	-3.39	-22.46	-3.65	-2.69	-16.88	-4.59	-4.61	-40.19
	03/1994								
US – Netherlands	-3.98	-2.75	-15.21	-3.70	-1.90	-11.00	-3.44	-3.52	-25.96
US – Switzerland	-4.19	-2.15	-11.74	-3.52	-1.56	-8.16	-4.32	-2.63	-14.02
US – UK	-5.02	-3.22	-19.20	-4.76*	-3.52	-24.95	-4.83*	-4.14	-35.53
				03/1994			07/1998		
US – Sweden	-4.37*	-3.51	-23.65	-3.51	-2.21	-13.37	-4.01	-3.21	-23.48
	12/1992								

Notes: The numbers in the first row indicate values of the corresponding test statistic. The numbers in the second row indicate the estimated date of the structural break in case of significant test statistic. ***, **, * indicate significance at the 1%-, 5%- and 10%-level, respectively.

Table 21: Gregory and Hansen (1996) Test Results: Bivariate Relationships between US and Asian Markets

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
US – Australia	-4.95** 03/1994	-3.60	-25.44	-4.66	-4.15	-31.86	-5.48** 12/1992	-3.63	-24.77
US – Hong Kong	-3.21	-2.59	-17.61	-2.85	-1.85	-11.28	-3.22	-2.84	-19.52
US – Japan	-3.37	-3.31	-19.16	-2.97	-2.15	-12.97	-3.09	-2.86	-17.34
US – Singapore	-3.78	-2.78	-16.71	-3.47	-1.57	-3.40	-3.40	-2.72	-15.83

Notes: The numbers in the first row indicate values of the corresponding test statistic. The numbers in the second row indicate the estimated date of the structural break in case of significant test statistic. ***, **, * indicate significance at the 1%-, 5%- and 10%-level, respectively.

European markets

When we consider bivariate relations between the European markets, we find a number of relations (Table 22). We find that Dutch, Swiss, and Swedish real estate indices are influenced by the French one, suggesting that France may be a regional “core” market. Furthermore, the Dutch index seems to co-move with Swiss, Belgian, and Swedish indices. There is also weak evidence in favour of the cointegration relation between the Netherlands and France with France being independent variable, in case of regime change specification, albeit only at the 10% level of significance. Our results confirm earlier findings of Schindler (2010) who points out that France and the Netherlands can be considered as the two “core” markets of the Western Europe. Interestingly, Sweden appears to be affected by at least three European markets: French, Dutch and Swiss, and accordingly to Engle and Granger (1987) test, also by the Italian and German indices. It is therefore probably not an optimal choice for a diversified real estate portfolio. Interestingly, accordingly to Gregory and Hansen (1996) test, German, UK and Italian securitized real estate markets remain isolated in the long run. The literature increasingly finds that the EMU credit and stock markets are becoming more integrated, at least in terms of co-movements (Rangvid (2001), Aggarwal et al. (2005), Hardouvelis et al. (2006)). This seems not to be the case for the EU securitized real estate markets. Therefore, these markets might be of interest for investors seeking diversification opportunities, at least when we consider markets individually rather than within a group. These results also suggest that the larger and smaller (in terms of market capitalization) European securitized real estate markets display different degree of long-run co-movements. Our findings echo those of Liow et al. (2005) who find weak evidence in favour of one cointegration relationship among Italy, Germany, and the UK over the period from 1993 to 2003.

Asia-Pacific markets

The four Asian markets display two sets of bivariate relations over the sample period (Table 23). One pair of the relations is between Japan and Hong Kong (in both directions) and another pair of relations is found between Japan and Singapore (also in both directions). The null of no cointegration for these two pairs of markets has been rejected by all unit root tests for all the three model specifications at the 5% or 1% level of significance.

We also find weaker evidence in favour of cointegration between Hong Kong and Singapore, when Hong Kong market index is the dependent variable. The ADF test rejects the null of no cointegration in case of level and trend shift models for this pair of markets. For the pair Japan and Australia we find only very weak evidence against the null of no cointegration: in case of trend specification at the 10% level of significance. From the above results, Japan emerges as the “core” market for this group. Australian market, on the other hand, seems to be isolated from the other markets in the group and may therefore be seen as offering diversification benefits to investors seeking to diversify risks due to holding Asia-Pacific real estate securities.

Gerlach et al. (2006) analyses the impact of the Asian financial crisis in 1997 on the integration of the Asia-Pacific real estate markets. They conclude that judging benefits from long-run diversification by Johansen cointegration methodology might be misleading. It is shown that the number of cointegration relations increases when the Asian financial crisis of 1997 is explicitly considered in the analysis by testing for breaks in the trend. However, Gerlach et al. (2006) impose the date of the structural break exogenously. When we estimate the date of structural break from the data, it is interesting that with the exception of the relationship between Japan and Hong Kong, none of the identified break dates falls on 1997-1998, the years of the Asian and subsequent Russian financial crises. This illustrates the point made earlier in this section that setting up the date of the structural break exogenously may not correctly reflect the nature of the underlying data process.

Table 22: Gregory and Hansen (1996) Test Results: European Bivariate Relationships

Countries	Level			Trend			Regime		
	ADF Test	**PP Z(t)**	**PP Z(a)**	**ADF Test**	**PP Z(t)**	**PP Z(a)**	**ADF Test**	**PP Z(t)**	**PP Z(a)**
Belgium – France	-4.42* 01/1998	-4.16	-31.21	-4.27	-4.32	-34.56	-4.47	-4.57	-39.20
Belgium – Germany	-3.65	-3.14	-15.06	-4.08	-3.72	-24.71	-4.02	-3.01	-14.66
Belgium – Italy	-2.96 103	-2.78 172	-13.07 172	-4.68 111	-4.40 119	-32.47 119	-3.97 111	-3.03 115	-17.40 115
Belgium – Netherlands	-4.73** 12/1998	-4.45* 01/2000	-36.69* 01/2000	-4.87* 11/2005	-4.53	-37.96	-4.78* 12/1998	-4.64	-40.01
Belgium – Switzerland	-3.39	-3.28	-22.28	-4.37	-4.12	-31.00	-3.93	-3.28	-21.59
Belgium – UK	-3.48	-2.81	-13.67	-4.29	-3.64	-25.50	-3.45	-2.93	-13.81
Belgium – Sweden	-3.26	-2.69	-14.27	-3.89	-3.48	-24.43	-5.06** 10/1998	-4.17	-32.28
France – Belgium	-3.88	-3.89	-28.24	-4.05	-4.00	-30.41	-3.78	-3.85	-30.73
France – Germany	-3.75	-3.63	-24.75	-3.84	-3.64	-23.18	-3.72	-3.77	-27.32
France – Italy	-3.63	-3.53	-24.15	-4.79* 09/1998	-3.78	-25.45	-3.64	-3.52	-24.14
France – Netherlands	-3.94	-4.07	-29.80	-4.57	-4.58	-38.56	-4.76* 10/1998	-4.07	-35.27
France – Switzerland	-3.89	-3.90	-27.55	-3.72	-3.73	-26.87	-4.51	-4.44	-37.44
France – UK	-3.34	-3.28	-21.71	-3.42	-3.25	-20.81	-3.34	-3.27	-21.37
France – Sweden	-3.76	-3.76	-25.59	-3.61	-3.50	-24.02	-4.04	-4.51	-39.62

Table 22 continues on the next page

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
Germany – Belgium	-4.09	-2.28	-9.14	-4.14	-2.58	-10.83	-4.14	-3.01	-16.66
Germany – France	-4.09	-2.18	-9.88	-4.16	-2.65	-11.14	-3.83	-2.61	-11.67
Germany – Italy	-3.64	-2.86	-13.53	-3.81	-3.41	-20.47	-3.82	-3.50	-21.13
Germany – Netherlands	-4.18	-2.34	-9.31	-4.25	-2.68	-11.23	-4.12	-2.67	-14.44
Germany – Switzerland	-3.98	-2.18	-10.19	-4.05	-2.18	-10.33	-4.15	-2.34	-10.52
Germany – UK	-4.05	-3.24	-15.29	-3.87	-3.20	-15.34	-4.37	-4.38	-35.42
Germany – Sweden	-3.92	-2.38	-9.90	-4.05	-2.34	-10.79	-4.34	-2.91	-17.15
Italy – Belgium	-3.10	-2.56	-13.63	-4.17	-4.07	-29.94	-3.26	-2.96	-17.53
Italy – France	-2.50	-2.30	-10.84	-4.95* 09/1998	-3.72	-26.38	-2.50	-2.52	-12.38
Italy – Germany	-2.94	-2.84	-13.01	-4.17	-3.44	-22.57	-2.85	-2.65	-12.29
Italy – Netherlands	-2.90	-2.28	-11.16	-4.55	-3.87	-27.8	-3.06	-2.60	-13.85
Italy – Switzerland	-3.25	-2.47	-13.60	-3.36	-3.12	-17.73	-3.11	-2.34	-12.41
Italy – UK	-2.65	-2.32	-11.23	-3.88	-3.50	-23.54	-3.23	-3.09	-16.91
Italy – Sweden	-3.46	-2.87	-14.68	-3.41	-2.82	-14.33	-3.07	-2.86	-16.51
Netherlands – Belgium	-4.78** 10/2004	-4.30	-34.45	-4.91* 05/2005	-4.28	-34.07	-5.28** 02/2004	-5.23** 02/2004	-48.79** 02/2004
Netherlands – France	-4.38* 08/1998	-4.18	-30.38	-4.49	-4.68	-39.69	-4.88* 05/1998	-4.84* 04/1998	-43.80* 04/1998

Table 22 continues on the next page

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
Netherlands – Germany	-3.55	-3.24	-16.14	-3.85	-3.54	-22.32	-3.59	-3.25	-18.00
Netherlands – Italy	-2.98	-2.89	-13.97	-4.99* 12/1998	-4.09	-27.82	-3.36	-3.13	-15.14
Netherlands – Switzerland	-3.76	-3.24	-21.85	-3.87	-3.58	-25.67	-3.70	-3.35	-22.53
Netherlands – UK	-2.81	-2.81	-12.31	-3.26	-3.40	-22.42	-2.85	-2.91	-13.01
Netherlands – Sweden	-2.87	-2.78	-16.23	-3.26	-3.19	-20.62	-5.23** 08/1999	-5.33** 10/1999	-50.76** 10/1999
Switzerland – Belgium	-3.51	-3.28	-23.22	-4.42	-4.19	-30.96	-3.43	-3.13	-21.45
Switzerland – France	-4.21	-4.22	-32.12	-4.23	-4.24	-31.63	-4.77* 12/1994	-4.94** 09/1994	-44.71* 09/1994
Switzerland – Germany	-2.66	-2.71	-14.39	-3.51	-3.48	-19.32	-2.78	-2.83	-15.54
Switzerland – Italy	-2.91	-2.45	-13.77	-4.46	-4.04	-24.15	-2.83	-2.50	-13.99
Switzerland – Netherlands	-3.72	-3.50	-25.99	-4.13	-4.08	-30.88	-3.73	-3.50	-25.95
Switzerland – UK	-2.31	-1.95	-8.95	-3.44	-3.36	-21.71	-2.87	-2.94	-16.61
Switzerland – Sweden	-3.42	-4.19	-39.56* 09/1992	-4.14	-5.54*** 04/2000	-55.35** 04/2000	-3.82	-3.65	-27.31
UK – Belgium	-3.63	-2.45	-14.36	-3.64	-2.27	-13.23	-4.02	-3.05	-18.41
UK – France	-2.68	-1.84	-9.52	-2.70	-2.40	-15.15	-3.26	-3.62	-24.96
UK – Germany	-3.28	-2.86	-12.46	-3.23	-2.81	-13.08	-4.34	-4.35	-35.78
UK – Italy	-3.00	-2.60	-14.26	-3.33	-3.01	-17.76	-4.27	-4.02	-29.75

Table 22 continues on the next page

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
UK – Netherlands	-2.63	-1.69	-9.48	-2.94	-2.09	-11.60	-2.64	-2.39	-10.73
UK – Switzerland	-3.17	-1.10	-5.01	-3.14	-1.70	-9.04	-3.29	-2.82	-18.01
UK – Sweden	-3.00	-2.57	-16.85	-3.01	-2.95	-20.43	-3.05	-3.09	-18.76
Sweden – Belgium	-4.01	-3.11	-13.43	-4.64	-3.77	-24.41	-3.86	-3.74	-26.56
Sweden – France	-4.83^{**} 10/1993	-4.14	-27.69	-4.03	-4.00	-25.91	-4.63	-6.52^{***} 03/1993	-72.21^{***} 04/1993
Sweden – Germany	-4.05	-3.57	-11.72	-4.43	-3.70	-21.12	-4.41	-3.23	-11.88
Sweden – Italy	-3.72	-3.83	-14.51	-4.05	-3.54	-18.63	-4.59	-4.58	-29.31
Sweden – Netherlands	-3.37	-3.14	-16.44	-4.27	-4.06	-28.02	-4.24	-4.96^{**} 10/1999	-44.67^{*} 10/1999
Sweden – Switzerland	-4.69^{**} 10/1993	-4.42^{*} 09/1992	-35.41	-4.22	-5.12** 04/2000	-48.20^{**} 04/2000	-4.56	-5.24^{**} 11/1993	-50.50^{**} 11/1993
Sweden – UK	-3.79	-3.39	-16.64	-4.41	-4.46	-35.57	-4.01	-3.28	-22.94

Notes: The numbers in the first row indicate values of the corresponding test statistic. The numbers in the second row indicate the estimated date of the structural break in case of significant test statistic. ***, **, * indicate significance at the 1%-, 5%- and 10%-level, respectively.

Table 23: Gregory and Hansen (1996) Test Results: Asia-Pacific Bivariate Relationships

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
Australia – Hong Kong	-2.56	-1.07	-4.80	3.25	-0.74	-3.47	-2.13	-1.13	-5.84
Australia – Japan	-2.92	-2.58	-15.53	-3.23	-1.91	-11.17	-2.88	-2.26	-10.83
Australia – Singapore	-3.62	-1.86	-10.14	-3.63	-0.92	-4.43	-2.62	-2.06	-12.16
Hong Kong – Australia	-3.66	-3.39	-21.62	-4.20	-4.35	-35.48	-3.75	-3.50	-22.19
Hong Kong – Japan	-5.58^{***}	-5.16^{***}	-40.55	-5.76^{***}	-5.28^{**}	-44.14^{*}	-5.91^{***}	-5.49^{***}	-57.38^{***}
	01/1996	01/1996		09/2001	08/2002	08/2002	05/1993	05/1993	05/1993
Hong Kong – Singapore	-4.62^{**}	-3.93	-27.99	-5.39**	-4.61	-37.67	-4.64	-4.01	-28.62
	08/1997			10/1996					
Japan – Australia	-4.12	-4.13	-25.01	-4.91*	-4.89*	-39.51	-4.45	-4.51	-33.23
Japan – Hong Kong	-5.30^{***}	-5.13^{**}	-35.99	-5.77^{***}	-5.57^{***}	-46.14^{*}	-6.33^{***}	-5.93^{***}	-57.58^{***}
	02/1996	10/1995		06/2004	02/2004	02/2004	05/1998	05/1998	05/1998
Japan – Singapore	-6.11^{***}	-6.18***	-53.03^{***}	-6.23^{***}	-6.34^{***}	-59.27^{***}	-6.48^{***}	-6.53^{***}	-56.75^{**}
	12/2003	09/2003	09/2003	08/2003	11/2003	11/1993	09/2003	09/2003	09/2003
Singapore – Australia	-3.10	-2.61	-15.80	-3.46	-3.60	-25.44	-3.33	-3.34	-23.99
Singapore – Hong Kong	-4.26	-3.70	-25.41	-4.75^{*}	-4.11	-30.77	-4.19	-3.66	-26.18
Singapore – Japan	-5.12^{**}	-5.22^{***}	-41.09^{**}	-5.78^{***}	-5.87^{***}	-54.02^{**}	-5.23^{**}	-5.34^{**}	-42.35^{*}
	09/2003	08/2003	08/2003	08/1993	11/1993	11/1993	08/2003	08/2003	08/2003

Notes: The numbers in the first row indicate values of the corresponding test statistic. The numbers in the second row indicate the estimated date of the structural break in case of significant test statistic. ***, **, * indicate significance at the 1%-, 5%- and 10%-level, respectively.

Regional indices

Finally, we test for bivariate cointegration across the regions using three regional indices: Asia-Pacific, European, and North American regional indices. We find evidence of the null being rejected in case of the following pairs of indices (Table 24). For the pair Developed Europe – North America (with the European index being the dependent variable), we find that the null is rejected by two unit root tests: one rejection happens for the level specification at the 1% level and another happens for the trend specification at the 10% level. For the pairs North America and Developed Europe (with the latter being the independent variable) and well as Asia-Pacific and North America (with the latter being the independent variable) the null is rejected only by one unit root test at 10% level of significance. Since the null hypothesis is rejected at the low significance level by only one of the three unit root tests and only for a single model specification in each of the cases, we conclude that there is no sufficient evidence in favour of cointegration relations for these two pairs of markets in these specific groupings. This result is in line with the earlier findings for this group using conventional and recursive Johansen cointegration tests. This finding is in accordance with that of Gallo and Zhang (2009).

Summing up the results described above, we would like to note the following issues. The extent of international integration between the securitized real estate markets within the Asia-Pacific, European and North-American regions seem to differ. The two North-American markets, the US and Canada, seem to be fully integrated, at least according to the results of the Gregory and Hansen (1996) test. US investors will benefit from diversifying into European and Asia-Pacific markets, except of probably UK and Australia. However, we do not find evidence of substantial integration for the other two regions. We find nine out of fifty six possible cointegration relations for the group of eight European markets. This means that only 16% of all possible cointegration relations exist for these markets. We find more bivariate relations between larger and smaller real estate markets than among the larger real estate markets. According to our results, Dutch and French markets seem to be “core” real estate markets in the regions, to which smaller real estate markets in the region seem to be adjusting in the long-run. On the other hand, German, Italian, and the UK markets seem to be isolated. This means that investing in these markets could potentially offer diversification benefits seeking to diversify their holding of the European real estate securities. In case of the Asia-Pacific markets, we find five bivariate relations at 5% level of significance out of possible twelve, which means that about 40% of the possible relations actually exist in our sample.

Table 24: Gregory and Hansen (1996) Test Results: International Bivariate Relationships, Regional Indices

Countries	Level			Trend			Regime		
	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)	ADF Test	PP Z(t)	PP Z(a)
North America – Asia Pacific	-3.73	-3.36	-24.83	-3.70	-2.75	-19.26	-3.90	-3.58	-26.98
North America – Developed Europe	-4.50*	-3.34	-21.59	-4.65	-2.87	-18.73	-3.89	-4.07	-31.00
Developed Europe – North America	-5.18*** 02/1994	-2.98	-16.98	-4.35	-3.01	-16.76	-4.71* 12/2005	-3.99	-25.43
Developed Europe – Asia Pacific	-3.52	-2.42	-11.85	-3.24	-2.55	-13.48	-3.53	-3.73	-25.43
Asia Pacific – Developed Europe	-4.07	-3.06	-18.83	-4.30	-3.91	-29.09	-3.96	-3.31	-21.18
Asia Pacific – North America	-4.03	-3.33	-24.23	-4.85*	-4.18	-32.73	-3.89	-3.40	-23.49

Notes: The numbers in the first row indicate values of the corresponding test statistic. The numbers in the second row indicate the estimated date of the structural break in case of significant test statistic. ***, **, * indicate significance at the 1%-, 5%- and 10%-level, respectively.

This shows higher extent of integration among these markets in comparison with the European markets, at least when the extent of integration is measured by the total number of existing equilibrium relations. In this group, Australia emerges as an isolated market, while other markets seem to be linked, albeit by relations of differing significance

5.3.2 Recursive Hansen and Johansen (1999) Test

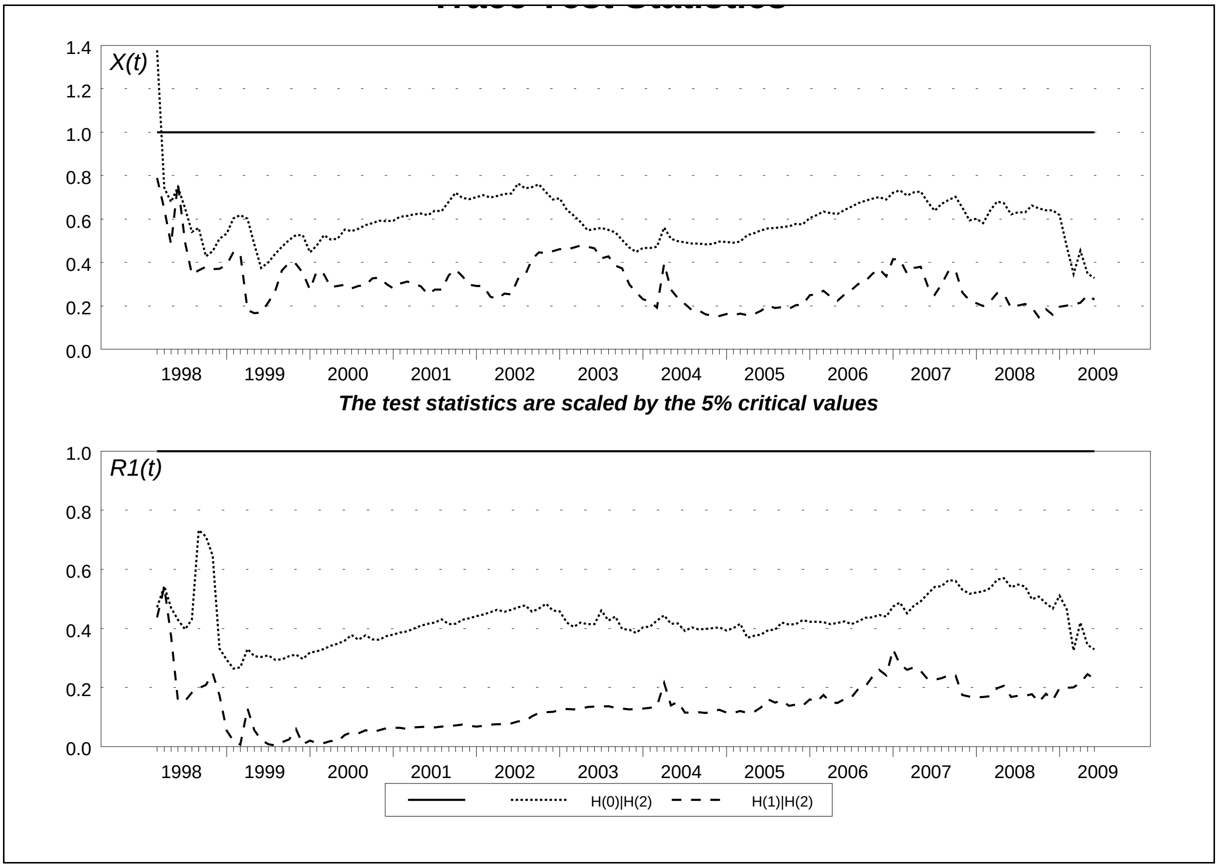
To explore the dynamics of the equilibrium relationships in more detail, we turn to the multivariate recursive cointegration methodology by Hansen and Johansen (1999). Figure 4- to Figure 14 present the rescaled recursive Hansen and Johansen (1999) trace statistic as described in Section 3 for the groups of EPRA/NAREIT indices described in Section 4. The results are presented for the whole sample period.[10]

North-American markets

We find no cointegration relations between the US and Canadian indices with the trace statistic being always below one (Figure 4). This result is in line with the findings of the Johansen cointegration test presented earlier. The exception is a very short period in 1998 in case of *X*-form specification. However, a sharp increase in the recursive test statistic at the start of the estimations is not unusual for this statistic and should rather be attributed to the particularities of the statistic calculations, rather than to economic forces (Juselius (2007)). This result is contrary to that of Gallo and Zhang (2009) who find a cointegration relation between these two markets over the period from 1992 to 2007. However, it should be pointed out that Gallo and Zhang (2009) study a shorter period, use different indices and methodology (conventional Johansen and Juselius cointegration test) to arrive at this result.

[10] The estimations were performed using RATS7.2 software. The length of the initial sample period t_0 (see Section 3 for more information) was set in case of each particular group of countries and takes into account the number of variables in the system and the number of lags in the VAR specification.

Figure 4: Recursive Trace Test: North America (US and Canada)



US and European markets

The US real estate securities and securities of the core European markets exhibit evidence of one cointegration relation which seems to be unstable and ceases altogether in the second half of 2003 (Figure 5). There is no evidence that the relationships have been strengthening.

The relationship among the US and periphery European markets, to the contrary seem to remain stable since 2002 in case of *R*-form, and according to X-form, there was one stable long-run relation for the whole sample period. There is also evidence in favour of the second equilibrium relation. However, the latter ceased between the second half of 1998 and early 2002. It ceased again in late 2003 and it has not re-emerged since then (Figure 6). Since only one equilibrium relationship among the five markets has been found and the second relationship has been rather unstable, we conclude that there is no evidence of intensifying co-movements for this group of markets.

When all European markets are included, there is an evidence of one relationship for the whole sample. A second equilibrium relationship is present except for the period between the late 1997 and early 2002. Since about January 2002 the second relationship seem to have re-emerged and persisted despite the drop in the test statistic in the wake of the recent financial crisis (Figure 7). Gallo and Zhang (2009) also find limited evidence of cointegration between the US and the EU markets in that they report a single relationship between the US and seven EU markets within a bivariate set-up.

Figure 5: Recursive Trace Test: US and European Core Markets

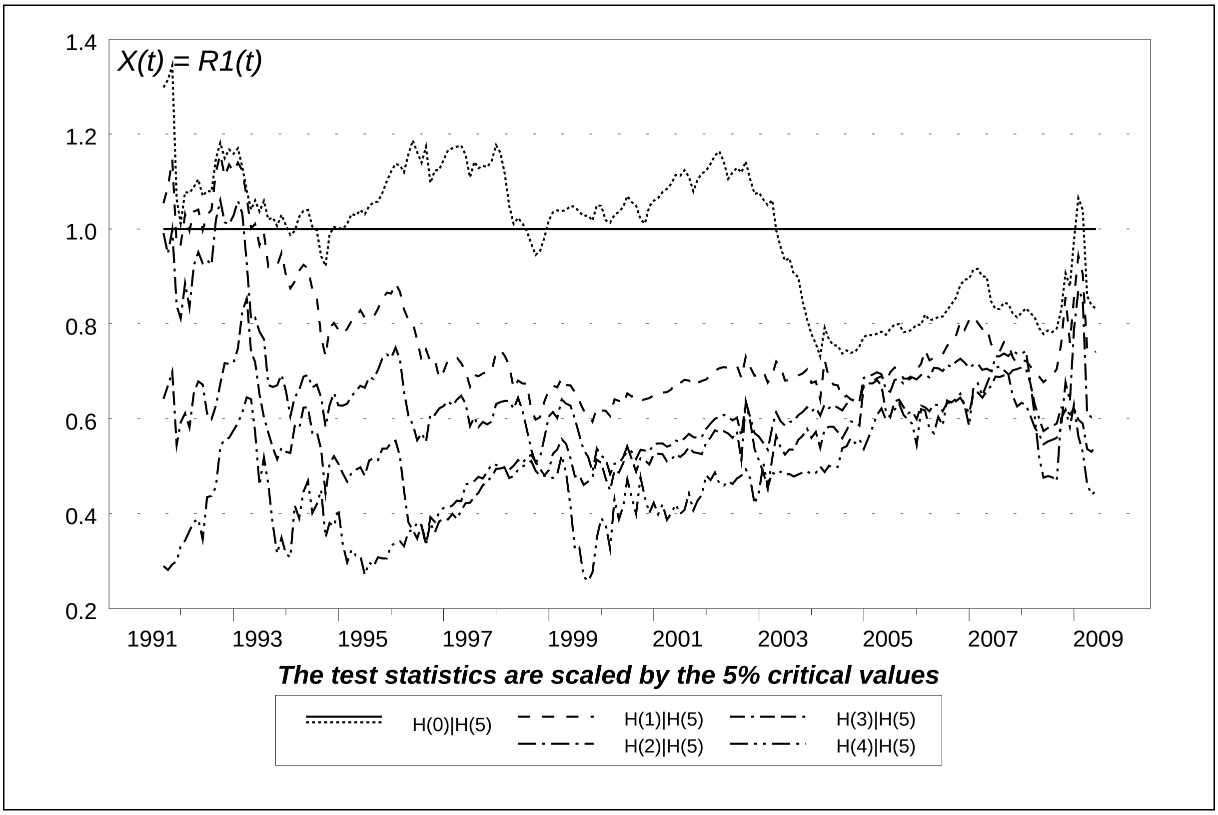


Figure 6: Recursive Trace Test: US and European Periphery Markets

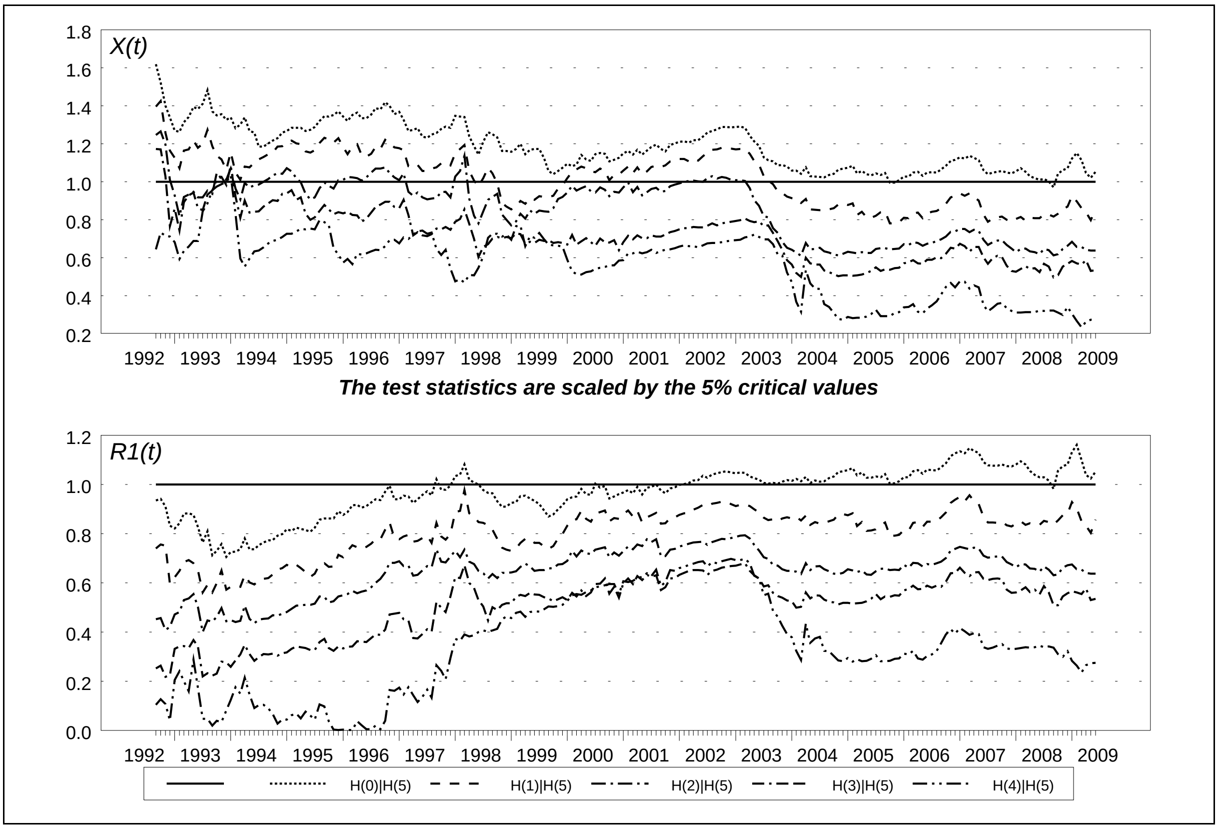
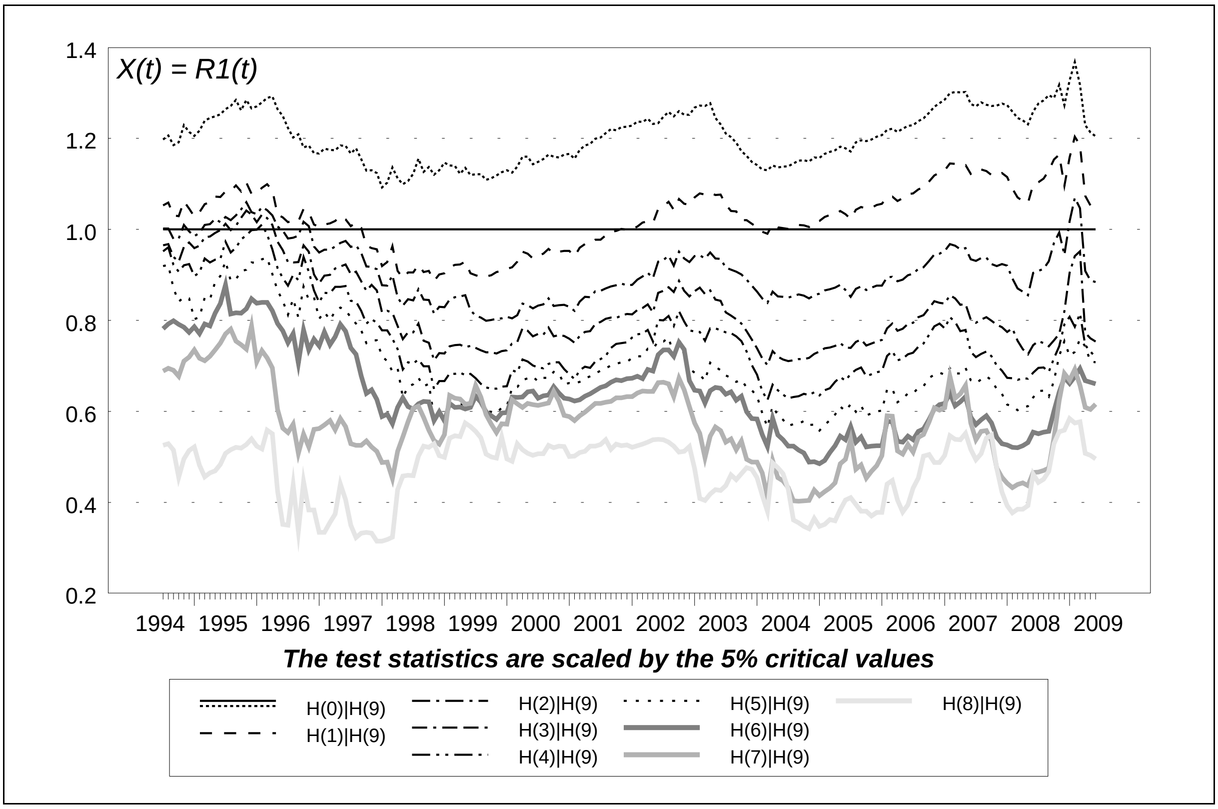


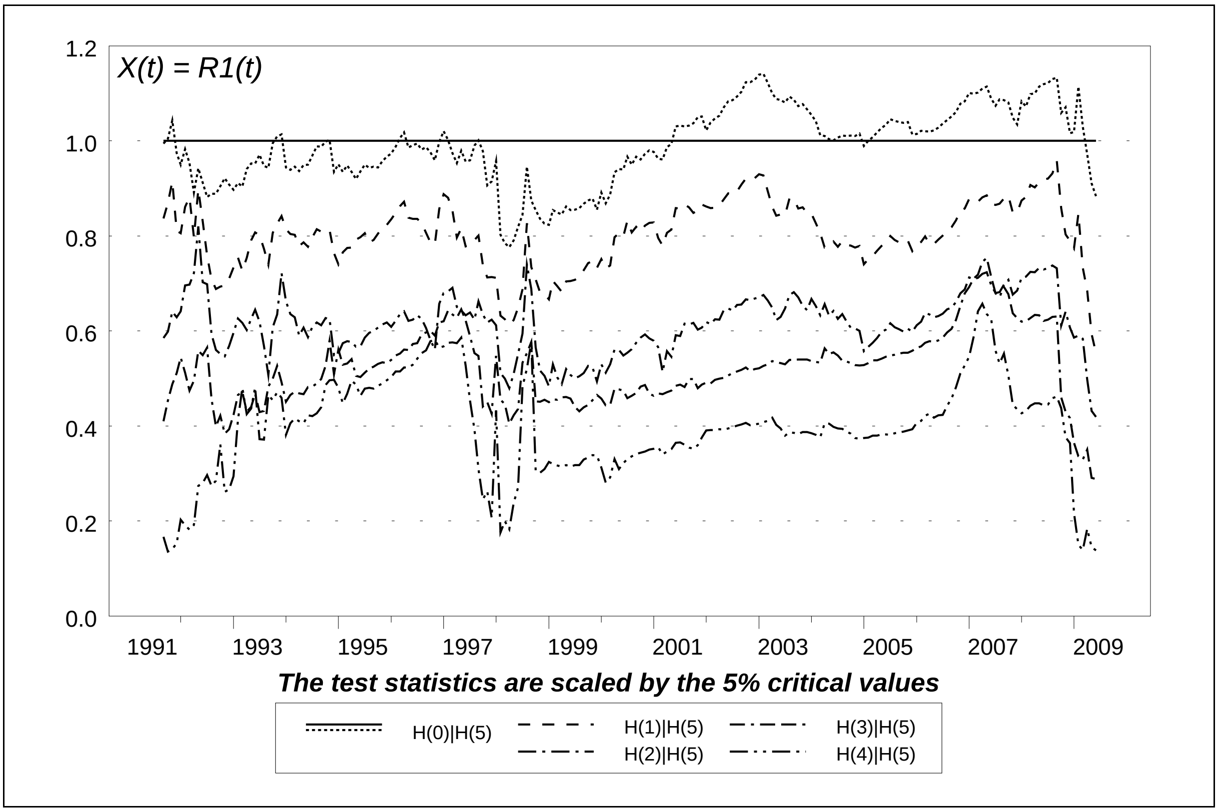
Figure 7: Recursive Trace Test: US and All European Markets



US and Asia-Pacific markets

As Figure 8 demonstrates, there is no stable relationship in the group consisting of the US and the four Asia-Pacific markets. The normalized trace statistic fluctuates around the value of 1 until 1997 and drops below one during the period of 1997-2001. The statistic reaches one from above again in the second half of 2004 and drops below one in the second half of 2009. It should be pointed out that the periods of no cointegration (1997-2001 and middle of 2009) seem to have coincided, at least to some extent, with the major financial crises. We can suggest that the long-run relationships in Johansen sense may not be immune to the periods of financial instability. Whereas it is well known that short-term correlations tend to increase during volatile periods (Longin and Solnik (2001) and Goetzmann et al. (2001)), this does not seem to be the case in long-term relationships. This result is in line with findings by Lucey and Voronkova (2008) in case of long-run relationships shared by the Russian stock market in the wake of the Russian financial crisis. The absence of stable long-run relationships suggests that markets tend to deviate from long-run equilibrium and implies presence of diversification benefits among these markets. The latter result is in line with the findings by Yunus and Swanson (2007) and Schindler (2009b).

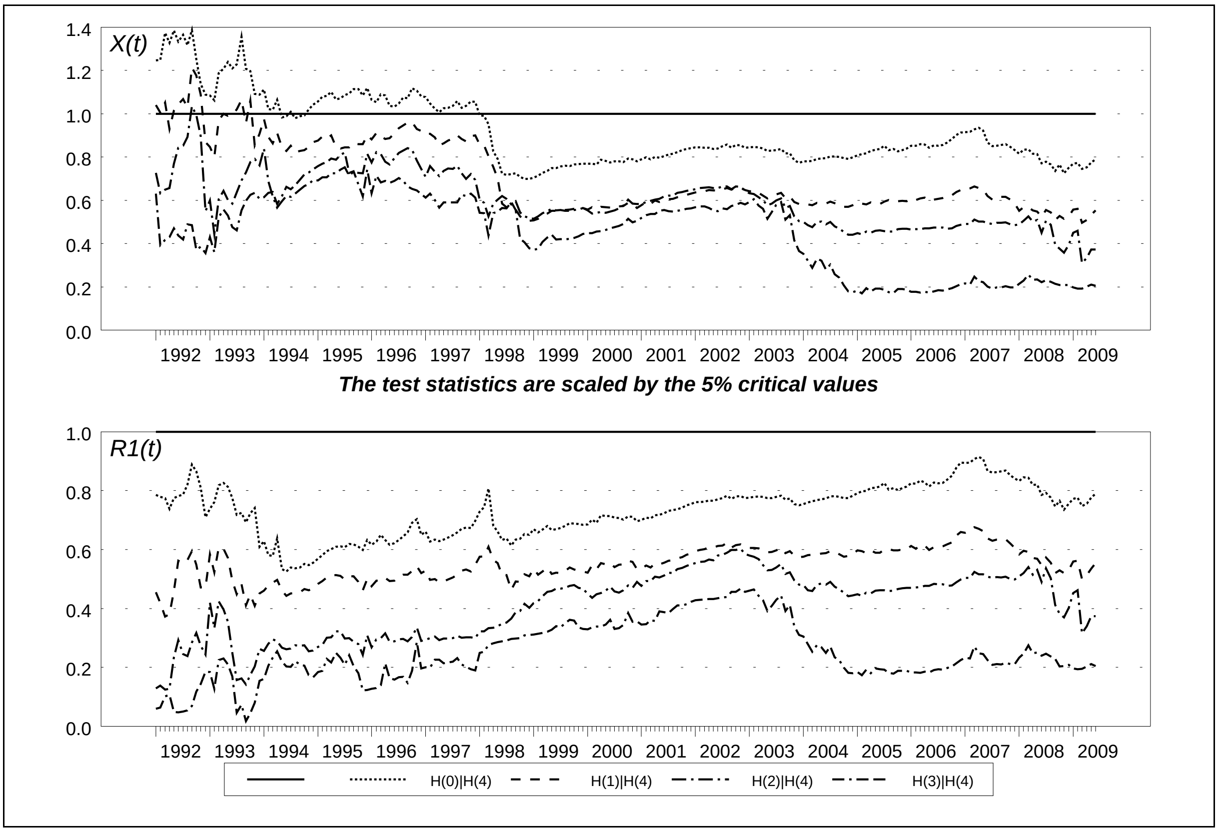
Figure 8: Recursive Trace Test: US and Asia-Pacific Markets



European core markets

There is an extensive literature on increasing integration among the Western European stock markets, yet a number of studies (see Section 5.2.2.) find that European real estate markets have not been converging as fast as European equity markets. The reasons behind this phenomenon may be the pronounced domestic focus of the real estate companies' portfolios and small size of real estate markets. Our longer sample may help to identify whether the integration processes have intensified over the last eight years. We find that core European real estate securities markets (Germany, France, the Netherlands and the UK) show no stable equilibrium relationships (Figure 9). A single relationship in this group of markets ceased in early 1998, accordingly to the results based on *X*-form. According to the results based on *R*-form, there are no relationships in this group. Liow et al. (2005) find weak evidence in favour of one cointegration relationship among France, Italy, Germany, and the UK over the period from 1993 to 2003.

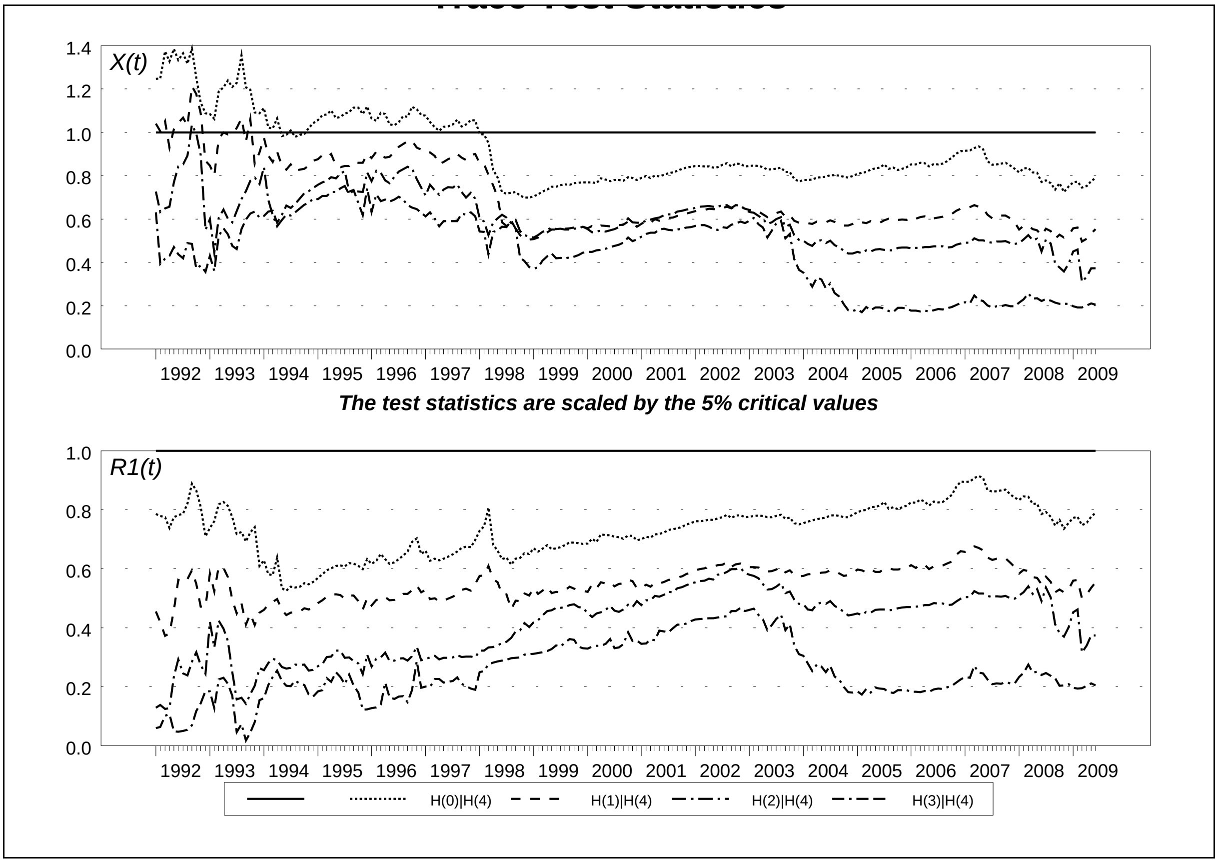
Figure 9: Recursive Trace Test: Core European Markets



European periphery markets

As shown by Figure 10, very similar results hold for the group including the European markets periphery (Belgium, Italy, Sweden, and Switzerland). These findings suggest presence of diversification benefits in the two groups of European markets.

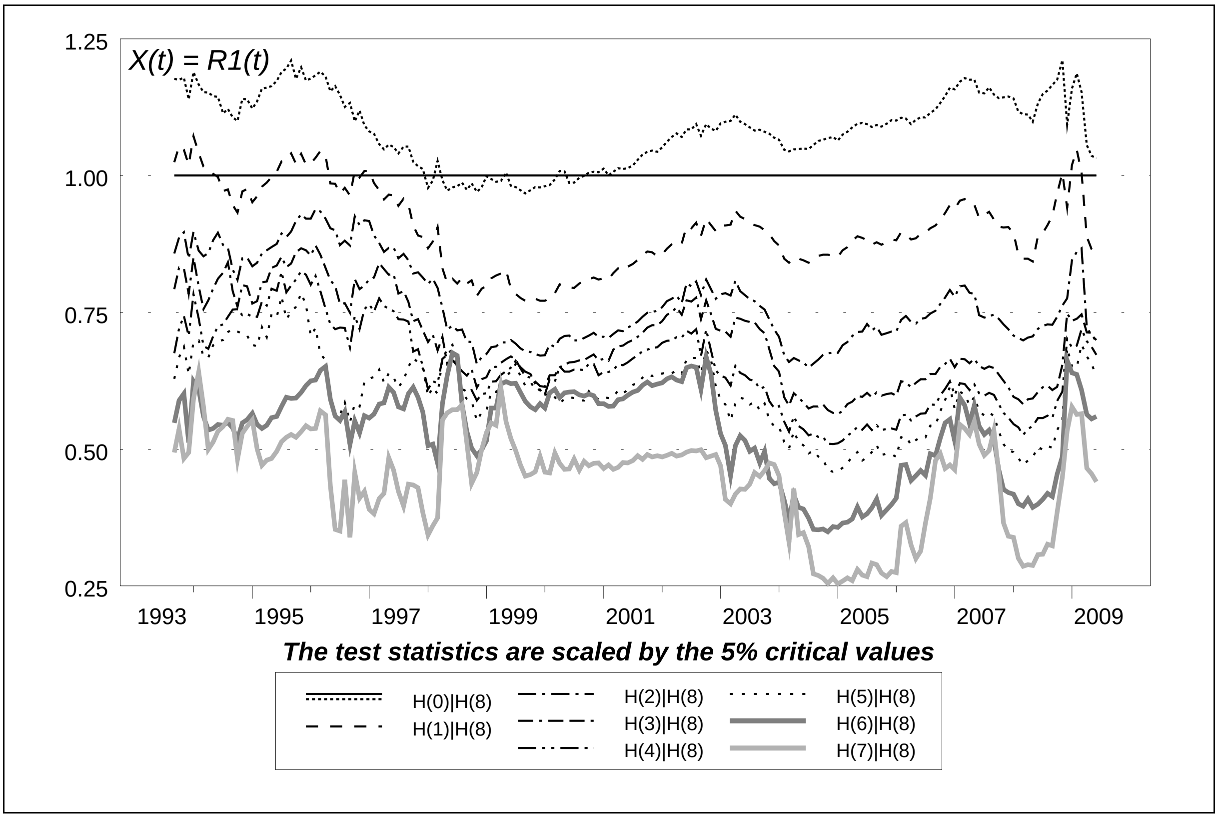
Figure 10: Recursive Trace Test: European Periphery Markets



All European markets

When all European markets are considered as a group, we find evidence in favour of a single cointegration relation (Figure 11). However, the relationship is not stable as it ceased between the second half of 1998 and the early 2001. We also observe a decline in the value of the normalized trace statistic towards one starting from the late 2008. It seems that the Asian and Russian financial crises and the sub-prime mortgage crises disrupted long-run equilibrium relationships among the European real estate securities markets. These results suggest that the real estate securities seem to be prone to the fluctuations in the securities markets. This observation may be useful for when studying pricing of real estate securities.

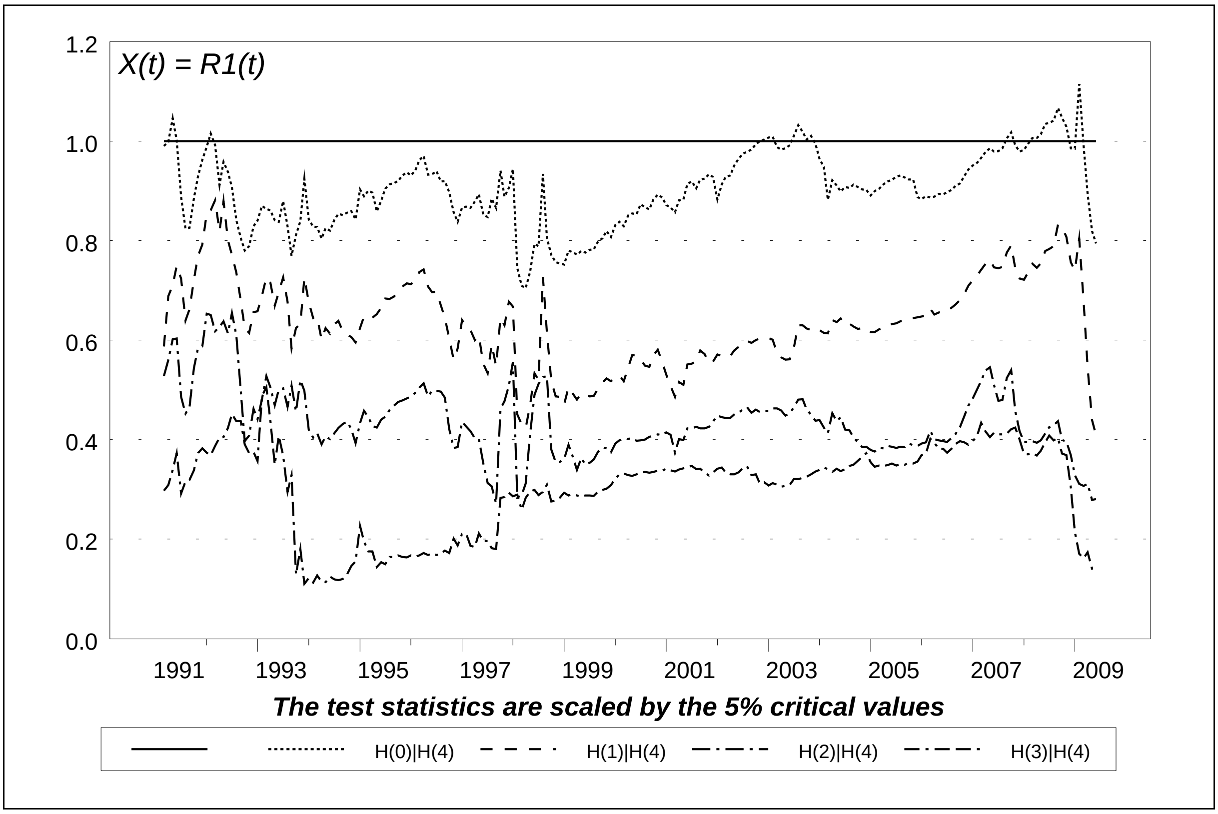
Figure 11: Recursive Trace Test: All eight European markets



Asia-Pacific markets

The four Asia-Pacific markets display no cointegration relation over the sample period (Figure 12). Although we observe an increase in the value of the normalized trace statistic since 1999, and despite trace statistic was above one for very short periods of time in 2002, 2003, and in 2008 and early 2009, it is notable that the statistic drops below one in early 2009. We suppose that the latter decline was driven by the recent financial crisis. The lack of cointegration among the Asia-Pacific markets is in line with the results in Garvey et al. (2001). Garvey et al. (2001) studied the Asia-Pacific markets over the period from 1975 through 2001. It seems that over the eight years following the end of their sample, the integration process in these markets has not accelerated.

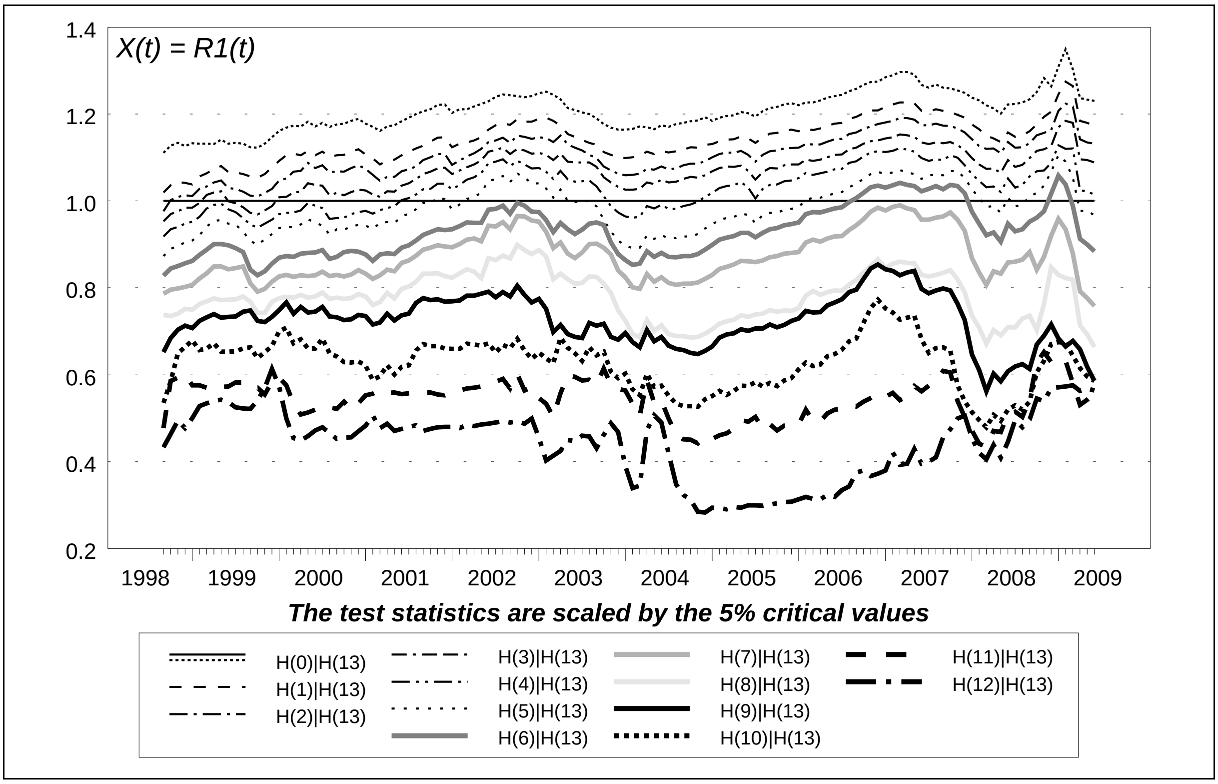
Figure 12: Recursive Trace Test: Asia-Pacific Markets



All markets

Finally, we consider all markets in one group, taking the point of view of an investor interested in broad international diversification. For the purposes of the estimations, we exclude Canada from the group since the data for the Canadian index is only available from December 1996. We find that in the group of all markets excluding Canada, the diversification benefits are unlikely to be expected, as we find between two and seven equilibrium relationships in this group at different points of the sample period. Two equilibrium relationships exist throughout the sample. The number of the long-run relationships has been steadily increasing for this group and it reached seven and the start of 2009. As in some cases described above, we observe a decline in the test statistic close to the end of the sample period. In the case of this group, the number and the values of the normalized trace statistic experienced a drop starting from the early 2009. However, notwithstanding the drop, the number of the equilibrium relations was equal to five at the end of the sample period. (Figure 13).[11] Yunus (2009) studied a much smaller set of international markets including Asia-Pacific, European markets and the US market. She finds one cointegration relation in her sample, with the second relation emerging after year 2003. Based on this finding, Yunus (2009) concluded that real estate securities markets move towards more integration. However, her sample period does not include the recent financial crisis period. When the recent financial crisis period is included, we observe a decline in the values of the corresponding test statistic and in the number of the cointegration relations. The reversal in the trend of the test statistic makes it difficult to draw conclusions about the future movements and about the future of the integration process among the real estate securities markets.

Figure 13: Recursive Trace Test: 13 markets

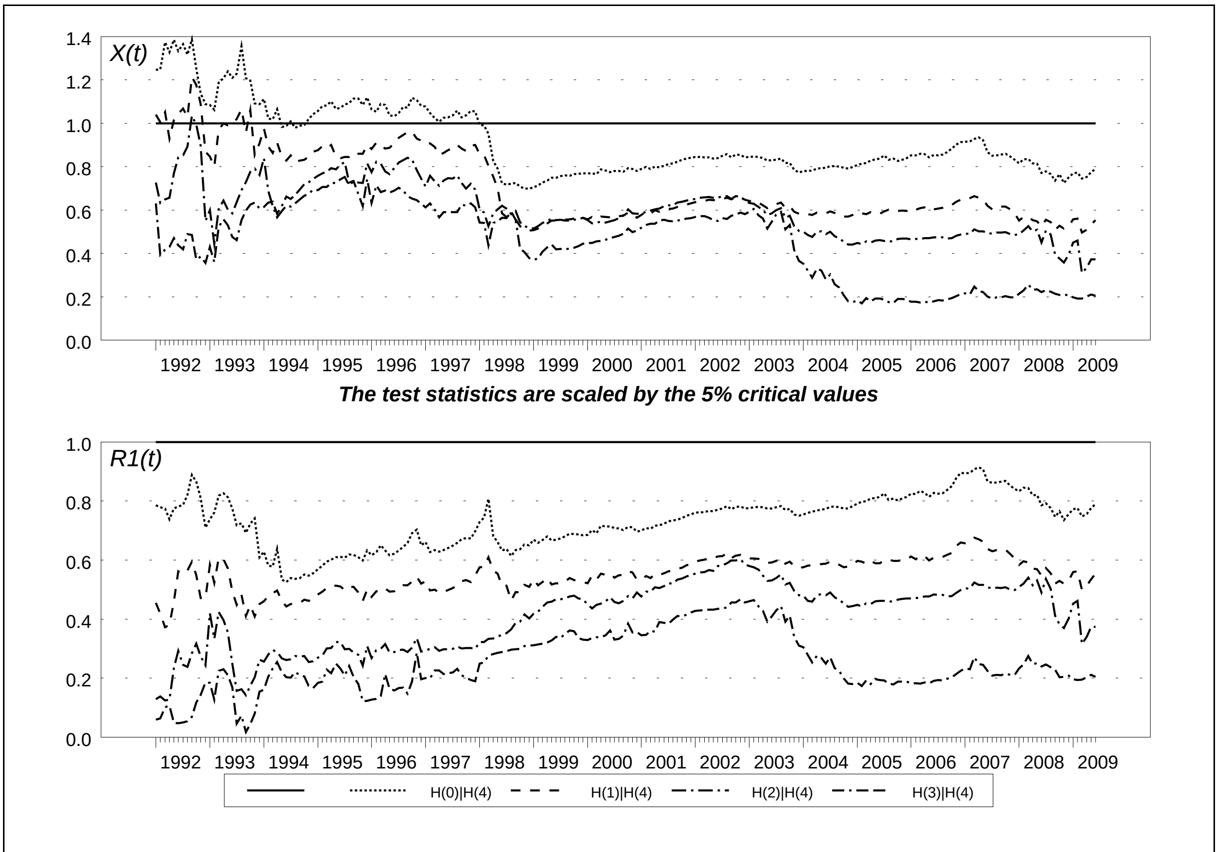


[11] It should be pointed out, however, that when we exclude Australia, Hong Kong, and Singapore, the number of cointegration relations drops to one (for the whole sample period) or two (for the sub-samples before November 1997 and after November 2004). The results are not reported here for the sake of space, but are available from the authors upon request. This drop in the number of cointegration relations may be explained by the interrelatedness of the three markets due to cross-border investments in the Asia-Pacific region. We thank Shaun Bond for pointing out this issue.

Regional indices

Previous studies (Gallo and Zhang (2009)) have considered inter-regional integration using the region-wide real estate securities indices. For the purposes of comparison with the earlier research, we conduct the tests for the set of the three regional indices: Asia-Pacific, Europe and North America. When the three indices are considered in a group, we find no evidence of cointegration relations as of early 1998 according to *X*-form and no evidence for the whole sample according to *R*-form (Figure 14). This finding is in accordance with that of Gallo and Zhang (2009). However, it should be born in mind that when we test for cointegration using national indices as in the preceding sub-section, we find stronger evidence of cointegration. The conclusions with regard to the extent of diversification benefits thus depend on the benchmarks used.

Figure 14: Recursive Trace Test: Asia-Pacific, Europe, Northern America



Summing up the results described above, we should point out the following issues. The extent of international integration and therefore the extent of diversification benefits may be sensitive to the testing methodology, to the sample period and selected benchmarks representing national and regional markets. In the case of the fourteen international real estate securities markets, we find that, on balance, within the multivariate time-varying framework, there is no conclusive evidence of increased integration within the three considered regions. Applying time-varying test does not yield evidence of increased international integration in the real estate securities markets. The results for the intra-continental integration depend on the used benchmarks. When regional, as opposed to national indices are used, we find even weaker evidence in favour of integration among the international real estate securities markets. Based on this, we conclude that the real estate securities markets have not been following the path of increased integration as it is the case for the international non-real-estate equity and bonds. Therefore, the considered markets may still offer diversification benefits to international investors. However, the extent of these benefits may or may not be economically significant depends on transaction costs, information asymmetries and other market frictions.

5.3.3 McCabe et al. (2003) Stochastic Cointegration Test

Harris, McCabe, and Leybourne (2002), suggest considering cointegration in a sense wider than that of Engle and Granger (1987) by loosening the strict requirement of stationarity of first differences of the series and requiring only the absence of stochastic I(1) trends. They allow for the presence of a non-linear heteroscedasticity that results in a volatile behaviour of the first differences of the time series. McCabe et al. (2003) develop stochastic cointegration test. They suggest testing the null hypothesis of stochastic cointegration against the alternative of no cointegration. Should the null not be rejected, the null hypothesis of stationary cointegration is tested against the alternative of heteroscedastic cointegration.[12]

North-American markets

As the Table 25 shows, the null hypothesis of stochastic cointegration is rejected by the stochastic cointegration test for the US and Canadian real estate markets in both directions, albeit at 10% level of significance only. Furthermore, the null hypothesis of stationary cointegration is rejected at 5% level of significance in both directions, suggesting presence of heteroscedastic cointegration relationship between the two North American markets. It is worth pointing out that the results of Engle and Granger (1987) test do not support existence of the equilibrium relationship between these two markets, but the Gregory and Hansen (1996) cointegration test does. Finding cointegration relation between the US and Canada is in line with the result of Gallo and Zhang (2009) who find a cointegration relation between these two markets over the period from 1992 to 2007 using Johansen and Juselius (1990) cointegration test.

[12] GAUSS code for calculation of the test statistics was kindly provided by Brendan McCabe.

Table 25: Stochastic Cointegration Test Results: North American Bivariate Relationships

Countries	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}
Canada – US	1.569	1.473*	1.528*
US – Canada	0.623	1.455*	1.657**

Notes: $\hat{\beta}_{AIV}$ denotes the slope coefficient from the cointegration equation. S_{nc} denotes the test statistic for the null hypothesis of stochastic cointegration against the alternative of no cointegration. S_{hc} denotes the test statistic for the null hypothesis of stationary cointegration against the alternative of heteroscedastic cointegration. ***, **, * indicate significance at the 1%-, 5%-, and 10%-level respectively. The relationships, for which the null of stochastic cointegration is not rejected, are shaded in grey. Thus shaded cell indicates the presence of a cointegration relationship between the given two markets. The country indicated in a row is the dependent variable. The country indicated in a column is the independent variable.

US and European markets

We find evidence in favour of two bivariate cointegration relations between the US and European markets (UK and Switzerland), and Italy, albeit for the latter the null of stochastic cointegration is rejected at the 10% level (Table 26). All of these three relations appear to be heteroscedastic cointegration relations. Notably, the relation between the US and the UK securitized real estate markets is also found by the Gregory and Hansen (1996) test. The results of stochastic cointegration in general suggest limited integration among the US and the European markets.

Table 26: Stochastic Cointegration Test Results: European Bivariate Relationships among US and European markets

Countries	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}
US – Belgium	0.524	2.196**	2.079**
US – France	0.458	1.778**	2.182**
US – Germany	1.786	1.794**	1.169
US – Italy	0.400	1.756*	2.490***
US – Netherlands	0.507	2.072**	2.296**
US – Switzerland	0.708	1.285	1.606*
US – UK	1.006	1.343	2.011**
US – Sweden	0.149	2.104**	1.845*

Notes: $\hat{\beta}_{AIV}$ denotes the slope coefficient from the cointegration equation. S_{nc} denotes the test statistic for the null hypothesis of stochastic cointegration against the alternative of no cointegration. S_{hc} denotes the test statistic for the null hypothesis of stationary cointegration against the alternative of heteroscedastic cointegration. ***, **, * indicate significance at the 1%-, 5%-, and 10%-level respectively. The relationships, for which the null of stochastic cointegration is not rejected, are shaded in grey. Thus shaded cell indicates the presence of a cointegration relationship between the given two markets. The country indicated in a row is the dependent variable. The country indicated in a column is the independent variable.

US and Asia-Pacific markets

The stochastic cointegration test finds higher number of cointegration relations in the group of Asian markets than either Engle and Granger (1987) or Gregory and Hansen (1996) tests (Table 27). Two relationships at 1% level of significance have been detected: one with Australia and another one with Hong Kong. Both of the relations appear to be stationary cointegration relations. Notably, Gregory and Hansen (1996) test also finds evidence in favour of cointegration between the US and Australia. The absence of stable long-run relationships implies presence of diversification benefits among the US and Japanese and the US and Singaporean markets.

Table 27: Stochastic Cointegration Test Results: Asia-Pacific Bivariate Relationships

Countries	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}
US – Australia	1.239	0.206	0.190
US – Hong Kong	1.190	1.236	1.174
US – Japan	0.389	2.428***	2.168**
US – Singapore	-0.045	2.396***	0.895

Notes: $\hat{\beta}_{AIV}$ denotes the slope coefficient from the cointegration equation. S_{nc} denotes the test statistic for the null hypothesis of stochastic cointegration against the alternative of no cointegration. S_{hc} denotes the test statistic for the null hypothesis of stationary cointegration against the alternative of heteroscedastic cointegration. ***, **, * indicate significance at the 1%-, 5%-, and 10%-level respectively. The relationships, for which the null of stochastic cointegration is not rejected, are shaded in grey. Thus shaded cell indicates the presence of a cointegration relationship between the given two markets. The country indicated in a row is the dependent variable. The country indicated in a column is the independent variable.

European markets

When we consider bivariate relations between the European markets, we find a several cointegration relations (Table 28). We find bivariate relations between the French and German, UK and Swedish markets (for the first two – in both directions). The Dutch index seems to affect German, Belgian, and Swedish real estate indices, with all three relationships running in both directions. Like in case of Engle and Granger (1987) and Gregory and Hansen (1996) tests, Sweden appears to be affected by several European markets: French, German, Italian, and Swiss. It is therefore probably not an optimal choice for a diversified real estate portfolio. Interestingly, accordingly to the stochastic cointegration test, German and UK markets appear to be less isolated than accordingly to the results of the other two bivariate cointegration tests. In fact, the German market shows relations with all of the European markets in at least one direction. The UK securitized real estate index seems to be linked to French and Swiss indices. All in all we find 20 out of possible 56 bivariate cointegration relations among the eight markets, which amount to about 36%. In case of Engle and Granger (1987) and Gregory and Hansen (1996) tests the figure stands at 13% and 16% respectively.

Thus stochastic cointegration test supplies much stronger evidence in favour of integration among the European markets than either Engle and Granger (1987) or Gregory and Hansen (1996) test. However, even this number is not suggestive of extremely tight integration among these markets which might have been expected based on the findings of integration among the EU stock markets (Rangvid (2001), Aggarwal et al. (2005), Hardouvelis et al. (2006)).

Table 28: Stochastic Cointegration Test Results: European Bivariate Relationships

	Belgium	France	Germany	Italy	Netherlands	Switzerland	UK	Sweden
Belgium		1.753^{**} 1.296^{*}	1.021 0.645	1.614^{**} 0.066	0.192 0.810	2.213^{**} 0.452	1.915^{**} 0.757	1.228 0.870
France	2.516^{***} 1.982^{**}		0.591 1.624^{*}	2.233^{***} 0.240	2.680^{***} 2.058^{**}	2.776^{***} 1.058	1.020 1.997^{**}	1.411^{*} 2.127^{**}
Germany	0.550 1.159	0.513 1.963^{**}		1.004 0.978	0.431 1.325^{*}	0.760 2.068^{**}	0.855 0.646	0.692 0.607
Italy	2.249^{**} 0.010	2.121^{**} 0.344	2.513^{***} 0.176		2.228^{**} 0.158	2.102^{**} 1.088	1.988^{**} 1.706^{**}	1.292^{*} 2.272^{**}
Netherlands	0.196 0.826	2.312^{**} 1.524*	0.484 0.924	2.205^{**} 0.237		2.631^{***} 0.281	2.250^{**} 0.025	1.863^{**} 2.054^{**}
Switzerland	2.763^{***} 2.391^{***}	2.412^{***} 2.021^{**}	1.405^{*} 1.003	1.957^{**} 1.221	2.655^{***} 2.487^{***}		0.574 2.696^{***}	2.215^{**} 2.233^{**}
UK	1.907^{**} 0.668	1.048 2.118^{**}	1.751^{**} 1.643^{*}	2.235^{**} 0.334	1.943^{**} 0.815	0.644 2.821^{***}		1.895^{**} 0.957
Sweden	1.426^{*} 2.336^{***}	1.115 2.361^{***}	1.200 1.022	1.114 2.379^{***}	1.624^{*} 2.704^{***}	1.561^{*} 2.404^{***}	1.536^{*} 1.563^{*}	

Notes: The number in the first line of each cell denotes S_{nc} - the test statistic for the null hypothesis of stochastic cointegration against the alternative of no cointegration. The number in the second line denotes. S_{hc} - the test statistic for the null hypothesis of stationary cointegration against the alternative of heteroscedastic cointegration. ***, **, * indicate significance at the 1%-, 5%-, and 10%-level, respectively. The relationships, for which the null of stochastic cointegration is not rejected, are shaded in grey. Thus shaded cell indicates the presence of a cointegration relationships between the given two markets. The country indicated in a row is the dependent variable. The country indicated in a column is the independent variable.

Asia-Pacific markets

The four Asian markets display no bivariate relations over the sample period (Table 29) with the possible exception of Japan and Hong Kong (in one direction), for which the null of stochastic cointegration has been rejected at the 10% level of significance only. In case of Asian markets the stochastic cointegration test supplies the weakest evidence in favour of cointegration. The Engle and Granger (1987) and Gregory and Hansen (1996) test both find six relationships for this group of markets.

Table 29: Stochastic Cointegration Test Results: Asia-Pacific Bivariate Relationships

Countries	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}
Australia – Hong Kong	0.695	1.917**	2.583***
Australia – Japan	0.316	2.428***	0.598
Australia – Singapore	-0.244	1.878**	1.753**
Hong Kong – Australia	1.044	1.661**	1.400*
Hong Kong – Japan	0.834	1.482*	2.434***
Hong Kong – Singapore	0.414	2.540***	1.413*
Japan – Australia	0.780	2.786***	1.167
Japan – Hong Kong	0.609	2.317***	1.735**
Japan – Singapore	0.444	2.266***	1.783**
Singapore – Australia	0.255	1.995**	0.351
Singapore – Hong Kong	0.386	2.050**	0.066
Singapore – Japan	1.079	2.077**	0.928

Notes: $\hat{\beta}_{AIV}$ denotes the slope coefficient from the cointegration equation. S_{nc} denotes the test statistic for the null hypothesis of stochastic cointegration against the alternative of no cointegration. S_{hc} denotes the test statistic for the null hypothesis of stationary cointegration against the alternative of heteroscedastic cointegration. ***, **, * indicate significance at the 1%-, 5%-, and 10%-level respectively. The relationships, for which the null of stochastic cointegration is not rejected, are shaded in grey. Thus shaded cell indicates the presence of a cointegration relationship between the given two markets. The country indicated in a row is the dependent variable. The country indicated in a column is the independent variable.

Regional indices

Finally, we test for bivariate integration across the regions using three regional indices: Asia-Pacific, European and North American regional indices (Table 30). We find evidence of the null being not rejected only in case of the Asia-Pacific and North American indices with the Asia-Pacific index being the dependent variable. For the pairs Developed Europe – North America (in both direction) and Asia-Pacific – Developed Europe (in one direction) we find that the null is rejected at the 10% level. Given these results we conclude that there is one certain relation among the three regional indices, the one Asia-Pacific and North American indices. Finding of one clear cointegration relation for the group of regional indices is in line with the results of the Engle and Granger (1987) and Gregory and Hansen (1996) tests. This result suggests weak international integration as measured by the regional indices.

Table 30: Stochastic Cointegration Test Results: International Bivariate Relationships

Countries	$\hat{\beta}_{AIV}$	S_{nc}	S_{hc}
North America – Asia Pacific	1.323	1.391*	2.269**
Asia Pacific – North America	0.651	1.221	2.141*
North America – Developed Europe	1.018	1.692*	2.138**
Developed Europe – North America	0.720	1.566*	2.342*
Asia Pacific – Developed Europe	0.667	1.658*	0.625
Developed Europe – Asia Pacific	1.232	1.979**	0.958

Notes: $\hat{\beta}_{AIV}$ denotes the slope coefficient from the cointegration equation. S_{nc} denotes the test statistic for the null hypothesis of stochastic cointegration against the alternative of no cointegration. S_{hc} denotes the test statistic for the null hypothesis of stationary cointegration against the alternative of heteroscedastic cointegration. ***, **, * indicate significance at the 1%-, 5%-, and 10%-level respectively. The relationships, for which the null of stochastic cointegration is not rejected, are shaded in grey. Thus shaded cell indicates the presence of a cointegration relationship between the given two markets. The country indicated in a row is the dependent variable. The country indicated in a column is the independent variable.

Stochastic cointegration test results also suggest differing degrees of integration across the three considered regions. In short, we find stronger evidence of cointegration among the eight European markets than using Engle and Granger (1987) or Gregory and Hansen (1996) cointegration tests. The stochastic cointegration test does not find conclusive evidence of integration among the two North-American markets, a result in line with that of Engle and Granger (1987) test. The test also finds weaker evidence of integration among the Asian markets than either Engle and Granger (1987) or Gregory and Hansen (1996) test. The results of for the US and European, US and Asia-Pacific and regional indices are generally comparable across the three tests. Summing this up, we conclude that dropping the assumption of homoscedasticity does not lead to finding larger number of cointegration relations in most of the cases, except for the case of the eight European markets. The conclusion of the previous studies regarding the low extent of integration is therefore likely not be driven by this assumption, except for perhaps European markets.

6 Conclusion and Implications

In this paper, long-run co-movements between international securitized real estate markets are analyzed. The examination is based on cointegration methodology and to our knowledge it is the first analysis that applies time-varying and stochastic cointegration analysis to international securitized real estate markets. We compare the results to traditional bivariate and multivariate cointegration frameworks suggested by Engle and Granger (1987) and Johansen (1988).

By explicitly controlling for structural breaks in the cointegration relationships and by considering time-varying cointegration as well as stochastic cointegration when analyzing long-run co-movements between 14 international securitized real estate markets, the main results and the contribution to existing literature are as follows: First, the analysis covers a longer time period than pervious studies that ranges from 1990 to 2009. Thus, the period is not only characterized by fast growing and upward moving real estate stock markets as many previous studies but also by the period of the current and still ongoing financial crisis that started in 2007. Second, the detected cointegration relationships are much stronger between national markets within one economic and geographic region than between national markets located in different regions. Thus, from an investor's point of view, the results indicate that broadening the investment horizon from the domestic continent to others regional markets might be more beneficial than diversifying within one region. This conclusion applies particularly to investors located in Europe. Third, it is shown that most cointegration relationships are unstable and time-varying and that the results from traditional cointegration methodologies suggested by Engle and Granger (1987) and Johansen (1988) might be misleading in that common long-run co-movements are time-varying and are much stronger when structural breaks are considered. A summary of the cointegration tests based on different methodological approaches is provided by Table 31. Fourth, there is no evidence for increasing integration between international securitized real estate markets over time as it is often stated for international stock markets.

Summarizing the empirical results from the analysis, it can be concluded that there exist several long-run relationships between international securitized real estate markets but there still exist vast benefits from international diversification in the long run. Since structural breaks in the cointegration relationships are detected and are of significance the applied methodological framework for analyzing long-run co-movement should take this into consideration and the structural breaks have to be implemented into the analysis. These findings should also be incorporated in practical portfolio management. How to implement the results in practical portfolio management is left for further research.

Table 31: Summary of the Results from Cointegration Tests for the International Securitized Real Estate Markets

	Multivariate Framework			Bivariate Framework		
	Johansen Test	Johansen Test (before 06/2007)	Recursive Johansen Test (R-form)	Engle and Granger (1987) Test	Gregory and Hansen (1996) Test	McCabe et al. (2003) Test
Multivariate Tests, National Indices						
1. North America (US and Canada)	- (0)	+ (1)	- (0)	- (0)	+ (2)	- (0)
2. Asia-Pacific Markets (Australia, Japan, Hong Kong, Singapore)	- (0)	- (0)	- (0)	+ (4)	+ (5)	- (0)
3. European Core Markets (Germany, France, Netherlands, UK)	- (0)	- (0)	- (0)	**n/a**	**n/a**	**n/a**
4. European Periphery Markets (Belgium, Italy, Sweden and Switzerland)	- (0)	- (0)	- (0)	**n/a**	**n/a**	**n/a**
5. All European Markets	+ (1)	+ (2)	+ (1, except for 1998-2000)	+ (4)	+ (9)	+ (21)
6. US and Core European Markets	- (0)	- (0)	+/- (1 before 2003)	**n/a**	**n/a**	**n/a**
7. US and Periphery European Markets	- (0)	- (0)	-/+ (1 after 2001)	**n/a**	**n/a**	**n/a**
8. US and All European Markets	+ (2)	+ (2)	+ (1; 2 except for 1998-2001)	+ (2)	- (0)	+ (2)
9. US and Asia-Pacific Markets	- (0)	- (0)	-/+ (1, 2001-2009)	- (0)	+ (1)	+ (2)
10. All markets (except Canada)	+ (5)	+ (6)	+ (2-8)	**n/a**	**n/a**	**n/a**
Multivariate Tests, Regional Indices						
11. Asia-Pacific, Europe, Northern America	- (0)	- (0)	- (0)	+ (1)	+ (1)	+ (1)

Notes: +(-) denotes presence (absence) of cointegration relationships. The figure below in parentheses indicates the total number of cointegration relationships found in for the group of countries under consideration at at least 5% level of significance, where appropriate. “n/a” indicates relationships which could not been calculated within a given approach.

References

Aggarwal, R., B. M. Lucey, and C. Muckley, 2005. *Dynamics of Equity Market Integration in Europe: Evidence of changes over time and with events* (Institute for International Integration Studies Dublin).

Andrews, D. W. K., 1993. Tests for Parameter Instability and Structural Change with Unknown Change Point, *Econometrica* 61, 821-856.

Ang, A., and G. Bekaert, 2002. International Asset Allocation With Regime Shifts, *Review of Financial Studies* 15, 1137-1187.

Arshanapalli, B., and J. Doukas, 1993. International Stock Market Linkages: Evidence from the Pre- and Post-October 1987 Period, *Journal of Banking & Finance* 17, 193-208.

Bachman, D., J. J. Choi, B. N. Jeon, and K. J. Kopecky, 1996. Common factors in international stock prices: Evidence from a cointegration study, *International Review of Financial Analysis* 5, 39-53.

Baele, L., A. Ferrando, P. Hordahl, E. Krylova, and C. Monnet, 2004. Measuring European Financial Integration, *Oxford Review of Economic Policy* 20.

Bekaert, G., and C. Harvey, 1995. Time Varying World Market Integration, *Journal of Finance* 50, 403-444.

Bond, S. A., and J. L. Glascock, 2006. The Performance and Diversifi cation Benefits of European Public Real Estate Securities, *European Public Real Estate Association* September 2006.

Bond, S. A., G. A. Karolyi, and A. B. Sanders, 2003. International Real Estate Returns: A Multifactor, Multicountry Approach, *Real Estate Economics* 31, 481-500.

Boswijk, H. P., and P. H. Frances, 1992. Dynamic Specification and Cointegration, *Oxford Bulletin of Economics and Statistics* 54, 369-381.

Brounen, D., M. P. Prado, and S. Stevenson, 2008. Kurtosis and Consequences – the Case of International Property Shares, *SSRN Working Paper*.

Campbell, J., and P. Perron, 1991. *Pitfalls and Opportunities: What Macroeconomists Should Know about Unit Roots*(MIT Press, Cambridge, MA).

Campos, J., N. R. Ericcson, and D. F. Hendry, 1996. Cointegration Tests in the Presence of Structural Breaks, *Journal of Econometrics* 70, 187-220.

Chaudhry, M. K., F. C. N. Myer, and J. R. Webb, 1999. Stationarity and Cointegration in Systems with Real Estate and Financial Assets, *Journal of Real Estate Finance and Economics* 18, 339-349.

Cheung, Y.-W., and K. S. Lai, 1993. Finite-sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration, *Oxford Bulletin of Economics and Statistics* 55, 313-328.

Cheung, Y.-W., and K. S. Lai, 1995. Lag Order and Critical Values of the Augmented Dickey-Fuller Test, *Journal of Business and Economic Statistics* 13, 277-280.

Eichholtz, P., R. Huisman, K. Koedijk, and L. Schuin, 1998. Continental Factors in International Real Estate Resturns, *Real Estate Economics* 26.

Eichholtz, P. M. A., 1996a. The Stability of the Covariance of International Property Share Returns, *The Journal of Real Estate Research* 11, 149-158.

Eichholtz, P. M. A., 1996b. Does International Diversification Work Better for Real Estate than for Stocks and Bonds?, *Financial Analysts Journal* 52, 56-62.

Elliot, G., T. J. Rothenberg, and J. H. Stock, 1996. Efficient Tests for an Autoregressive Unit Root, *Econometrica* 64, 813-836.

Engle, R., and C. Granger, 1987. Cointegration and Error Correction; Representation, Estimation and Testing, *Econometrica* 55, 251-276.

Fama, E. F., 1965. The Behaviour of Stock Market Prices, *Journal of Business* 38.

Gallo, J. G., and Y. Zhang, 2009. Global Property Market Diversification, *Journal of Real Estate Financial Economics* forthcoming.
Garcia Pascual, A., 2003. Assessing European Stock Markets (Co)integration, *Economics Letters* 78, 197-203.
Garvey, R., G. Santry, and S. Stevenson, The Linkages between Real Estate Securities in the Asia-Pacific. 2001, *Pacific Rim Property Research Journal* 7, 240-258.
Gelos, G., and R. Sahay, 2000. Financial Spillovers in Transition Economies, *IMF Working Paper* WP/00/71.
Gerlach, R., P. Wilson, and R. Zurbruegg, 2006. Structural Breaks and Diversification: The Impact of the 1997 Asian Financial Crisis on the Integration of Asia-Pacific Real Estate Markets, *Journal of International Money and Finance* 25.
Glascock, J. L., C. Lu, and R. W. So, 2000. Further Evidence on the Integration of REIT, Bond, and Stock Returns, *Journal of Real Estate Finance and Economics* 20, 177-194.
Goetzmann, W. N., L. Li, and K. G. Rouwenhorst, 2001. Long-Term Global Market Correlations, *NBER Working Paper* 8612.
Goetzmann, W. N., L. Li, and K. G. Rouwenhorst, 2005. Long-term Global Market Correlations, *Journal of Business* 78, 1-38.
Gregory, A. W., and B. E. Hansen, 1996. Residual-based Tests for Cointegration in Models with Regime Shifts, *Journal of Econometrics* 70, 99-126.
Grubel, H. G., 1968. Internationally Diversified Portfolios: Welfare Gains and Capital Flows, *The American Economic Review* 58.
Hall, A., 1994. Testing for a Unit Root in Time Series with Pretest Data-Based Model Selection, *Journal of Business and Economic Statistics* 12, 461-470.
Hansen, H., and S. Johansen, 1999. Some Tests for Parameter Constancy in Cointegrated VAR-Models, *Econometrics Journal* 2, 306-333.
Hardouvelis, G. A., D. Malliaropulos, and R. Priestley, 2006. EMU and European Stock Market Integration, *The Journal of Business* 79, 365-392.
Harris, D., B. McCabe, and S. Leybourne, 2002. Stochastic Cointegration: Estimation and Inference, *Journal of Econometrics* 111, 363-384.
Johansen, S., 1988. Statistical Analysis of Cointegration Vectors, *Journal of Economic Dynamics and Control* 12, 251-254.
Johansen, S., and K. Juselius, 1990. Maximum Likelihood Estimation and Inference of Cointegration with Application to the Demand for Money, *Oxford Bulletin of Economics and Statistics* 52, 169-209.
Juselius, K., 2007. *The Cointegrated VAR Model: Methodology and Applications*(Oxford University Press, Oxford).
Kasa, K., 1992. Common Stochastic Trends in International Stock Markets, *Journal of Monetary Economics* 29, 95-124.
Koedijk, K. C. G., R. A. J. Campbell, and P. Kofman, 2002. Increased Correlation in Bear Markets, *Financial Analysts Journal* 58, 87-94.
Kwiatkowski, D., P. C. B. Phillips, P. Schmidt, and Y. Shin, 1992. Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root, *Journal of Econometrics* 54, 159-178.
Ling, D. C., and A. Naranjo, 1999. The Integration of Commercial Real Estate and Stock Markets, *Real Estate Economics* 27, 483-515.
Liow, K. H., 1997. The Historical Performance of Singapore Property Stocks, *Journal of Property Finance* 8, 111-125.
Liow, K. H., 2007. The Dynamics of Return Volatility and Systematic Risk in International Real Estate Security Markets, *Journal of Property Research* 24, 1-29.
Liow, K. H., J. Ooi, and Y. Gong, 2005. Cross-Market Dynamics in Property Stock Markets – Some International Evidence, *Journal of Property Investment and Finance* 23, 55-75.

Liow, K. H., and M. C. Sim, 2006. he Risk and Return Profile of Asian Real Estate Stocks, *Pacific Rim Property Research Journal* 12, 283-310.
Liow, K. H., and H. Yang, 2005. Long-Term Co-Memories and Short-Run Adjustment: Securitized Real Estate and Stock Markets, *The Journal of Real Estate Finance and Economics* 31, 283–300.
Lizieri, C., P. McAllister, and C. Ward, 2003. Continental Shift? An Analysis of Convergence Trends in European Real Estate Equities, *Journal of Real Estate Research* 25, 1-21.
Longin, F., and B. Solnik, 1995. Is the Correlation in International Equity Returns Constant: 1960-1990?, *Journal of International Money and Finance* 14, 3-26.
Longin, F., and B. Solnik, 2001. Extreme Correlation of International Equity Markets, *The Journal of Finance* 56, 649-676.
Lucey, B. M., and S. Voronkova, 2008. Russian Equity Market Linkages before and after the 1998 crisis: Evidence from Stochastic and Regime-Switching Cointegration Tests, *Journal of International Money and Finance* 27, 1303-1324.
MacKinnon, J. G., A. A. Haug, and L. Michelis, 1999. Numerical Distributions of Likelihood Ratio Tests for Cointegration, *Journal of Applied Econometrics* 14, 563-577.
Maddala, G. S., and I.-M. Kim, 1998. *Unit Roots, Cointegration and Structural Change*(Cambridge University Press, Cambridge).
Mandelbrot, B. B., 1963. The Variation of Certain Speculative Prices, *Journal of Business* 34, 393-413.
McCabe, B., S. Leybourne, and D. Harris, 2003. Testing for Stochastic Cointegration and Evidence for Present Value Models, *www.RePEc:wpa:wuwpem:0311009*.
Morawski, J., H. Rehkugler, and R. Füss, 2008. The Nature of Listed Real Estate Companies: Property or Equity Market?, *Financial Markets and Portfolio Management* 22, 101-126.
NAREIT, 2009. *The Investor's Guide to REITs. NAREIT's Guide to the Real Estate Investment Trust Industry*(Washington D.C.).
Newell, G., and K. W. Chau, 1996. Linkages between Direct and Indirect Property Performance in Hong Kong, *Journal of Property Finance* 7, 9-29.
Ogaki, M., and J. Y. Park, 1997. A Cointegration Approach to Estimating Preference Parameters, *Journal of Econometrics,* 82, 107-134.
Oikarinen, E., M. Hoesli, and C. Serrano, 2009. Linkages between Direct and Securitized Real Estate, *Swiss Finance Institute Research Paper Series* N°09-26.
Okunev, J., and W. P. J., 1997. Using Non-Linear Tests to Examine Stock Market Integration Between Real Estate and Stock Markets, *Real Estate Economics* 25, 487-503.
Phillips, P. C. B., and P. Perron, 1988. Testing for a Unit Root in Time Series Regression, *Biometrica* 75, 395-346.
Rangvid, J., 2001. Increasing Convergence among European Stock Markets? A Recursive Common Stochastic Trends Analysis, *Economic Letters* 71, 383-389.
Schindler, F., 2009a. Correlation Structure of Real Estate Markets over Time, *Journal of Property Investment and Finance* 27, 579-592.
Schindler, F., 2009b. Long-Term Benefits from Investing in Asia-Pacific Real Estate, *Pacific Rim Property Research Journal* 15, 335-357.
Schindler, F., 2010. Long-term Benefits from Investing in International Securitized Real Estate, *International Real Estate Review* forthcoming.
Schwert, G. W., 1989. Tests for Unit Roots: A Monte Carlo Investigation, *Journal of Business and Economic Statistics* 7, 147-159.
Seiler, M. J., J. R. Webb, and M. F. C. N., 1999. Are EREITs Real Estate?, *Journal of Real Estate Portfolio Management* 5, 171-181.
Serrano, C., and M. Hoesli, 2009. Global Securitized Real Estate Benchmarks and Performance, *Journal of Real Estate Portfolio Management* 15, 1-19.

Serrano, C., and M. Hoesli, 2010. Are Securitized Real Estate Returns More Predictable than Stock Returns?, *Journal of Real Estate Finance and Economics* 41, 170-192.

Urzúa, C. M., 1996. On the Correct Use of Omnibus Tests for Normality, *Economics Letters* 53, 247-251.

Westerheide, P., 2006. Cointegration of Real Estate Stocks and REITs with Common Stocks, Bonds and Consumer Price Inflation – an International Comparison, *ZEW Discussion Paper* No. 06-057.

Worzala, E., and C. F. Sirmans, 2003. Investing in Internationall Real Estate Stocks: a Review of the Literature, *Urban Studies* 40, 1115-1149.

Yang, J., J. W. Kolari, and G. Zhu, 2005. European Public Real Estate Market Integration, *Applied Financial Economics* 15, 895-905.

Yunus, N., 2009. Increasing Convergence between U.S. and International Securitized Property Markets: Evidence Based on Cointegration Tests, *Real Estate Economics* 37, 383-411.

Yunus, N., and P. E. Swanson, 2007. Modelling Linkages between US and Asia-Pacific Securitized Property Markets, *Journal of Property Research* 24, 95-122.

Zivot, E., and D. W. K. Andrews, 1992. Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis, *Journal of Business & Economic Statistics* 10, 251-270.