J
R
E
P
M

Are Public and Private Asset Returns and Risks the Same? Evidence from Real Estate Data

Martin Hoesli
Elias Oikarinen

Executive Summary. Real estate constitutes a good laboratory to investigate the similarity of public and private asset returns and risks. We find evidence of a one-to-one long-term relation between public and private real estate performance. Also, the return volatilities do not differ significantly between the public and private markets regardless of investment horizon. The findings have important implications for portfolio management: (1) public and private real estate are close substitutes in a portfolio with a several-year investment horizon and (2) public real estate-related ETFs and derivatives are useful to hedge risks associated with direct real estate holdings or lenders' mortgage inventory.

Publicly traded securities represent indirect claims on lumpy privately traded assets such as factories and equipment or real estate.[1] Therefore, it could be expected that the returns and risks of privately traded direct investments and of securities that are based on similar direct assets are alike, at least in the long run and after catering for the effects of leverage and management costs. After all, the security cash flows are generated from the underlying direct assets (i.e., the expected cash flows and their volatility should be similar). Nevertheless, due to factors such as higher liquidity and lower transaction costs of the securitized assets traded in public market places, the returns on securities may notably deviate from those on private assets.[2] In particular, a lower liquidity premium and lower transaction costs could induce a lower required (and therefore also expected) return on securitized assets. Also the diversification benefits offered by securities versus direct assets can differ, at least in the relatively short term, possibly affecting the required rates of return. In addition, it is well known that the public asset markets are notably more informationally efficient than their private counterparts. Therefore, it is essentially an empirical question to examine whether the trading "platform" influences the asset returns and return volatilities.

The question of whether publicly traded securitized assets provide similar overall returns and return volatilities as privately traded direct investments is of great importance for asset and risk management. For one, the equivalence of returns and return volatilities would indicate that publicly traded assets are close substitutes for the underlying private assets in an investment portfolio. In addition, a tight connection between the returns would suggest that investors, banks, and other financial institutions having a private investment market exposure by either holding private assets directly or through their outstanding lending inventory can use derivatives on public assets to hedge the risk exposure arising from private investment portfolios.

An empirical examination of the question is usually not possible, since there are no reliable time series data on the typical privately traded underlying assets. The "duality" of the real estate markets offers an opportunity to test whether securitized asset returns reflect the performance of underlying private assets: relatively reliable data are available both for securitized and direct real estate performance. In addition to providing an opportunity to study the equivalence of public and private asset performance, a study using real estate data is of substantial interest per se given the attractiveness of real estate assets for a great number of investors (e.g., Hudson-Wilson, Fabozzi, and Gordon, 2003; Andonov, Eichholtz, and Kok, 2015) and thereby the significant role of real estate investments in, for instance, institutional investors' portfolios, and because of the importance of real estate in the credit institutions' and households' balance sheets.

Despite the importance of the topic, the literature on the similarity of public and private asset performance is scarce. Geltner and Kluger (1998), Pagliari, Scherer, and Monopoli (2005), Riddiough, Moriarty, and Yeatman (2005), and Ling and Naranjo (2015) compare the mean returns for real estate assets in the market in the United States controlling for the property-type mix and the leverage in real estate investment trust (REIT) indices. Geltner and Kluger's (1998) results indicate that REIT portfolio returns and volatilities were generally greater than those in the private market during the 1987–1996 period. Pagliari, Scherer, and Monopoli (2005) conclude that REIT and direct real estate returns did not differ from each other, in the statistical sense, over the 1981–2001 period when also controlling for appraisal smoothing in the private market index. Furthermore, they accept the hypothesis of equivalent REIT and direct real estate return variances based on an F-test. Riddiough, Moriarty, and Yeatman (2005) account for the influence of management fees in addition to property-type mix and leverage. They report a three percentage point difference between REIT and direct real estate returns during 1980–1998, but do not test for the statistical significance of this difference. Ling and Naranjo (2015) present evidence of the REIT market outperforming the private real estate market in the office and retail sectors over the period 1994–2011. In the multifamily and industrial sectors, they observe the opposite. While these authors control for property type, leverage, and management fees, they do not test for the statistical significance of the return differences.

In addition, pairwise cointegration tests have been applied in several studies to investigate the long-term dynamics between various assets and the substitutability of different assets with respect to diversification benefits and to exposure to market fundamentals. Oikarinen, Hoesli, and Serrano (2011), Boudry, Coulson, Kallberg, and Liu (2012), Hoesli and Oikarinen (2012), and Yunus, Hansz, and Kennedy (2012) report evidence of cointegration between REIT and direct real estate markets in several countries. Most of these studies do not control for the leverage or property type issues, and none of them consider the influence of management costs. Moreover, none of the studies formally test for the equivalence of the long-run returns.

In sum, the previous investigations are based on simple F-tests on return and volatility correspondence or do not formally test for the similarity. We are the first to propose an alternative method, i.e., cointegration analysis, to test for the equivalence of the returns over the long run. Cointegration analysis allows us to test formally for a long-term one-to-one relation between REIT and direct real estate total return indices. This analysis has several advantages over the simple F-test on means that have been used in a few previous studies on the topic. For example, a one-to-one cointegrating relation between two series implies that there are tight economic forces keeping the series together in the long run and therefore an observed equivalence of the mean returns is not just a coincidence that is likely to vanish in the future, but is due to a real economic phenomenon (i.e., due to an equilibrium relation between the series).

In addition to the different statistical methodology used to investigate the long-run return similarity, our analysis includes several other notable contributions compared to extant studies. First, we use the readily available direct real estate and REIT sector level indices for the U.S. and recent data that include the Global Financial Crisis (GFC) period of the 2000s. Besides being able to investigate the impacts of the GFC on the relation between private and public real estate performance, using recent data is deemed desirable, since structural changes in the REIT market in the early 1990s may have significantly affected the linkages between private and public real estate (Clayton and MacKinnon, 2001, 2003; Pagliari, Scherer, and Monopoli, 2005). Second, we study the U.K. market in addition to the U.S. market. Third, we are the first to rigorously compare the return volatilities at different investment horizons. This is important, as real estate investments generally exhibit notable horizon effects due to the positive autocorrelation in price growth; thus the planned investment horizon may significantly affect the conclusions (Campbell and Viceira, 2005). The volatility comparison is based on variance ratio computations using the Wild bootstrap approach (Kim, 2006). Fourth, we conduct robustness checks for the results and examine the stability of the long-term relations during the sample period. This will, for instance, make it possible to gauge whether such relationship has been altered by the GFC.

Given the well documented fact that public real estate returns co-move more with the general stock market returns than with the underlying real estate performance over the relatively short horizon but public and private real estate appear to provide similar diversification benefits in a mixed-asset portfolio in the longer horizon (e.g., MacKinnon and Al Zaman, 2009; Yunus, Hansz, and Kennedy, 2012), it is reasonable to concentrate on long-term returns in our analysis. That is, it is at the several-year investment horizon that public and private real estate investments can be considered as substitutes if the returns and return volatilities are similar. In addition, private market data complications may distort the short-term analysis, and investors who consider investing in direct real estate typically have an investment horizon of several years due to the illiquidity and large transaction costs in the private real estate market. Thus, our results are relevant with respect to long-term strategic asset allocation, in particular.

Our analysis is based on U.S. data for the period 1994–2011 and U.K. data for the period 1991–2011. We cater for leverage, property-type mix, and management fees, and detect a long-term one-to-one relation between the public and private markets in four out of the six real estate sectors included in the analysis. The return volatilities generally do not differ significantly regardless of sector and time horizon. Thus, our findings indicate that securitized and direct real estate investments generally can be considered to work as substitutes in an investment portfolio with an investment horizon of several years. Given the equivalence of returns and volatilities, and similar diversification benefits in the long term, REITs may be a better option than private real estate for an investor for whom the market liquidity and transaction costs are of notable importance. The results also support the suggestion in the literature that house price-linked financial assets such as REITs can be used to take housing exposure in a more continuous manner than by using the direct market only, thereby achieving welfare gains (Kraft and Munk, 2011).

Furthermore, the close linkages between REIT and direct real estate returns suggest that REIT-related ETFs and derivatives offer opportunities to hedge risks brought about by direct real estate holdings. This is particularly important for banks and other financial institutions that are not actual real estate investors, but are significantly exposed to the private real estate market through their mortgage lending. The use of such hedging strategies could help banks to survive better during periods of economic distress and drastically decreasing real estate prices, and could diminish, to some extent, the illiquidity risk of privately traded assets such as direct real estate investments. Finally, as the duality of the real estate markets offers a rare opportunity to study empirically the correspondence of public and private asset performance, the results can provide insights for the broader investment universe as well.

Data

For the U.S., we include four real estate sectors (apartments, offices, industrial, and retail) and for the U.K. two sectors (offices and retail) in the analysis. For securitized real estate, the FTSE/NAREIT Equity REIT sector level indices are used for the U.S., whereas for the U.K. we construct the REIT indices from the company level price, dividend, and market cap data provided by EPRA.[3] While the sector level direct real estate indices for the U.S. are transaction-based NCREIF (TBI) indices, the IPD indices we use for the U.K. are appraisal-based. Both the REIT and private indices we use can be considered as well diversified and thereby broad measures of the market performance. The sample period is 1994:Q1–2011:Q4 for the U.S. and 1991:Q1–2011:Q4 for the U.K., and all the real estate indices employed in the analysis are total return indices. The data frequency is constrained by the direct market data. The sample period spans over several business cycles thus making it possible to make conclusions about longer-term linkages between the markets. The use of sector level indices enables us to control for index compositional differences. This is important given that the overall direct, and securitized real estate indices typically differ notably with respect to the property-type mixes and because the return dynamics and performance between various real estate sectors may vary substantially (Yavas and Yildirim, 2011; Hoesli and Oikarinen, 2012).

While the REIT returns are net of portfolio-level management fees, such fees are not deducted from the TBI and IPD returns. Therefore, to make the returns comparable, we need to deduct portfolio-level management costs from the TBI and IPD data. According to Riddiough, Moriarty, and Yeatman (2005), these fees range between 50 and 120 bps per year. We follow Riddiough, Moriarty, and Yeatman (2005) and Ling and Naranjo (2015), and use an annual 80 bps assumption, i.e., deduct 0.2% from the quarterly TBI and IPD returns, in our baseline computations. This assumption also is well in line with the management fees that Andonov, Eichholtz, and Kok (2015) report for pension funds in a range of countries. In any matter, we conduct robustness checks using the 50 bps and 120 bps assumptions.

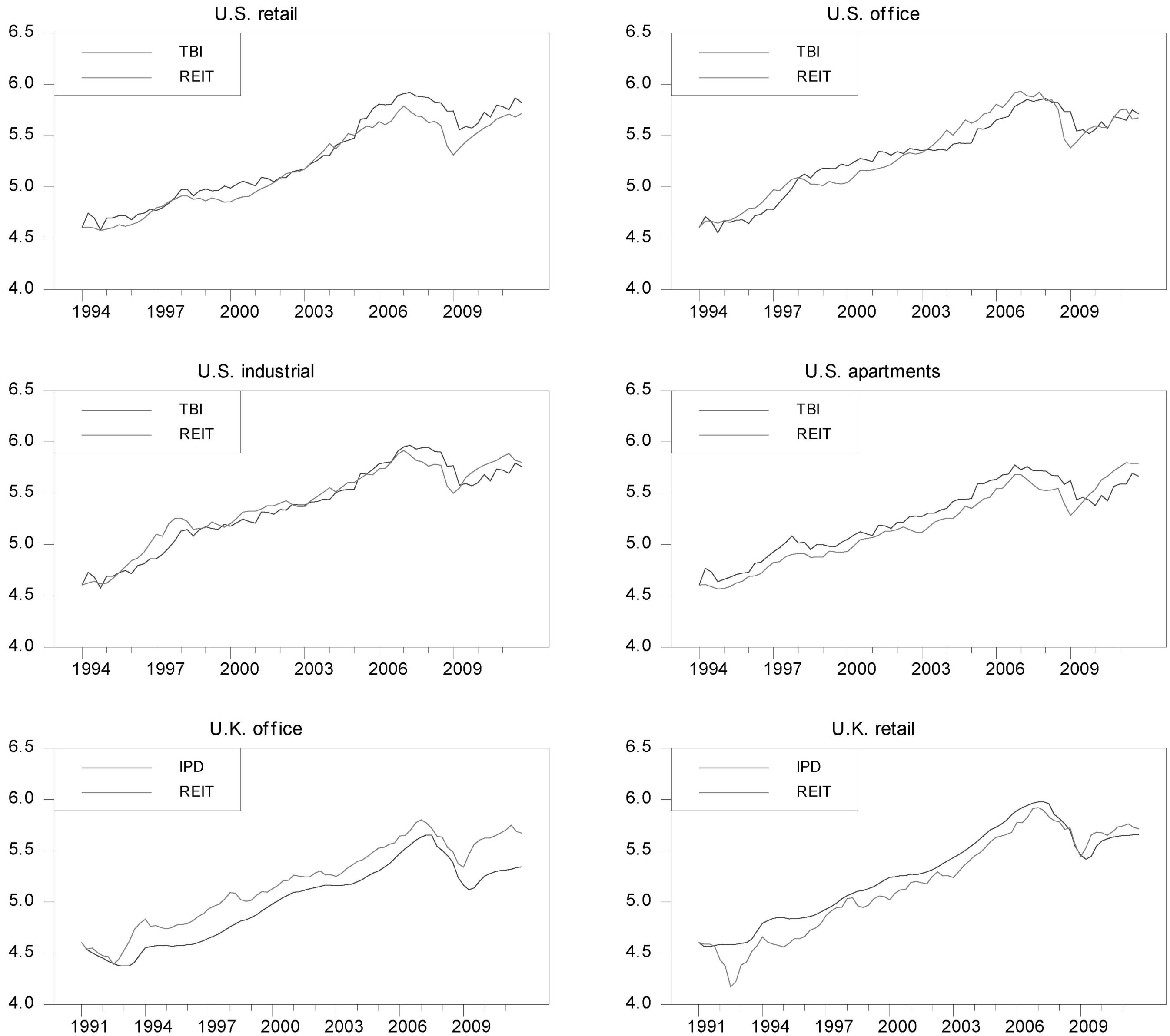
Moreover, while REIT returns include the impact of leverage, the direct market indices consist of unleveraged properties. The magnitude of leverage naturally affects the mean and volatility of the returns. Therefore, we restate the REIT returns for the effect of leverage. Similar to Pagliari, Scherer, and Monopoli (2005), the unlevered returns are computed using the following formula that is based on the well-known proposition of Modigliani and Miller (1958):

$$r_{uit} = r_{eit}(1 - LTV_{it}) + r_{dt}LTV_{it}, \qquad (1)$$

where r_{uit} is the unlevered REIT return of sector i in period t, r_{eit} is the return on equity of REIT sector i in period t, r_{dt} is the cost of debt in period t, and LTV_{it} is the loan-to-value ratio of sector i in period t. The quarterly leverage time series data are provided by NAREIT. The average leverage of U.S. REITs during the sample period is 43% in the industrial sector, 47% in the apartment sector, 48% in the office sector, and 51% in the retail sector. The leverage is quite volatile, being at the lowest around 30% in the mid-1990s and at the highest some 70%–75% in 2009. In the U.K., the mean leverage ratios are similar to the U.S. (49% for retail and 54% for office) but less volatile with the minimum being 40% and maximum less than 70%. The cost of debt used in the computations is the corporate bond middle rate for the corresponding country sourced from Datastream.

After the aforementioned data adjustments, we deflate the indices by the U.S./U.K. Consumer Price Index and take natural logs of the real indices. Exhibit 1 shows the real unlevered REIT indices and

Exhibit 1 | Sector Level (unlevered) REIT and Private Real Estate Real Total Return Indices



the real direct market indices net of 80 bps management fees. To give an idea about how the unlevering of REIT returns affects the REIT series, Exhibit A1 in the Appendix presents the unlevered REIT indices together with the levered (original) ones.

The TBI indices for the U.S. are "quality-adjusted" transaction-based indices, as they are calculated by means of the price changes between the appraised values of properties two quarters ago and the transaction prices of properties that have been sold during a given quarter. The quality control is achieved in that the characteristics of properties in the current quarter and two quarters ago should be very similar. The IPD indices that we use for the U.K. are also constant-quality; however, they exhibit appraisal smoothing as they are solely appraisal-based. Smoothing refers to the fact that the index values and returns exhibit high levels of serial correlation as a result of appraisers largely relying on the past value when estimating the contemporaneous value of properties (Clayton, Geltner, and Hamilton, 2001). Therefore, for the purpose of the tests conducted on volatility equivalence, we desmooth the

IPD returns. We use a simple reverse filter to uncover desmoothed returns that exhibit similar levels of serial correlation as in the U.K. transaction-based index developed by Devaney and Martinez Diaz (2011), i.e., approximately 0.3.[4] The desmoothing parameter is set equal to 0.6 to achieve such levels of serial correlation of the desmoothed series. To test the robustness of our results to the value selected for the desmoothing parameter, we also consider parameters of 0.5 and 0.7.

Methodology

Baseline Analysis

A tight long-term relation in terms of cointegration may or may not exist between variables that do or do not "look" cointegrated, and the only way to find out if data are actually cointegrated is through a careful statistical analysis, rather than to rely on visual inspection (Hendry and Juselius, 2000). Hence, we test for cointegration between REIT and TBI/IPD indices using the Johansen (1996) trace test.[5] The cointegration tests are conducted separately for each sector. The vector error correction model (VECM) used in the trace test is the following:

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_k \Delta X_{t-k} + \alpha\beta' X_{t-1} + \Omega D_t + \varepsilon_t, \quad (2)$$

where ΔX_t is $X_t - X_{t-1}$, X_t is a two-dimensional vector of return index values in period t, μ is a two-dimensional vector of drift terms, Γ_i is a 2×2 matrix of coefficients for the lagged differences of the return indices at lag i, k is the number of lags in differences included in the model, α is a vector of the speed of adjustment parameters, β' forms the cointegrating vector (i.e., includes the long-term coefficient estimates on the public and private market indices), and ε is a vector of white noise error terms. The models for the U.S. also include one point dummy variable (D), which takes the value of one in 2008:Q4 and is zero otherwise to cater for the unique outlier observations induced by the collapse of Lehman Brothers. The dummy variable helps to fulfil the assumption of normally distributed residuals and hence to achieve more reliable p-values in the trace test. Similarly, the U.K. models include two point dummy variables, 1992:Q3 and 2008:Q4 for retail, and 2007:Q4 and 2008:Q4 for offices. We also test for the need for seasonal dummy variables in the VECMs. Based on the Schwarz information criterion, seasonal dummies are not needed in any of the VECMs.

The lag length is selected based on the Hannan-Quin information criterion (HQ) as suggested by Johansen, Mosconi, and Nielsen (2000). However, more lags are included if the assumption of no autocorrelation in residuals cannot be accepted by the Lagrange multiplier test at lag length two. As the models include point dummies, we report trace test p-values based on the simulated statistics computed with the program CATS2 (Dennis, 2006). Because asymptotic distributions can be rather bad approximations of the finite sample distributions, the Bartlett small sample corrected values suggested by Johansen (2002) are employed throughout the cointegration analysis.

The null hypothesis in the trace test is that of no cointegration between the variables. If this hypothesis can be rejected at the conventionally used levels of statistical significance, we conclude that the series are cointegrated (i.e., exhibit a tight long-term relation). This is not enough to conclude that the returns are similar over the long run, however. The similarity of returns can be tested by imposing a one-to-one restriction on the cointegrating relation. In case we detect cointegration, we further test for the hypothesis TBI or IPD = REIT in the cointegrating relation. The one-to-one hypothesis is tested by the Bartlett small-sample corrected likelihood ratio (LR) test reported in Johansen (2000). In case either of the variables can be restricted to be weakly exogenous, the test on the one-to-one hypothesis also includes the assumption of weak exogeneity. If the one-to-one restriction cannot be rejected at the conventional significance levels, the hypothesis that REITs and direct real estate for the given property type provide the same mean return over the long horizon is accepted.

Previous studies on the topic have used simple comparisons of the mean returns and corresponding F-tests to investigate the equivalence of REIT and direct real estate returns. Cointegration analysis has several advantages over the F-test. First, the F-test results can be highly dependent on the ending and starting dates of the sample period, especially given that direct real estate prices (returns) appear to react notably slower to changes in the fundamentals than REIT prices (Hoesli and Oikarinen, 2012). This can be problematic particularly if the starting or ending period represents an abnormal time period, such as a financial crisis time. Although the ending date can affect the cointegration results as well, the cointegration analysis is less prone to the complications in relation with sample period timing. This is because such analysis is based on the relation between the variables during the whole sample period, not only on the starting and ending values of return indices (as the F-test essentially is to a large extent), and because the concept of cointegration allows for even large *temporary* deviations from a long-run equilibrium relation.

Second, a one-to-one cointegrating relation between two series indicates that there are strong economic forces keeping the series together in the long run. Stated differently, it means that the equivalence of mean returns is not just a coincidence that is likely to vanish in the future, but is due to a real economic phenomenon (i.e., due to an equilibrium relation between the series). Third, the F-test results are known to be highly dependent on outlier observations and sensitive to the violations of the normality assumption. In cointegration tests, we can add point dummy variables to cater for outliers and thereby fulfill the normality of residuals assumption while still getting reliable test values through simulation. Fourth, cointegration analysis allows us to conduct robustness checks that are not possible with the F-test, in particular the recursive estimation that makes it possible to investigate the stability of the long-term relations. Finally, (abnormally) prominent cycles and thereby return volatility in the sample period, due to a financial crisis for instance, increases the likelihood of accepting the null of similar returns in the F-test. Nevertheless, we also report the conventional F-test p-values for the hypothesis of similar mean returns on REITs and direct real estate.

Regarding the analysis of the similarity of risks, we use the return volatility (i.e., the standard deviation of returns) as the measure of risk. We test the similarity of volatilities in the two markets using the F-test.

Sensitivity Analyses

We conduct a number of robustness checks for the empirical findings. As the assumed direct real estate portfolio management costs can influence the results, we also conduct cointegration analysis and F-tests on the mean returns assuming 50 bps and 120 bps management fee assumptions.

Also, the planned investment horizon (i.e., the employed data frequency) may affect our F-test conclusions. This is particularly relevant regarding the test on the equivalence of return volatilities. Since real estate returns are known to exhibit substantial "momentum" (positive autocorrelation) in the relatively short term and reversion (negative autocorrelation) in the long run and the momentum and reversion patterns can differ between public and private markets, the relative volatilities may be dependent on the assumed investment horizon. Additionally, the direct market volatility may be downgraded in the relatively short term by the time-varying liquidity in the market (Fisher, Gatzlaff, Geltner, and Haurin, 2003; Pagliari, Scherer, and Monopoli, 2005). Therefore, we conduct the F-tests assuming three-year and five-year investment horizons as well (instead of the baseline one-quarter horizon).

A complication with the longer-horizon F-test statistics is the loss of observations in the early sample period. This may significantly affect the results especially in those cases where there is substantial volatility in the indices during the early sample period (such is the case for the U.K. retail sector). This gives another reason to rely more on the cointegration statistics rather than on the F-statistics of the mean returns in the long-horizon analysis.

We also compute variance ratios (VRs) for the returns to illustrate the impact of the investment horizon on the riskiness of the markets. Given that the number of observations is relatively small, the wild bootstrap approach of Kim (2006) is used to compute the VRs and their standard deviations. This approach has better small sample properties than the conventional VR statistics. The VRs and standard deviations are then used to compute and graph the asset volatilities and their confidence bands at each investment horizon up to 20 quarters. The standard deviation of market i at the x quarter horizon is calculated as $\sigma_i(1) * VR_i(x)$, where σ denotes standard deviation.

As a diagnostic check regarding the cointegration analysis, we examine the stability of the long-term relations by the recursive and backwards recursive max test statistics (in the R-form) of constancy of the estimated long-run relation (Juselius, 2006). This will, for instance, make it possible to gauge whether such a relation has been altered by the GFC.

For the U.K., we use the original appraisal-based IPD indices in the cointegration analysis and in the F-tests for mean returns. This is because appraisal-smoothing should not affect the long-term relations between the public and private markets. In the volatility tests, however, we use the desmoothed IPD returns, which should provide a much more reliable measure of direct real estate risk. As the volatility comparison depends on the assumed first-order serial correlation and thereby on the imposed desmoothing parameter, we conduct robustness checks using desmoothing parameters of 0.5 and 0.7 (the parameter being 0.6 in the baseline analysis).

Empirical Findings

Baseline Analysis

Exhibits 2 and 3 report the mean returns, return volatilities, and other descriptive statistics for the real unlevered asset returns for the U.S. and U.K., respectively. While the observed first-order autocorrelations are positive for REIT and IPD returns, they are negative (although not statistically significant) for TBI returns. The negative first-order autocorrelations of TBI returns are likely due to short-term measurement error in the TBI indices. The exhibits also provide the F-test p-values for the hypothesis of equivalent returns and return volatilities between the public and private real estate markets.

Exhibit 2 | Baseline Descriptive and F-test Statistics: U.S. Market Returns

	Mean	Std. Dev.	Jarque-Bera	Autocorrelation
Panel A: Quarterly Returns				
Retail TBI	0.017	0.058	0.01	−0.152
Retail REIT	0.016	0.045	0.00	0.335**
F-test (p-value)	.860	.030		
Office TBI	0.016	0.054	0.00	−0.144
Office REIT	0.015	0.057	0.00	0.316**
F-test (p-value)	.950	.660		
Industrial TBI	0.016	0.057	0.00	−0.158
Industrial REIT	0.017	0.049	0.00	0.343**
F-test (p-value)	.950	.180		
Apartment TBI	0.015	0.056	0.00	−0.204
Apartment REIT	0.017	0.041	0.00	0.399**
F-test (p-value)	.830	.010		
Panel B: Three-year Returns				
Retail TBI	0.202	0.235	0.73	
Retail REIT	0.187	0.208	0.08	
F-test (p-value)	.710	.350		
Office TBI	0.192	0.217	0.17	
Office REIT	0.174	0.248	0.00	
F-test (p-value)	.680	.320		
Industrial TBI	0.192	0.231	0.01	
Industrial REIT	0.200	0.194	0.94	
F-test (p-value)	.840	.190		
Apartment TBI	0.160	0.200	0.00	
Apartment REIT	0.192	0.153	0.00	
F-test (p-value)	.320	.040		
Panel C: Five-year Returns				
Retail TBI	0.382	0.246	0.45	
Retail REIT	0.336	0.235	0.29	
F-test (p-value)	.330	.760		
Office TBI	0.349	0.198	0.21	
Office REIT	0.332	0.266	0.07	
F-test (p-value)	.570	.040		
Industrial TBI	0.356	0.217	0.03	
Industrial REIT	0.324	0.184	0.20	
F-test (p-value)	.420	.240		
Apartment TBI	0.292	0.205	0.05	
Apartment REIT	0.326	0.119	0.26	
F-test (p-value)	.290	.000		

Notes: Jarque-Bera denotes the Jarque-Bera test for normally distributed returns. Autocorrelation is not reported for the longer-run returns, since these returns are computed on an overlapping window basis. The TBI values are based on an 80 bps management cost assumption.
*Significant at the 5% level.
**Significant at the 1% level.

The F-test results clearly provide support for the hypothesis of similar mean returns for all tested horizons for each sector, except for U.K. retail. Also the hypothesis of similar return volatilities is generally

Exhibit 3 | Baseline Descriptive and F-test Statistics: U.K. Market Returns

	Mean	Std. Dev.	Jarque-Bera	Autocorrelation
Panel A: Quarterly Returns				
Retail IPD	0.013	0.062	0.00	0.300**
Retail REIT	0.013	0.059	0.00	0.320**
F-test (p-value)	.930	.660		
Office IPD	0.009	0.059	0.00	0.313**
Office REIT	0.013	0.046	0.24	0.439**
F-test (p-value)	.530	.020		
Panel B: Three-year Returns				
Retail IPD	0.138	0.286	0.00	
Retail REIT	0.202	0.203	0.03	
F-test (p-value)	.070	.000		
Office IPD	0.134	0.260	0.00	
Office REIT	0.176	0.156	0.00	
F-test (p-value)	.180	.000		
Panel C: Five-year Returns				
Retail IPD	0.262	0.328	0.00	
Retail REIT	0.373	0.230	0.11	
F-test (p-value)	.010	.000		
Office IPD	0.268	0.269	0.00	
Office REIT	0.304	0.162	0.17	
F-test (p-value)	.290	.000		

Notes: Jarque-Bera denotes the Jarque-Bera test for normally distributed returns. Autocorrelation is not reported for the longer-run returns, since these returns are computed on an overlapping window basis. The TBI values are based on an 80 bps management cost assumption. The IPD volatility is based on a 0.6 desmoothing parameter.
*Significant at the 5% level.
**Significant at the 1% level.

accepted. Nevertheless, the F-statistics imply that the TBI volatility is greater than that of REITs at each horizon in the U.S. apartment sector. In the U.S. office sector, in turn, REIT market volatility is greater at the five-year horizon, whereas the observed quarterly TBI volatility is greater in the U.S. retail sector. The latter observation may well be due to measurement error induced noise in the TBI series (this is also supported by the VR statistics). The results further indicate that the private market volatility is greater than that of unlevered REITs in the U.K. office sector. Note also that the observed higher longer-term mean return for the REIT market than for the private market in the U.K. retail sector is to a large extent due to the loss of observations and substantial return volatility in the early sample period (i.e., this result is unreliable). In addition, the non-normal distribution of asset returns makes the p-values unreliable in many cases, especially at the quarterly frequency.

The baseline cointegration analysis results are summarized in Exhibit 4. Note that all the test statistics are small-sample corrected. Except for the U.K. office sector, the trace test statistics indicate cointegration between the public and private real estate total return indices. However, the U.S. apartment sector is a borderline case with a p-value of 0.08.[6] In each U.S. sector, REITs can be restricted to be weakly exogenous based on the LR test (i.e., only the TBI adjusts towards the cointegrating relation). This is in line with previous empirical evidence and the assumption that the direct market reacts more sluggishly than the REIT market to shocks. In contrast, in the U.K., it is the REIT market that adjusts towards the long-term relation in the retail sector (i.e., REIT returns can be predicted by deviations from the relation). This suggests that the REIT market is less mature and informationally efficient in the U.K. than in the U.S.

Most importantly, all the long-term coefficients on REITs, which are estimated based on the VECM presented in equation (2) without imposing any restrictions on the long-term coefficients (β), are close to one and, except for the U.S. retail sector, a one-to-one restriction on the long-term coefficients can be accepted. This indicates similarity in the public and private market long-term returns in the U.S. office, industrial, and apartment sectors and in the U.K. retail sector. The coefficient 1.10 on REITs in the U.S. retail sector implies slightly greater mean returns for direct real estate than for unlevered REITs. The point estimate indicates that, on average, when REIT returns are 10%, the corresponding TBI returns are 11%. The other "unconstrained" coefficients (i.e., the coefficients estimated on REITs without imposing the one-to-one restriction) vary from 0.89 in the U.K. retail sector to 1.05 in the U.S. industrial sector.[7]

Exhibit 5 shows the deviations of private market indices from the equilibrium relations. Except for U.S. retail, the deviations are those for the one-to-one relations. Generally, the greater the speed of adjustment parameter, the shorter and smaller the temporary deviations from the long-term equilibrium

Exhibit 4 | Baseline Cointegration Analysis Results

	Trace Test (p-value on r = 0)	LR Test, α(REIT) = 0 (p-value)	LR Test, α(TBI/IPD) = 0 (p-value)	LR Test, TBI = REIT (p-value)	Unconstrained Coeff. on REITs	α(TBI)	α(REIT)
Panel A: The U.S. Market							
Retail	0.00	0.77	0.00	0.03	1.100 (0.031)	−0.417 (0.070)	—
Office	0.00	0.62	0.00	0.65	0.941 (0.057)	−0.229 (0.051)	—
Industrial	0.00	0.13	0.00	0.26	1.050 (0.046)	−0.305 (0.062)	—
Apartments	0.08	0.43	0.00	0.57	0.910 (0.076)	−0.177 (0.063)	—
Panel B: The U.K. Market							
Retail	0.00	0.00	0.15	0.12	0.887 (0.039)	—	0.155 (0.047)
Office	0.62						

Notes: The direct market indices are based on the 80 bps management cost assumption. The U.S. models include one lag in differences and a point dummy variable for 2008:Q4. The U.K. models include two lags in differences and two point dummy variables, 1992:Q3 and 2008:Q4 for retail and 2007:Q4 and 2008:Q4 for office. The reported trace test statistics are Bartlett small-sample corrected and simulated to consider the influence of the dummy variables. The LR test on TBI/IPD = REIT is Bartlett small-sample corrected. The LR test on the one-to-one relation is a test on the joint hypothesis of REIT/direct weak exogeneity and the 1–1 relation in case the hypothesis of REIT/direct weak exogeneity is accepted. "Unconstrained" coefficient denotes the estimated coefficient in the case of no restrictions in the cointegrating vector, but restrictions in the alpha vector if accepted. The reported speed of adjustment parameters (α) are based on a 1–1 relation if such relation is not rejected. Standard deviations are in parentheses.

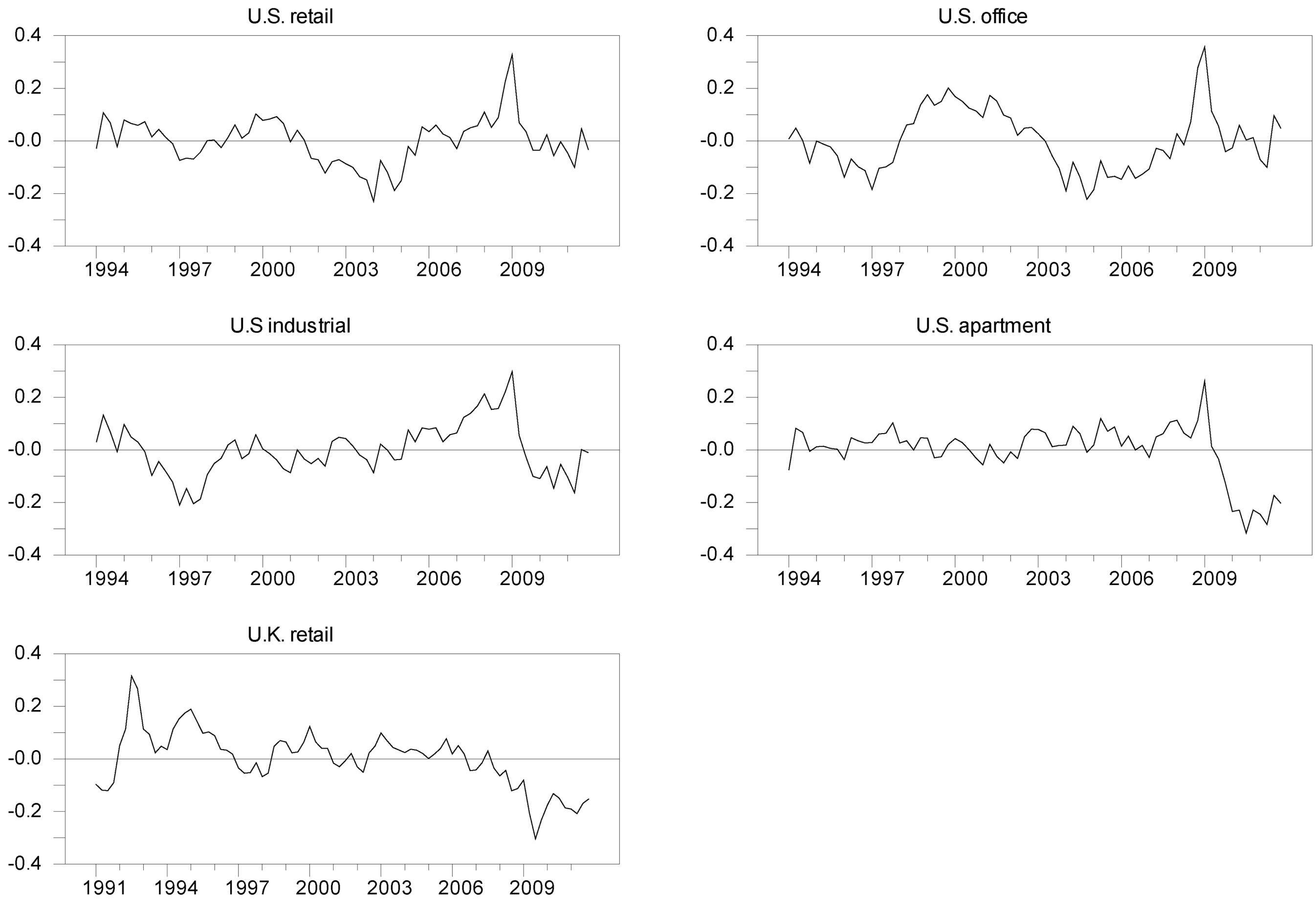
are likely to be, thereby enhancing the substitutability between the two markets and the possibilities to hedge direct real estate portfolio risks by public market vehicles.[8] In line with this argument, the U.S. office sector (speed of adjustment parameter of 23%) shows much longer-lasting deviations from the equilibrium than the more rapidly adjusting U.S. retail (42%) and industrial (31%) sectors during the late 1990s and early 2000s. The apartment sector results suggest that the inability to get stronger evidence of cointegration in the trace test as well as the small estimated alpha are due to the aftermaths of the GFC. While the other U.S. sectors were close to the long-term relations as of 2011, the apartment TBI remained approximately 20% undervalued relative to REITs. The apartment TBI followed closely the REIT index before the GFC, however. The deviation in the U.K. retail sector was −15% as of 2011:Q4.

Further Sensitivity Analyses

We discussed the impact of the investment horizon on the F-test statistics above. As the investment horizon for private real estate assets is typically several years, we further illustrate (in Exhibits 6 and 7) the influence of horizon on the return volatility by graphing the volatilities and their confidence bands in each market at each investment horizon up to 20 quarters based on VRs and their standard deviations. The U.K. direct market curves are based on the baseline 0.6 desmoothing parameter. Even though the wild bootstrap VR statistics have improved small-sample properties compared with the conventional VR statistics, the inference based on the computed VR statistics gets less reliable as the considered horizon becomes large compared with the overall sample period. Thus, the statistics at the longest horizons considered in Exhibits 6 and 7 should be interpreted cautiously. Despite this complication, we show the curves up to a five-year horizon, since as Campbell, Lo, and MacKinlay (1997) suggest, it may be only at the longer horizon (lower frequencies) that the impact of business cycles is detectable.

Exhibits 6 and 7 indicate short-term momentum in each of the return series. The annualized volatility

Exhibit 5 | Deviations of Private Market Indices from the Long-run Cointegrating Relations

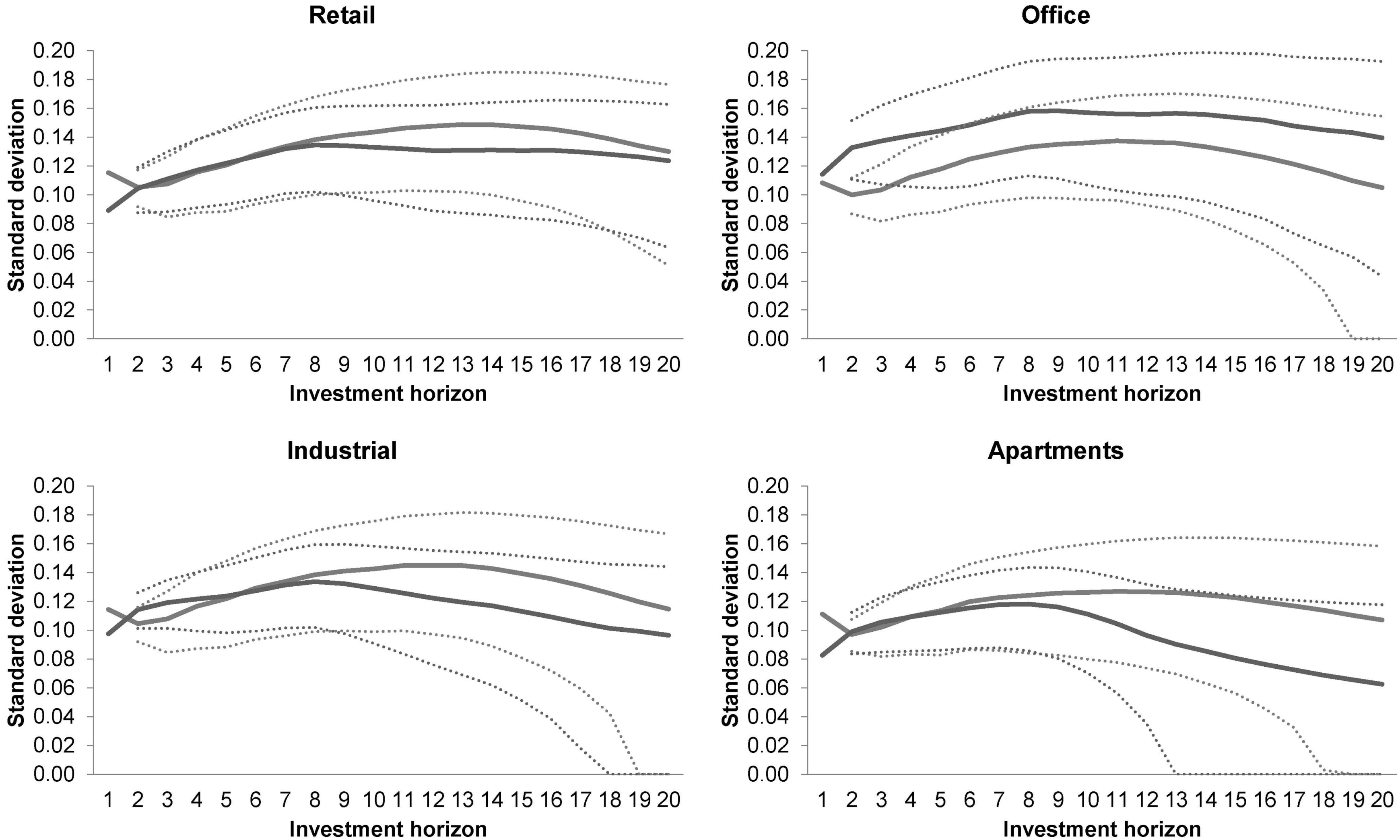


curves peak at around a two- to three-year horizon. Exceptions are the U.K. REIT returns, whose volatilities peak at the four-quarter (retail) and five-quarter (office) horizons. At longer horizons, the curves bend downwards, indicating longer-term mean reversion in the return series. The reversion is particularly strong in the U.S. apartment REIT market. The initial drop in the TBI volatility curves is likely due to the short-term measurement error in the indices, rather than because of actual negative short-term autocorrelation. This "noise" also contributes to the inability to find a statistically significant momentum effect for the U.S. direct market returns.[9]

While the investment horizon can affect both the absolute and relative riskiness of the markets, the main message of Exhibits 6 and 7 is that the shapes of the private and public real estate volatility curves are quite similar. Generally, REIT market volatility seems somewhat lower than private market volatility when leverage is catered for. Nevertheless, in each sector the hypothesis of equivalent standard deviations can be accepted regardless of the assumed investment horizon for the U.S. and at each horizon from three quarters onwards for the U.K.

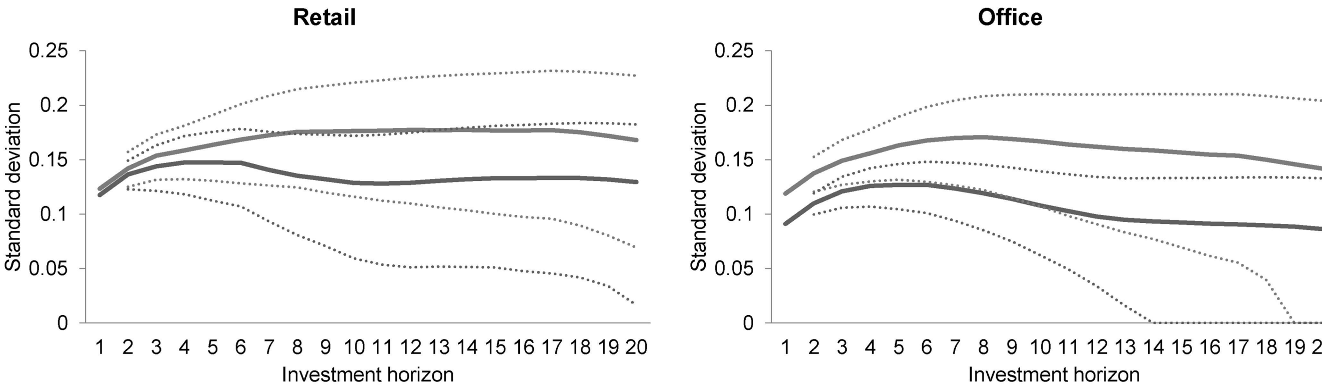
Note that regarding the U.K. office sector, the findings based on the more reliable VR statistics differ from the F-test results reported in Exhibit 3. The employed desmoothing parameter naturally affects the U.K. private market volatility curves. Nevertheless, the main message remains the same even if the parameter is 0.5 or 0.7 instead of the baseline 0.6. There are some slight changes, though: in the 0.5 case, there are no statistically significant volatility

Exhibit 6 | Annualized Standard Deviations of U.S. Private Real Estate and REIT Returns and Their Confidence Bands (±2 std. dev.)



Note: The dark gray line indicates REIT returns while the lighter gray line indicates private real estate returns.

Exhibit 7 | Annualized Standard Deviations of U.K. Private Real Estate and REIT Returns and Their Confidence Bands (±2 std. dev.)



Note: The dark gray line indicates REIT returns while the lighter gray line indicates private real estate returns.

differences even at the short horizon, and in the 0.7 case, the IPD volatility seems to be notably greater than that of REITs, although still not in a statistically significant manner.

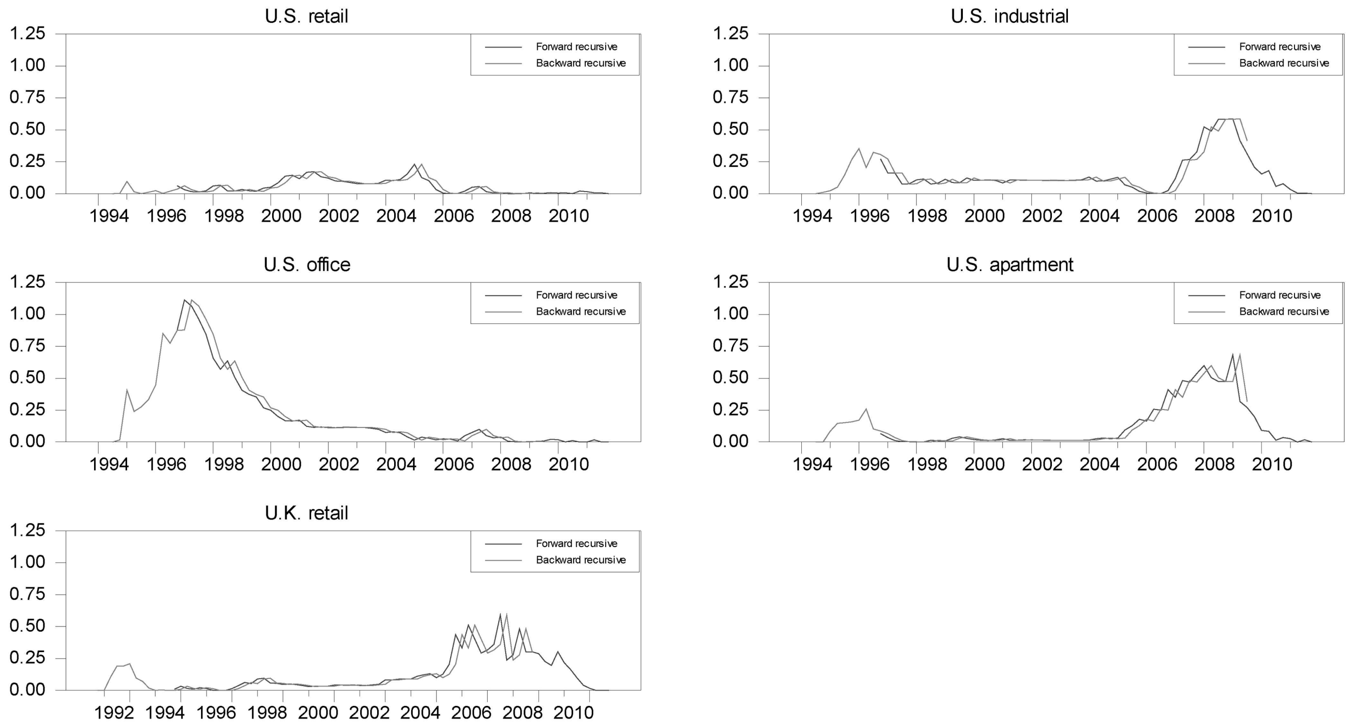
Next, we investigate whether the estimated long-run relations, which are one-to-one relations in four out of five sectors, are stable over the sample period. In Exhibit 8, we graph the recursive and backwards recursive max test statistics for the hypothesis of constancy of the estimated long-run relation. The statistics are scaled by the 5% critical values, so that a value exceeding one indicates rejection of the null. The stability is clearly accepted in the U.S. retail sector throughout the sample period. In the U.S. office sector, there is evidence of instability during the early sample period, but the relation remains constant thereafter. The U.S. apartment and industrial sectors as well as the U.K. retail sector, in turn, show temporary instability during the GFC, which is not unexpected given a visual inspection of the deviations from the relations (Exhibit 5). However, there is no evidence of a permanent structural change and the instability around the GFC is not statistically significant. Therefore, it is reasonable to believe that the estimated relations hold at the end of the sample period.[10]

We also test the robustness of our findings with respect to the assumed direct real estate portfolio management fees. These fees may range between 50 and 120 bps per year (Riddiough, Moriarty, and Yeatman, 2005), and in the baseline analysis we follow those authors by using the 80 bps points assumption. Exhibit 9 presents cointegration test results for the 50 bps and 120 bps assumptions.

While the trace test results are generally in line with the baseline analysis, there are some changes.[11] In the U.S. apartment sector, the conclusion on cointegration is dependent on the management fee assumption: the hypothesis of no cointegration is rejected at the 4% significance level in the 50 bps case, but only at the 11% level in the 120 bps case. Assuming cointegration, the hypothesis of a one-to-one relation can be accepted in the apartment sector for all management fee assumptions. Regarding the one-to-one hypothesis, the management fee assumption is of significance in the retail sector, as the hypothesis is accepted assuming 120 bps fees, but not otherwise. In the industrial sector, in turn, the one-to-one relation is a borderline case if the true management fees are 50 bps per year.

Generally, the speed of adjustment parameters are the greatest in the baseline case. This suggests that the interdependence between the REIT and TBI/IPD indices is the tightest when the 80 bps level of management costs is assumed. This is indirect support for the relevance of the 80 bps assumption and is sensible given that 80 bps is considered to be a typical value for the management costs, while the 50

Exhibit 8 | Forward Recursive and Backward Recursive Max Test Statistics (in the R-form) of Constancy of the Estimated Long-run Relation Scaled by the 5% Critical Value



bps and 120 bps cases are deemed the lowest and highest possible. In other words, the baseline analysis generally can be considered to yield the most reliable conclusions. Again, an exception is the U.S. apartment sector, where the linkage between the markets is strongest in the 50 bps case. Also, in the U.K. retail sector, the REIT speed of adjustment parameter is the greatest when 50 bps management fees are assumed. These findings suggest that the management fees in these sectors are actually close to 50 bps.

The Influence of the Global Financial Crisis

As shown by Exhibit 8, the stability of the long-run relations cannot be rejected even during the GFC. Nevertheless, Exhibit 5 indicates that substantial deviations from the long-term relations emerged after the Lehman Brothers collapse. Therefore, it is worthwhile to briefly discuss the influence of the GFC on the relation between public and private real estate market performance.

Expectedly, the outbreak of the GFC had a notable adverse influence on asset prices both in public and private markets (Exhibit 1). Since the REIT market reacted to the adverse shock much earlier than the direct market, deviations from the long-run relations of around 30% in each U.S. sector took place in 2008:Q4–2009:Q1. In contrast, such overpricing of private real estate relative to REITs did not take place in the U.K. retail sector: also REITs reacted sluggishly in the U.K. market.

Given the total return indices shown in Exhibit 1 and the previous empirical evidence on the more sluggish adjustment of the direct market than the REIT market to shocks in the fundamentals, the large initial deviations in the U.S. after the Lehman collapse were most probably due to the substantial private real estate market frictions. These frictions include low liquidity: when an investor needs cash rapidly, due to an inability to refinance short-term debt for instance, the investor will typically sell the more liquid publicly traded assets first as those can

Exhibit 9 | Cointegration Analysis Results Assuming 50 bps and 120 bps Management Fees in the Private Market

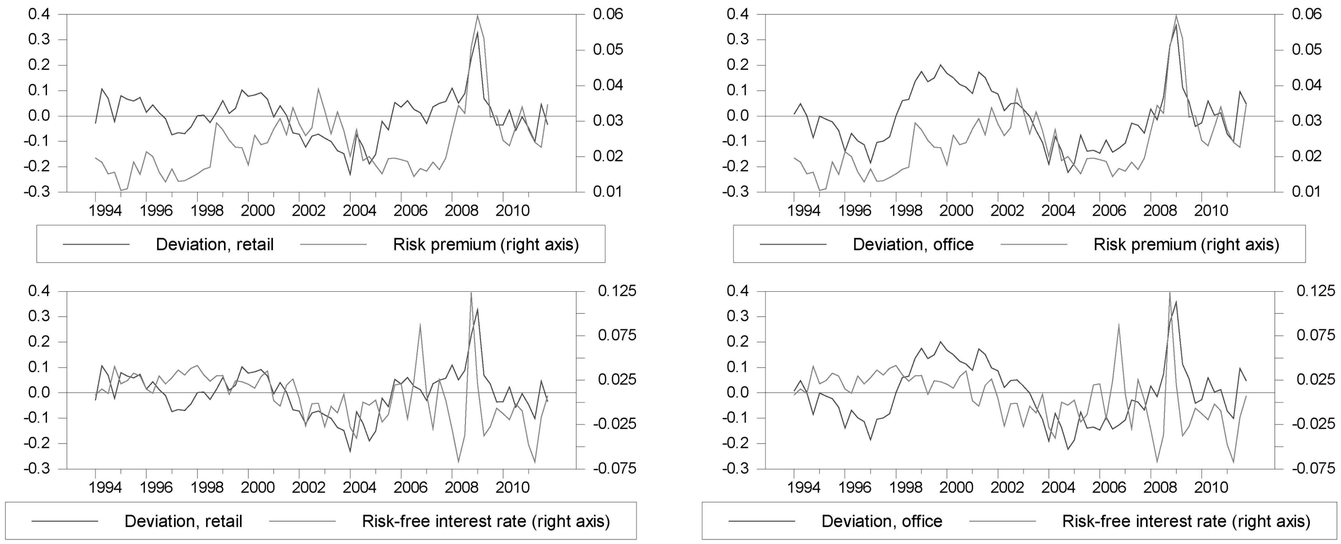
	Trace Test (p-value on r = 0)	LR Test, α(REIT) = 0 (p-value)	LR Test, α(TBI/IPD) = 0 (p-value)	LR Test, TBI = REIT (p-value)	Unconstrained Coeff. on REITs	α(TBI)	α(REIT)
Panel A: The U.S. Market							
50 Basis Points Management Fees							
Retail	0.00	0.65	0.00	0.00	1.140 (0.032)	−0.407 (0.068)	—
Office	0.00	0.54	0.00	0.73	0.954 (0.066)	−0.212 (0.047)	—
Industrial	0.00	0.10	0.00	0.09	1.090 (0.044)	−0.286 (0.060)	—
Apartments	0.04	0.33	0.00	0.57	0.951 (0.072)	−0.205 (0.066)	—
120 Basis Points Management Fees							
Retail	0.00	0.14	0.00	0.27	1.040 (0.035)	−0.390 (0.064)	—
Office	0.00	0.45	0.00	0.22	0.872 (0.061)	−0.194 (0.047)	—
Industrial	0.00	0.18	0.00	0.52	1.000 (0.048)	−0.303 (0.063)	—
Apartments	0.11	0.58	0.01	0.48	0.856 (0.082)	−0.138 (0.058)	—
Panel B: The U.K. Market							
50 Basis Points Management Fees							
Retail	0.00	0.00	0.17	0.19	0.920 (0.036)	—	0.250 (0.059)
Office	0.57						
120 Basis Points Management Fees							
Retail	0.00	0.01	0.11	0.09	0.842 (0.044)	—	0.209 (0.057)
Office	0.65						

Notes: The U.S. models include one lag in differences and a point dummy variable for 2008:Q4. The U.K. models include two lags in differences and two point dummy variables, 1992:Q3 and 2008:Q4 for retail and 2007:Q4 and 2008:Q4 for office. The reported trace test statistics that are Bartlett small-sample corrected and simulated to consider the influence of the dummy variables. The LR test on TBI/IPD = REIT is Bartlett small-sample corrected. The LR test on the one-to-one relation is a test on the joint hypothesis of REIT/direct weak exogeneity and the 1–1 relation in case the hypothesis of REIT/direct weak exogeneity is accepted. "Unconstrained" coefficient denotes the estimated coefficient in the case of no restrictions in the cointegrating vector, but restrictions in the alpha vector if accepted. The reported speed of adjustment parameters (α) are based on a 1–1 relation if such relation is not rejected. Standard deviations are in parentheses.

be sold relatively fast without having to accept as large a discount as with the less liquid privately traded assets (Brunnermeier, 2009). It should also be noted that part of the initial adjustment in the direct real estate market took place through lower liquidity (i.e., longer time-on-the-market and fewer transactions). This kind of adjustment is not visible from the total return series. Based on Kim and Lee (2014), the notable decrease in private real estate market liquidity during crisis periods such as the GFC is expected to increase the required return on private real estate assets. Nevertheless, our findings do not suggest that private market returns are generally greater than those of REITs.

The overpricing of TBI relative to REITs disappeared towards the end of the sample period, as the direct market gradually adjusted and the financial markets became more stable. Hence, the large deviations predicted the forthcoming collapse in direct market values. There was even some overshooting in the other direction (i.e., towards an undervaluation of direct real estate) in the industrial and apartment sectors since late 2009. This is partially due to the

Exhibit 10 | Deviation of Retail and Office TBI from the Long-run Relation, the Risk Premium, and the Risk-free Interest Rate



rapid "rebound" of the REIT market. While this deviation has vanished in the industrial sector, direct apartment investments remained about 20% undervalued relative to their public counterpart as of 2011:Q4. REITs rebounded faster in the U.K. market as well, inducing notable undervaluation of the direct retail market compared with REITs. The magnitude of the undervaluation was greatest (30%) in 2009:Q3. Relying on the reported statistics indicating that there have not been permanent structural changes in the long-term relations, the findings suggest that the U.S. private apartment assets and U.K. private retail real estate assets were expected to appreciate notably faster than their REIT counterparts in the few years after the end of our study period.

An interesting question is whether there could be a particular underlying fundamental variable that can be employed to determine the large temporary deviations. Hoesli and Oikarinen (2012) and Hoesli, Oikarinen, and Serrano (2015) show that REITs tend to react substantially faster to risk premium and real interest rate shocks than do direct real estate values, and that a notable increase in the interest rate preceded, while an increase in the risk premium coincided with, the emergence of the substantial deviations. These findings emphasize the role of adverse interest rate and risk premium shocks behind the deviation patterns during the GFC. Exhibit 10 illustrates the relation between the real risk-free interest rate, risk premium, and the deviations from the equilibrium relations for the U.S. market. The interest rate and risk premium are measured here as the three-month Treasury bill rate and the spread between corporate bond (Baa, Moody's) yield, and the 10-year government bond yield, respectively.

Regarding portfolio allocation implications, the observations suggest that an investor should not reallocate his portfolio from REITs to direct real estate after a drastic drop in REIT prices due to shocks such as the Lehman Brothers collapse. This is because the direct market is likely to follow the REIT market fall, and the expected returns for REITs are therefore greater than those for direct real estate for some time after such an adverse shock.

Finally, the experience from the GFC period suggests that hedging private real estate exposure by public real estate derivatives can work during a crisis period (i.e., when such hedging is needed the most). Despite the slower response of private real estate values, the linkage between the private and public markets remained generally quite constant during the GFC and its aftermath, as the notable long-run deviations vanished relatively quickly (with the exception of the U.S. apartment sector and the U.K.

retail sector to a lesser extent). Importantly, as it is typically hard to sell the more illiquid private real estate assets rapidly during a crisis without a notable discount, the gains on the derivatives used to hedge the downside risks can be used as a source of necessary liquidity instead of having to conduct distressed sales of private assets at a discount. That is, the use of such derivatives also could diminish, at least to some extent, the illiquidity risk of privately traded assets, such as direct real estate investments.

Conclusion

We use sector level real estate data for the U.S. and U.K. to investigate the similarity of public and private market returns and risks. The data are adjusted for the effects of leverage and management fees. The results provide evidence of tight long-term relations (cointegration) between the public and private markets in the four U.S. property types included in the analysis and in one of the two U.K. property sectors. Thus, while in the short run the observed comovement between REITs and direct real estate markets can be low due to factors such as data complications and market frictions in the private market, in the long term public and private real estate returns are tightly linked. In four of the five cointegrated sectors, the hypothesis of a one-to-one long-term relation between the adjusted total return indices is clearly accepted. We also find that volatilities generally do not differ significantly between REITs and direct real estate regardless of sector and investment horizon.

Our findings have important practical implications. First, public and private real estate investments can generally be considered to work as close substitutes in an investment portfolio with several years' investment horizon, since they provide similar total returns and return variances, and co-move tightly over the long horizon. This substitutability is somewhat weakened by the difference in the liquidity between public and private real estate assets. On the other hand, the good "funding liquidity" (Brunnermeier, 2009) of private real estate enhances the substitutability as a counter-force for the low "market liquidity." Anyhow, as securitized real estate assets enable diversification with smaller amounts of capital, and the liquidity is better and transaction costs are lower in the public market than in the private market, investors who have limited amounts of capital and highly value liquidity and low transaction costs should tilt their real estate holdings towards REITs. As would be expected, liquidity and transaction costs become less important as the planned investment horizon is increased, making private assets relatively more attractive for investors with long holding periods.

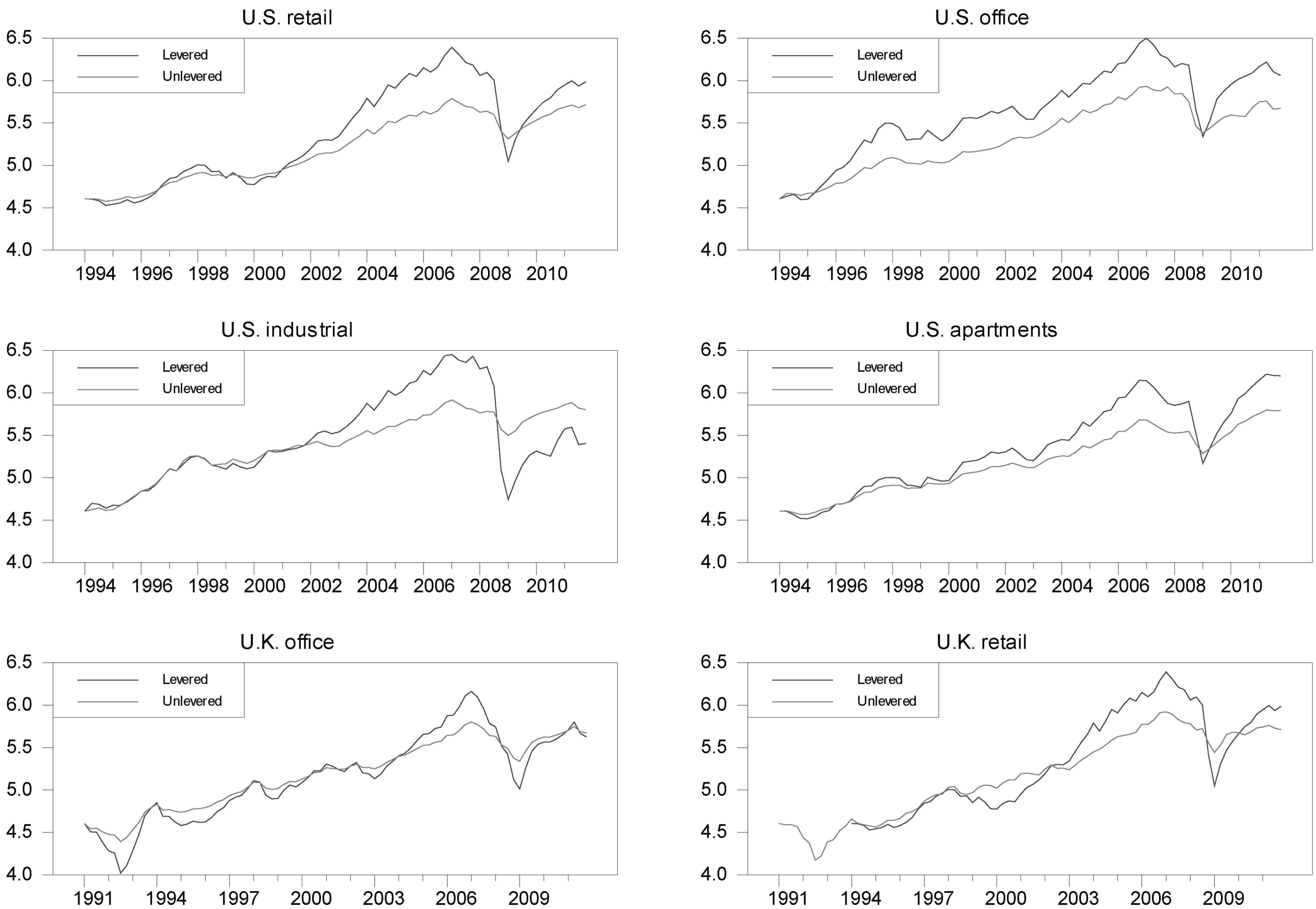
Second, the tight long-term relation between public and private returns suggests that REIT-related ETFs and derivatives can be used to hedge risks created by direct real estate holdings. Short positions on ETFs, for instance, offer a good opportunity to hedge risks in lending institutions' portfolios that arise due to their outstanding mortgage lending inventory. Among other potential benefits, such hedging could help banks to survive better through periods of economic distress and drastically decreasing real estate prices. From an investor's point of view, during crisis periods the gains on the derivatives used to hedge the downside risks could be used as a source of necessary liquidity rather than to have to sell private assets at a substantial discount, thus diminishing the illiquidity risk of private assets.

Due to the potentially lengthy deviations from the equilibrium relations between public and private real estate and the idiosyncratic risk of individual properties, hedging cannot totally remove risks. Moreover, in many markets the current public market related vehicles are not sufficient to properly exploit the hedging opportunities. That is, new financial vehicles, especially for taking longer-term short positions, and more liquid markets for them are needed in order to be able to take better advantage of hedging potentials. Generally, the longer the horizon and the faster the adjustment of the private market towards the equilibrium relation, the better are the hedging opportunities.

An empirical examination into the correspondence of public and private asset returns and return volatilities is usually not possible, since there are no reliable time series data on the typical underlying privately traded assets. The “duality” and data availability of the real estate market offers a laboratory to such empirical examinations. As such, this analysis has implications on the relation between publicly and privately traded assets in the broader investment universe as well.

Appendix

Exhibit A1 | Unlevered and Levered REIT Total Return Indices in Real Terms



Endnotes

1. In this article, “public” and ‘”securitized” investments are used as synonyms, both referring to securities traded in public market places. “Private” and “direct” investments, in turn, both refer to the underlying privately traded assets.
2. For the real estate sector, the influence of managers’ capability on fund performance and the economies of scale provided by large REITs are sometimes given as potential factors inducing higher returns for REITs than for private real estate. The assertion that these factors could lead to greater expected security returns is at odds with the efficient market hypothesis: any expected managers’ or economies of scale influence on security cash flows should be reflected in the prevailing asset prices, while the expected return-risk relations should not be increased. Indeed, if managers’ capability lessens the risks, the expected (required) return should accordingly be smaller.
3. The classification of companies by property type as of 2006 was used to construct the sector indices for the period from 1991 to 2005.
4. Devaney and Diaz (2011) use a hedonic model to construct transaction-based indices for the period 2002–2010; the period unfortunately is too short for us to use their indices directly. We use their results concerning selection corrected indices, however, as our benchmark for the level of serial correlation that should be inherent to real estate indices at the quarterly frequency.
5. There also are techniques, such as fully-modified OLS (FMOLS) or dynamic OLS (DOLS) that allow for the estimation of single equation cointegrating models with a preset dependent variable. However, there are several reasons to

use the Johansen method. First, while the aforementioned techniques assume one dependent variable that adjusts towards the long-term relation, the Johansen method allows both private and public markets to adjust. This is of importance, since in theory and based on extant empirical literature either market can react to deviations from the long-term cointegrating relation. Second, the Johansen approach allows for a formal test on weak exogeneity of the variables. Third, it also allows for a formal recursive investigation of the stability of the long-term relation(s). Moreover, the Johansen technique avoids the two-step complication present in the residual-based single-equation cointegration tests such as FMOLS and DOLS, and takes into consideration the short-term dynamics of the system. Finally, small-sample corrected test values are available for the trace test and for the likelihood ratio test for model restrictions, to increase the efficiency of the tests.

6. The hypothesis of no cointegration can clearly be rejected in the U.S. apartment sector if REITs are readily set to be weakly exogenous in the analysis to enhance the power of the trace test [for a discussion of the trace test in the presence of weakly exogenous variables, see Harbo, Johansen, Nielsen, and Rahbek (1998)]. This applies to the other U.S. sectors as well.
7. We also checked whether the inclusion of the point dummies notably alters the results. The influence of the point dummies generally is cosmetic. An exception is the U.K. retail sector where the one-to-one restriction is a borderline case with respect to statistical significance if the financial crisis point dummy is removed.
8. Note that the hedging potential does not necessitate a one-to-one relation between the returns: any cointegrating relation that is stable over time offers hedging opportunities. The greater the coefficient on REITs, the greater is the required exposure to REIT derivatives to hedge the direct real estate exposure.
9. Some individual autocorrelations for TBI returns are significantly greater than zero. The hypothesis of no significant momentum in the direct market is clearly rejected in each sector if the baseline is the two-quarter horizon instead of the one-quarter horizon.
10. The recursive analysis conclusions are unaffected by the inclusion or exclusion of the point dummies.
11. The max test statistics regarding the stability of the cointegrating relations are not notably affected by the variation in the management fees.

References

Andonov, A., P. Eichholtz, and N. Kok. Intermediated Investment Management in Private Markets: Evidence from Pension Fund Investments in Real Estate. *Journal of Financial Markets*, 2015, 22, 73–103.

Boudry, W.I., N.E. Coulson, J.G. Kallberg, and C.H. Liu. On the Hybrid Nature of REITs. *Journal of Real Estate Finance and Economics*, 2012, 44, 230–49.

Brunnermeier, M. Deciphering the Liquidity and Credit Crunch 2007–2008. *Journal of Economic Perspectives*, 2009, 23, 77–100.

Campbell, J.Y., A.W. Lo, and A.C. MacKinlay. *The Econometrics of Financial Markets*. Princeton, NJ: Princeton University Press, 1997.

Campbell, J.Y. and L.M. Viceira. The Term Structure of the Risk: Return Trade-Off. *Financial Analysts Journal*, 2005, 61, 34–44.

Clayton, J., D. Geltner, and S.W. Hamilton. Smoothing in Commercial Property Valuations: Evidence from Individual Appraisals. *Real Estate Economics*, 2001, 29, 337–60.

Clayton, J. and G. MacKinnon. The Time-Varying Nature of the Link between REIT, Real Estate and Financial Asset Returns. *Journal of Real Estate Portfolio Management*, 2001, 7, 43–54.

——. The Relative Importance of Stock, Bond and Real Estate Factors in Explaining REIT Returns. *Journal of Real Estate Finance and Economics*, 2003, 27, 39–60.

Dennis, J.G. *CATS in RATS. Cointegration Analysis of Time Series, Version 2*. Evanston, IL: Estima, 2006.

Devaney, S. and R.M. Diaz. Transaction Based Indices for the U.K. Commercial Real Estate Market: An Exploration Using IPD Transaction Data. *Journal of Property Research*, 2011, 28, 269–89.

Fisher, J., D. Gatzlaff, D. Geltner, and D. Haurin. Controlling for the Impact of Variable Liquidity in Commercial Real Estate Price Indices. *Real Estate Economics*, 2003, 31, 269–303.

Geltner, D. and B. Kluger. REIT-Based Pure-Play Portfolios: The Case of Property Types. *Real Estate Economics*, 1998, 26, 581–612.

Harbo, I., S. Johansen, B. Nielsen, and A. Rahbek. Asymptotic Inference on Cointegrating Rank in Partial Systems. *Journal of Business & Economic Statistics*, 1998, 16, 388–99.

Hendry, D.F. and K. Juselius. Explaining Cointegration Analysis: Part 1. *The Energy Journal*, 2000, 21, 1–42.

Hoesli, M. and E. Oikarinen. Are REITs Real Estate? Evidence from International Sector Level Data. *Journal of International Money and Finance*, 2012, 31, 1823–50.

Hoesli, M., E. Oikarinen, and C. Serrano. Do Public Real Estate Returns Really Lead Private Returns? *Journal of Portfolio Management*, 2015, 41, 105–17.

Hudson-Wilson, S., F.J. Fabozzi, and J.N. Gordon. Why Real Estate? *Journal of Portfolio Management*, 2003, 29, 12–25.

Johansen, S. *Likelihood-Based Inference in Cointegrated Vector Auto-Regressive Models*. Oxford: Oxford University Press, 1996.

——. A Bartlett Correction Factor for Tests on the Cointegrating Relations. *Econometric Theory*, 2000, 16, 740–78.

——, S. A Small Sample Correction for the Test of Cointegrating Rank in the Vector Autoregressive Model. *Econometrica*, 2002, 70, 1929–61.

Johansen, S., R. Mosconi, and B. Nielsen. Cointegration Analysis in the Presence of Structural Breaks in the Deterministic Trend. *Econometrics Journal*, 2000, 3, 216–49.

Juselius, K. *The Cointegrated VAR Model: Methodology and Applications*. New York, NY: Oxford University Press, 2006.

Kim, J.H. Wild Bootstrapping Variance Ratio Tests. *Economics Letters*, 2006, 92, 38–43.

Kim, S. and K. Lee. Pricing of Liquidity Risks: Evidence from Multiple Liquidity Measures. *Journal of Empirical Finance*, 2014, 25, 112–33.

Kraft, H. and C. Munk. Optimal Housing, Consumption, and Investment Decisions over the Life Cycle. *Management Science*, 2011, 57, 1025–41.

Ling, D.C. and A. Naranjo. Returns and Information Transmission Dynamics in Public and Private Real Estate Markets. *Real Estate Economics*, 2015, 43, 163–208.

MacKinnon, G.H. and A. Al Zaman. Real Estate for the Long Term: The Effect of Return Predictability on Long-Horizon Allocations. *Real Estate Economics*, 2009, 37, 117–53.

Modigliani, F. and M.H. Miller. The Cost of Capital, Corporation Finance and the Theory of Investment. *The American Economic Review*, 1958, 48, 261–97.

Oikarinen, E., M. Hoesli, and C. Serrano. The Long-Run Dynamics between Direct and Securitized Real Estate. *Journal of Real Estate Research*, 2011, 33, 73–103.

Pagliari Jr., J.L., K.A. Scherer, and R.T. Monopoli. Public Versus Private Real Estate Equities: A More Refined, Long-Term Comparison. *Real Estate Economics*, 2005, 33, 147–87.

Riddiough, T.J., M. Moriarty, and P.J. Yeatman. Privately Versus Publicly Held Asset Investment Performance. *Real Estate Economics*, 2005, 33, 121–46.

Yavas, A. and Y. Yildirim. Price Discovery in Real Estate Markets: A Dynamic Analysis. *Journal of Real Estate Finance and Economics*, 2011, 42, 1–29.

Yunus, N., J.A. Hansz, and P.J. Kennedy. Dynamic Interactions between Private and Public Real Estate Markets: Some International Evidence. *Journal of Real Estate Finance and Economics*, 2012, 45, 1021–40.

Martin Hoesli, University of Geneva, Geneva, Switzerland; University of Aberdeen Business School, Scotland, and Kedge Business School, France or martin.hoesli@unige.ch.

Elias Oikarinen, University of Turku, Turku, Finland or elias.oikarinen@utu.fi.