# Long-Term Structural Price Relationships in Real Estate Markets

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***Abstract.* This study investigates the long-run stochastic properties of real estate assets by geographical breakdown. We also study their linkages with financial assets. The initial tests find that almost all property types exhibit the presence of nonstationarity. Thus, cointegrated methodologies are used. Structural breakpoints identified in the literature are used as a guide to divide the data into two windows, 1983–1989 and 1990–1996. The results show that real estate in the different regions exhibit a closer relationship with each other in the second period, compared with the first. Also, strong linkages between real estate regions and financial assets are noted in the second period. The South is the only region to exhibit segmentation in both periods. Overall, the information derived from our analysis sheds light on linkages among real estate assets and between real estate and financial assets and also provides a framework for creating diversified portfolios.**

## Introduction

Studies that examine linkages between different types of real estate assets and between real estate and financial assets are relatively sparse. This is surprising, given the value of real estate, the cost of transacting real estate and its relative illiquidity compared with financial assets. Knowing the nature of such linkages will provide useful information to many investors. We know, for instance, that real estate professionals frequently seek suitable hedging vehicles to reduce their exposure to real estate, while investors, including institutions, evaluate the usefulness of real estate as investment and diversification tools. Institutional involvement in commercial real estate increased to approximately $232 billion by mid-1995 (Grissom and De Lisle, 1997). Fisher and Webb (1992) note that for companies in the United States, ownership of real estate is an important allocation of their capital.[1] Some also point out that real estate is becoming increasingly fungible prompting investors to move capital between real estate and other assets (McMahan, 1997). In addition, Webb, Miles and Guilkey (1992) identify significant anticipated and unanticipated inflation-hedging properties of real estate. Thus, identifying the price dynamic in real estate assets, and linkages between real estate and financial assets provides an idea of the opportunities real estate offers to investors.

To examine these linkages the nature of the underlying data used need to be accounted for. The statistical problems encountered with real estate data are well known.

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Evidence of autocorrelation and nonnormality in real estate data has been found (Myer and Webb, 1994). Others find evidence of nonstationarity in real estate data (Chaudhry, Myer and Webb, 1999). Young and Graff (1995) suggest that investment risk models with infinite variance provide a better description of distributions of individual property returns. Consequently, standard statistical procedures, especially in the presence of nonstationarity, may give misleading inferences. However, advances in cointegration methodologies provide an alternative framework for investigating equilibrium price adjustments in financial and real estate time series, especially their long-run relationships. Webb, Miles and Guilkey (1992) suggest that appraisal-based estimates may be unbiased in the long run. Cointegration methods can be applied to appraisal-based data to study long-term relationships among real estate assets while accounting for idiosyncrasies in the data. The resulting estimates should provide a rich source of information to investors, particularly institutions, given the increasing involvement of institutions in real estate markets. Grissom and De Lisle (1997) argue that institutional ownership in real estate is relatively long term and that institutional ownership groups require analysis of more than an individual property. Thus, these methods seem especially suitable.

Nonstationarity characterizes many different types of time series and is not a problem confined to real estate data. For instance, nonstationarity has been detected in financial time series warranting the use of cointegration methods to study linkages in such data. Bradley and Lumpkin (1992) use cointegration methodologies to examine relationships in Treasury debt markets. They find strong evidence of cointegration in seven Treasury rates, indicating that a long-run relationship exists among these rates. Their results indicate that a forecast that uses information about this long-term relationship outperforms one that does not. The use of cointegration methods in real estate studies although less frequent, is becoming increasingly popular (*e.g.*, Abeybayehu and Kuchler, 1993; Chaudhry, Myer and Webb, 1999; and Rudolph and Griffith, 1997). For instance, Abeybayehu and Kuchler use cointegration methods to identify the determinants of farmland prices. They find little evidence to reject the hypothesis that market fundamentals determine farmland prices.

It is reasonable to presume that linkages between markets and instruments arise due to common macroeconomic fundamentals. Assets that have common stochastic trends are said to be cointegrated. For instance, many have shown that different instruments respond to the release of macroeconomic news by governmental agencies (Harvey and Huang, 1991; and Ederington and Lee, 1993). While common macroeconomic factors may affect all assets, other factors may affect only a group of assets. Malliaris and Urrutia (1996) note that substitutability and complementarity create price interdependence or linkages in agricultural commodities. Thus, it is quite possible that the primary determinants of the price of real estate assets may differ from that of financial assets. These differences and linkages present important implications for fund managers and real estate investors pursuing risk management. Using cointegration methodologies, we study the extent of these differences and linkages between different categories of real estate assets and between real estate assets and financial assets such as Treasury bonds.

The results from this study make several contributions to the literature. First, the time series properties of several classes of real estate and financial assets are detailed. Real estate assets are examined by geographical classification. Twelve real estate categories are studied. The geographical sub-categorization of real estate is important as regional and local variations are important determinants of the price of real estate. For instance, Smith and Tesarek (1991) suggest that the economy-wide energy-induced recession in the 1970s and early 1980s may have not caused disequilibrium in the Houston housing market, as much as the excessive construction activity that took place in the 1980s. The increase in supply far outstripped the demand of buyers. The study also examines important financial assets such as the S&P 500 Index and the Lehman Govt/ Corp Index (Treasury Bonds) and determines their linkages with real estate assets. Thus, this study extends the work of Chaudhry, Myer and Webb (1999). They examine such linkages using a broader classification of real estate asset types and detail linkages in real estate assets by type, such as office and retail, and by region, such as East and Midwest. Building on this, we look at geographical subdivisions within this classification, for instance East North Central and Northeast. The results from this line of research will enable investors to structure manageable portfolios. Second, we establish whether different groups or subgroups of real estate are segmented or integrated. Investors seeking risk reduction should consider diversifying their investments into segmented markets. If markets are integrated, portfolio risk reduction may be less, depending on the degree of integration. Third, using the Johansen Test (Johansen, 1988, 1991, 1994), we detail the number of common stochastic trends, or cointegrating vectors, for different real estate assets and for different combinations of real estate and financial assets, enabling estimation of the number of common linkages. In general, the higher the number of linkages the better a hedge is expected to perform.

Others have also studied linkages involving real estate assets. To illustrate, Ling and Naranjo (1997), identify fundamental macroeconomic variables that systematically affect real estate returns. Lizieri and Satchell (1997) find that a wider economy leads the real estate market in the short term but that, with a longer lag structure, positive real estate returns may point to negative future returns with the rest of the economy. They use a two-sector analytic model to capture these relationships. Cheung, Tsang and Mak (1995) examine the linkage between sale price changes and rental rate changes in the Hong Kong real estate market for the period 1982–1991 and conclude that the two markets are efficient. Causal relationships are not found in twenty-nine of the forty cases examined. Darrat and Glascock (1993) use a vector autoregressive process to study the relationship between real estate and financial variables. They find evidence of a significant lagged relationship between real estate returns and fiscal policy moves. We add to this growing line of research by detailing linkages among real estate assets and between real estate assets and financial assets.

The remainder of this article is organized as follows. First, we present the methodology used to test for stationarity and integration/segmentation and describe the data. Next, we detail the results of the tests. We then present a summary and a conclusion.[2]

## Research Design and Data

Most time series used in real estate studies are nonstationary, and the use of cointegrated methodologies accounts for this characteristic.[3] Engle and Granger (1987) suggest that if a system of variables is cointegrated, then economic forces interact to bind these variables together in a long-run equilibrium relationship. In this event, an error correction model (ECM) can represent the cointegrated variables.[4] In general, the ECM shows the dependence of this periods price change on last periods price change, thus providing a measure of how far the system is out of long-term equilibrium. Before testing for cointegration between two or more series, it is necessary to test whether the different time series are integrated to the same order.[5] To do this, we apply the conventional unit roots tests described below.

***Stationarity*** *(Unit Root) Tests for Individual Time Series*

In general, most texts and manuals on tests for stationarity of a time series ($TS_t$) will probably begin with the estimation of the following regression equation, if no linear trend is considered:[6]

$$\Delta\, TS_t = \alpha_0 + \alpha_1 TS_{t-1} + \sum_{j=1}^{p} \gamma_j\, \Delta TS_{t-j} + \varepsilon_t, \tag{1}$$

and by Equation (2) when a linear trend and a parameter for drift are considered:

$$\Delta TS_t = \alpha_0 + \alpha_1 TS_{t-1} + \alpha_2 t + \sum_{j=1}^{p} \gamma_j\, \Delta TS_{t-j} + \varepsilon_t, \tag{2}$$

where the $\Delta$ represents differences (first differences unless otherwise noted), $\alpha_0$ represents the term for drift in the series, $\alpha_1$ allows testing for a unit root and $\alpha_2$ verifies the presence of trend. The error-correcting mechanism is represented by $\Delta TS_{t-j}$ in the model. If the hypothesis $\alpha_1 = 0$ cannot be rejected, then the series is said to have a unit root and is nonstationary and conversely, if the hypothesis, $\alpha_1 = 0$, is rejected, then it is concluded that the series does not contain a unit root and is stationary. Tests involving parameters $\alpha_0$ and $\alpha_2$ verify the presence of drift and trend. Inclusion of the $p$-lagged values ensures a white-noise series. The number of lags is determined by a test of significance such as the Akaike Information Criterion (AIC) (Akaike, 1973).[7] The distribution of the ordinary $t$- and $F$-Statistics computed for the regressions do not have the expected distributions. Thus, to test the various hypotheses critical values have been computed using Monte Carlo techniques and are tabulated in various references (*e.g.*, Davidson and MacKinnon, 1993).

We apply robust methodologies in our testing procedures. Tests for stationarity and cointegration use the Philips-Perron (P&P) non-parametric testing procedure. The P&P procedure is used as the crucial i.i.d. error assumption is not needed.[8]

***Integration/Segmentation Tests for Groups***

Once the unit root tests are completed, we proceed to test for cointegration for different groups. If two or more series are cointegrated, we can infer that the series are stationary. This part of the analysis uses the methodology developed by Johansen (1988). The description that follows draws heavily from Johansen (1988, 1991, 1994) and Johansen and Juselius (1990, 1991). The Johansen methodology presents some distinct advantages. For example, identification of the number of cointegrating vectors is possible with the Johansen test. Such inferences are based on the number of significant eigenvalues. Also, many argue that the statistical properties and power for Johansens test are generally higher than for alternative procedures.

To check for stationarity arising from a linear combination of variables, the following AR representation for a vector VTS made up of $n$ variables is used,

$$VTS_t = c + \sum_{i=1}^{s-1} \phi_i Q_{it} + \sum_{i=1}^{k} \pi_i VTS_{t-i} + \varepsilon_t, \tag{3}$$

where VTS is at most I(1), $Q_{it}$ are seasonal dummies (*i.e.*, a vector of non-stochastic variables) and $c$ is a constant. It is not necessary that all variables that make up VTS be I(1). To find cointegration in the system only two variables in the system need be I(1). If an error-correction term is appended, this becomes:

$$\Delta VTS_t = c + \sum_{i=1}^{s-1} \phi_i Q_{it} + \sum_{i=1}^{k-1} \Gamma_I \, \Delta VTS_{t-i} + \prod VTS_{t-k} + \varepsilon_t, \tag{4}$$

which is basically a vector representation of Equation (1) with seasonal dummies added. All long-run information is contained in the levels term $\Pi VTS_{t-k}$ and the short-run information in the differences $\Delta VTS_{t-i}$. The above equation would have the same degree of integration on both sides only if $\Pi = 0$ (the series are not cointegrated) or $\Pi V_{t-k}$ is I(0), which infers cointegration. In order to test for cointegration, the validity of $H_1(r)$, shown below, is tested:

$$H_1(r) : \Pi = \alpha\beta',$$

where $\beta$ is a matrix of cointegrating vectors and $\alpha$ represents a matrix of error correction coefficients. The hypothesis $H_1(r)$ implies that the process $\Delta VTS_t$ is stationary, $VTS_t$ is nonstationary and $\beta' VTS_t$ is stationary (Johansen, 1991). The Johansen method yields the Trace and the $L_{max}$ statistics that enable determination of the number of cointegrating vectors.

***Description of Data***

The data used in this study are obtained from the National Council of Real Estate Investment Fiduciaries (NCREIF) Classic Property Index (1996:2). For most time series, total values (prices) commence in 1983:1 and end in 1996:2. NCREIF provides quarterly rates of return. The use of quarterly data, and the lesser frequency of observations, may create some misgivings about the power of unit-root and cointegration tests against an alternative hypothesis, such as a trend stationary model. However, Hakkio and Rush (1991), from Monte Carlo simulations, show that the frequency of information plays a minor role, rather it is the length of the data series, the span, which is more important in discerning whether the time series are co-integrated or not. This finding is also supported by Shiller and Perron (1985) and Perron (1989) who find that changing the frequency of observations while keeping the sample length fixed is not helpful when testing for cointegration because it is mainly a long-run property.

The literature points to numerous problems associated with the NCREIF indices. For instance, Fisher and Webb (1992) detail that raw appraisal-based NCREIF data suffer from a number of limitations. Such limitations may potentially bias the results from a study that uses this database. Some of the limitations identified in the literature are serial correlation, appraisal smoothing of the data, quarterly appraisals and problems of seasonality. However, these appraisal-based data may be unbiased in the long run (Webb, Miles and Guilkey, 1992). The objective in this study is to estimate long-run relationships. Hence, some of the problems associated with Russell-NCREIF data series are mitigated when long-run relationships are examined. Importantly, some of the features in the Johansen methodology also address these problems. The Johansen cointegration methodology treats all variables as endogenous and corrects for serial correlation and seasonality (the lag length that corrects for these problems is identified using AIC).

The data used is ex post. To ensure the usefulness of such data possible breakpoints are noted. The literature suggests that demand lagged supply in the 1980s, and that credit restraints characterized the early 1990s (Fisher and Webb, 1992). Noting these structural breakpoints we shorten the window examined into two periods, 1983–1989 and 1990–1996. This provides a more (shorter) stable environment to study the relationships between assets. Such spanning is not without precedent (*e.g.*, Rudolph and Griffith, 1997). However, care has to be taken in shortening data windows as sufficient spanning is needed to ensure that the estimates obtained are reliable. Spanning is an important issue in cointegration analysis (Hakkio and Rush, 1991). Fortunately, the overall length of the time series allows us to compartmentalize the data. Some have shown that return computations based on appraisal data are generally biased (Giliberto, 1988). Others, like Gau and Wang (1990) suggest otherwise. They find the bias to be very small in the annual period returns for portfolio properties. However, we take precautions to avoid such biases, small as they might be. We use only index prices in this study. Cointegration permits this. Also, the seasonal adjustment in the Johansen tests account for residual biases, if any. We are confident that our estimates are reliable.

To provide a comprehensive analysis NCREIF data by geographical region is studied. The data is grouped on the following basis:

**Regions:** East, Midwest, South and West.
**Divisions:** Northeast Division and Mideast Division in the East Region.
East North Central Division and West North Central Division in the Midwest Region.
Southeast Division and Southwest Division in the South Region.
Pacific Division and Mountain Division in the West Region.

The disaggregation into divisions allows us to examine relationships in the real estate markets with a finer level of precision.

## Results of Empirical Tests

The summary statistics for the regional property are provided in Exhibit 1. Panel A provides statistics for the first time period, 1983–1989, and Panel B provides similar statistics for the second time period, 1990–1996. Notice that standard deviation and kurtosis are generally smaller in the second period, compared with the first. The significance of the Ljung-Box (Q) statistics for both time periods is very apparent. The Ljung-Box (Q) statistics test for the absence of autocorrelation. Significant values for this statistic reject the absence of autocorrelation. Remarkably, the *Q*-Statistic is significant in all cases (see the Ljung-Box Q-Statistics for twelve lags). The evidence suggests that all time series are characterized by autocorrelation. These results confirm the findings of Myer and Webb (1994).

Since, the time series exhibit the presence of autocorrelations, we determine the lag length that would eliminate such autocorrelations using the AIC. The lag length is selected by minimizing the AIC over different choices for the length of the lag. The values of AIC are formulated by computing the value of the equation $T \log(RSS) + 2K$, where $K$ is the number of regressors, $T$ is the number of observations and $RSS$ is the residual sum of squares. These results are shown in Exhibit 2 (*Nlags*) along with the results of the Philips-Perron unit root tests. Notice that the time series require a range of lags to correct for the presence of autocorrelation. For instance, the time series belonging to East require four lags to minimize the AIC and purge autocorrelations, whereas Midwest requires no autocorrelation correction. Any autocorrelation in the time series belonging to Midwest decays quickly. To check for robustness, we carry out alternate procedures (Schwarz Criterion) to correct for autocorrelation and find the results to be mostly unchanged. These tests are not reported but available on request.

### *Tests for Nonstationarity of Each Time Series Using the Philips-Perron Test*

The time series are tested for a unit root using the P&P tests. The number of lag lengths, chosen by minimization of AIC, enables elimination of autocorrelations in the time series. To check for nonstationarity in the time series, we examine the non-rejection of $\alpha_1 = 0$. The P&P tests without trend suggest that most of the time series are nonstationary (*i.e.*, non-rejection of $\alpha_1 = 0$), clearly suggesting the need for cointegrated methodologies. Few exceptions, like East North Central, reject Southeast nonstationarity. However, for the model with trend we find no instances of stationarity.

**Exhibit I**
**Sample Statistics**

| Series | Mean | Std. Dev. | Skewness | Kurtosis | Ljung-Box |
|---|---|---|---|---|---|
| Panel A: 1983–1989 | | | | | |
| Regions | | | | | |
| East | 409.28 | 95.62 | −0.07 | −1.20 | 88.86* |
| Midwest | 237.71 | 42.83 | −0.10 | −1.29 | 91.88* |
| South | 246.44 | 17.30 | −1.27* | 0.75 | 50.16* |
| West | 327.73 | 55.99 | −0.05 | −0.96 | 81.44* |
| Divisions | | | | | |
| Northeast | 175.54 | 43.01 | −0.07 | −1.12 | 86.42* |
| Midwest | 149.98 | 31.17 | −0.01 | −1.40 | 94.75* |
| East N. Central | 145.85 | 28.50 | −0.04 | −1.31 | 92.13* |
| West N. Central | 139.46 | 20.60 | −0.29 | −1.14 | 89.55* |
| Southeast | 147.89 | 23.29 | −0.52 | −0.89 | 54.11* |
| Southwest | 109.51 | 7.22 | 0.34 | −1.59 | 62.86* |
| Pacific | 148.35 | 28.84 | 0.12 | −1.00 | 82.97* |
| Mountain | 136.59 | 14.21 | −1.15** | 0.13 | 68.94* |
| Panel B: 1990–1996 | | | | | |
| Regions | | | | | |
| East | 540.24 | 32.81 | 0.41 | −0.67 | 45.95* |
| Midwest | 306.06 | 19.32 | 0.83*** | 0.07 | 47.55* |
| South | 266.70 | 21.53 | 0.93*** | 0.04 | 58.44* |
| West | 409.71 | 25.69 | −0.08 | −1.33 | 52.50* |
| Divisions | | | | | |
| Northeast | 220.38 | 16.07 | 0.43 | −1.30 | 52.95* |
| Midwest | 202.83 | 14.24 | 1.24** | 0.90 | 47.39* |
| East N. Central | 194.74 | 11.20 | 0.76 | 0.02 | 44.18* |
| West N. Central | 162.88 | 13.02 | 0.86*** | −0.04 | 53.56* |
| Southeast | 180.51 | 17.75 | 0.92*** | −0.03 | 60.24* |
| Southwest | 108.86 | 6.63 | 0.93*** | 0.12 | 54.12* |
| Pacific | 193.50 | 12.64 | −0.13 | −1.46 | 62.03* |
| Mountain | 141.77 | 15.54 | 0.65 | −0.54 | 58.56* |

Note: The quarterly data time series begins 1983:1 and ends 1996:2. The data are divided into two subperiods: 1983–1989 and 1990–1996. The Ljung-Box *Q*-Statistic contains twelve lags and is distributed as $\chi^2$ with *n* degrees of freedom.
*1% level of significance.
**5% level of significance.
***10% level of significance.

The results are not surprising. The literature identifies nonstationarity (*i.e.*, unit roots) in several time series (*e.g.*, Doukas and Rahman, 1987; Phillips and Perron, 1988; and Brenner and Kroner, 1995). For instance, Doukas and Rahman report the presence of unit roots in foreign currency futures price series. Consistent with these studies, the results provide evidence that prevents rejection of the null hypothesis of

**Exhibit 2**
**Tests of Stationarity using Phillipa Perron Tests**

| Series | Nlags | No Trend $\alpha_1 = 0$ | No Trend $\alpha_0 = \alpha_1 = 0$ | With Trend $\alpha_1 = 0$ | With Trend $\alpha_0 = \alpha_1 = \alpha_2 = 0$ | With Trend $\alpha_1 = \alpha_2 = 0$ |
|---|---|---|---|---|---|---|
| Panel A: 1983–1989 | | | | | | |
| Regions | | | | | | |
| East | 4 | −2.15 | 24.87* | −1.41 | 45.29* | 2.44 |
| Midwest | 0 | −0.87 | 14.48* | −1.22 | 26.83* | 1.07 |
| South | 2 | −3.30* | 9.03* | −1.94 | 6.17* | 5.62* |
| West | 0 | −0.37 | 12.61* | −1.76 | 8.14* | 1.58 |
| Divisions | | | | | | |
| Northeast | 4 | −1.87 | 26.14* | −2.16 | 16.46* | 3.72 |
| Midwest | 0 | −0.34 | 9.25* | −1.51 | 16.12* | 1.16 |
| East N. Central | 5 | −0.49 | 10.35* | −2.82 | 9.15* | 4.17 |
| West N. Central | 0 | −2.56 | 7.31* | −0.31 | 8.03* | 6.47* |
| Southeast | 2 | −2.67* | 19.73* | −0.36 | 14.29 | 30.21* |
| Southwest | 2 | −1.42 | 1.04 | −1.97 | 2.44 | 3.64 |
| Pacific | 4 | −1.32 | 1.24 | −1.52 | 1.77 | 2.25 |
| Mountain | 2 | −5.87* | 14.32* | −1.97 | 28.87* | 17.34* |
| Panel B: 1990–1996 | | | | | | |
| Regions | | | | | | |
| East | 1 | −0.30 | 0.34 | −0.69 | 6.81* | 9.43* |
| Midwest | 4 | −0.44 | 0.67 | −0.29 | 5.72* | 6.81* |
| South | 2 | −1.62 | 3.19 | −0.36 | 9.69* | 10.04* |
| West | 5 | −1.09 | 0.67 | −0.46 | 1.96 | 2.80 |
| Divisions | | | | | | |
| Northeast | 1 | −1.31 | 0.91 | −0.36 | 4.90* | 7.28* |
| Midwest | 1 | −2.28 | 5.07* | −0.90 | 9.83* | 10.29* |
| East N. Central | 4 | −0.26 | 0.56 | −0.18 | 3.71 | 4.32 |
| West N. Central | 2 | −0.99 | 1.22 | −0.94 | 20.28* | 25.95* |
| Southeast | 2 | −1.72 | 3.32 | −0.20 | 13.56* | 14.68* |
| Southwest | 2 | −1.19 | 2.42 | −0.10 | 4.45* | 4.01 |
| Pacific | 5 | −1.25 | 0.80 | −0.15 | 1.01 | 1.50 |
| Mountain | 3 | −0.42 | 0.59 | −0.42 | 8.52* | 10.75* |
| Asymptotic Critical Values | | | | | | |
| 10% level | | −2.60 | 3.94 | −3.18 | 4.30 | 5.60 |

Note: The Phillips-Perron test is computed with a constant term. The tests are also computed with and without linear trend. If the hypothesis $\alpha_1 = 0$ cannot be rejected, then the series is said to have a unit root and is nonstationary and conversely, if the hypothesis, $\alpha_1 = 0$, is rejected, then it is concluded that the series does not contain a unit root and is stationary. Tests involving parameters $\alpha_0$ and $\alpha_2$ verify the presence of drift and trend. Inclusion of the *p*-lagged values ensures a white-noise series. The number of lags is determined by a test of significance such as the AIC. The distribution of the ordinary *t*- and *F*-Statistics computed for the regressions do not have the expected distributions. Thus, to test the various hypotheses critical values have been computed using Monte Carlo techniques and are tabulated in various references (*e.g.,* Davidson and MacKinnon, 1993).
*Significant at the 10% level.

nonstationarity (*i.e.*, non-rejection of $\alpha_1 = 0$) at the 10% level of significance. Critical values at the 10% level are provided in the last row of Exhibit 2.

If the series has a unit root (*i.e.*, $\alpha_1 = 0$) and the hypothesis $\alpha_0 = \alpha_1 = 0$ is not rejected, this suggests the presence of the constant term or drift in the series. Overall, the results in Exhibit 2 suggest the absence of drift in most of the series, although the time series, such as Southwest, show the presence of drift. If the series has a unit root ($\alpha_1 = 0$) and the hypothesis $\alpha_1 = \alpha_2 = 0$ is not rejected, this denotes presence of trend and suggests that the model with trend is probably more appropriate for the series. We observe that the model with trend is appropriate in most instances. Thus, a model that includes a term for trend is necessary, while the inclusion of drift may not be as important.

### *Johansen Tests to Determine Cointegration Rank*

Many factors favor cointegration among real estate assets. It is well accepted that common macroeconomic fundamentals will produce comovements in prices. Other factors may also influence cointegration. Such factors include falling (post-1980) mortgage rates and increasing population levels. On the other hand, shifting demographics, differences in legislation across regions, and the lead-lag effects of economic forces across state and regional economies, will reduce the level of cointegration. Fisher and Webb (1992), for instance, draw reference to a First Chicago (1990) study that finds regional employment, population, income, unemployment, retail sales and housing starts to be a function of net reports, defense outlays, oil prices, industrial production and farm income. Importantly, different regions of the country responded differently to changes in the factors during the 1980s. This study attempts to discern regional and local differences by examining real estate in two periods (1983–1989 and 1990–1996) and by geographical breakdown. While we expect real estate assets to exhibit cointegration on a regional basis and with financial assets, especially given the evidence in Chaudhry, Myer and Webb (1999), evidence of differences in the degree of cointegration will not be unexpected.

The Johansen test is used to determine cointegration rank. When the data is nonstationary, the vector autoregressive model can be employed to distinguish long-run patterns about cointegrated relationships and this is the basis of the estimation procedure developed by Johansen (*e.g.*, Johansen and Juselius, 1991).[9]

## Tests for Cointegration Rank for Regions and Divisions

The results for the groups using Johansens method are presented in Exhibit 3. Two tests, the Trace statistic and the $L_{max}$ statistic, are reported. These are basically likelihood ratio tests where the null hypothesis is $L_{r+1} = L_{r+2} = \ldots = L_p = 0$, indicating that the system has $p - r$ unit roots, where $r$ is the number of cointegrating vectors. The rank is then determined using a sequential approach starting with the hypothesis of $p$ unit roots. If this is rejected then the next hypothesis $L_2 = L_3 = \ldots = L_p = 0$ is tested and so on.[10]

**Exhibit 3**
**Johansen's Integration / Segmentation Tests—Real Estate Assets**

| | | 1983–1989 | | 1990–1996 | |
|---|---|---|---|---|---|
| | *r* = | $L_{MAX}$ | Trace | $L_{MAX}$ | Trace |
| Regions | 0 | 19.25* | 48.02* | 41.38* | 82.44* |
| | 1 | 14.91* | 28.77* | 20.98* | 41.06* |
| | 2 | 10.01 | 12.86 | 16.46* | 20.08* |
| | 3 | 1.76 | 1.76 | 2.62 | 2.62 |
| Asymptotic Critical Values | | | | | |
| | 0 | 17.15 | 43.84 | | |
| | 1 | 13.39 | 26.70 | | |
| | 2 | 10.60 | 13.31 | | |
| | 3 | 2.71 | 2.71 | | |
| Divisions | | | | | |
| East | 0 | 17.37* | 18.62* | 31.65* | 35.45* |
| | 1 | 1.25 | 1.25 | 3.80* | 3.80* |
| Midwest | 0 | 10.15 | 11.75 | 16.96* | 16.97* |
| | 1 | 1.60 | 1.60 | 0.01 | 0.01 |
| South | 0 | 8.12 | 9.14 | 9.20 | 10.33 |
| | 1 | 1.01 | 1.01 | 1.13 | 1.13 |
| West | 0 | 16.47* | 22.30* | 23.93* | 40.14* |
| | 1 | 5.83* | 5.83* | 6.21* | 6.21* |
| Asymptotic-Critical Values | | | | | |
| | 0 | 10.60 | 13.31 | | |
| | 1 | 2.71 | 2.71 | | |

Where *r* refers to the number of cointegrating vectors. If $r = 0$ hypothesis is rejected whereas $r = 1$ is not, then the system is tied by one cointegrating vector. In general, it is possible to examine sequentially the hypotheses that $r \leq 1$, $r \leq 2$, etc. If the null hypothesis cannot be rejected for, say $r \leq r_0$ but it has been rejected for $r \leq r_0 - 1$, the conclusion would be that the number of cointegrating vectors are equal to $r_0$.
*Significant at the 10% level.

The results for the Johansen test for the regions and divisions are shown in Exhibit 3. The regions are composed of four time series, East, Midwest, South and West. The results suggest rejection of the hypothesis of no cointegration ($r = 0$) for the two time periods. Interestingly, real estate in the different regions seem to exhibit a closer relationship in the second period. We have three common factors binding the time series in this period, compared with two factors in the first. Also, the Trace statistic and the $L_{max}$ statistic values are higher in the second period, compared with the first. For example, for $r = 0$, the $L_{max}$ statistic for the 1990–1996 is 41.38, whereas it is less than half this value, 19.25, for 1983–1989. The Trace statistic and the $L_{max}$ values also suggest stronger binding in the second period among real estate in the different regions. Many plausible explanations for the stronger binding in the second period can be offered. Possible explanations include a stronger economy in the 1990s (compared with the 1980s) that may have eased the high vacancy rates that prevailed in the first period, contributing to weaker local influences in the second period. This might have heightened cointegration between the regions. The literature notes that the 1980s were a period when demand lagged supply (Fisher and Webb, 1992). Thus, local influences may have been stronger in the first period. Underpinning this possibility is the finding of Smith and Tesarek (1991). They suggest that excessive vacancy rates in Houston in the 1980s had an impact on the local real estate markets. Other issues, such as deregulation in the 1980s, the Tax Reform Act of 1986, and the higher frequency of secondary market transactions in the 1990s, may have also played a role.

The results for divisions are also provided in Exhibit 3. The divisional tests provide information that is more detailed than the first set of tests that examine relationships among regions. In general, we find relationships to have strengthened between the divisions that make up a region in the second period. Specifically, stronger relationships in the second period are found in divisions that make up the East, Midwest and West. To illustrate, for the divisions that make up the East region, the Trace statistic is 18.62 in the first period, but increases to 35.45 in the second period. However, this is not so with divisions that comprise the South region. Very little evidence of any difference between the two periods is detected. The South region comprises the Southeast division and the Southwest division. The results suggest no cointegration, implying that these two divisions are segmented. Lack of cointegration may have arisen because it was the Southeast market that suffered heavily from the effects of the energy-induced recession and excessive vacancy rates. These effects may have taken a longer time to dissipate in this division.

### ***Tests for Cointegration Rank between Real Estate and Financial Assets***

A series of similar tests, is carried out on systems of assets composed of real estate groups and a financial asset. An example of a system would be the real estate Regions group (composed of East, Midwest, South and West) and the S&P 500 Index. This part of the analysis also examines cointegration between individual real estate regions, such as East, and financial assets, and between divisions, such as Northeast and financial assets. Financial assets are represented by the S&P 500 (equity) and the Lehman Govt./Corp. Index (Treasury bonds). Prices for financial assets are also obtained from the NCREIF database. Given the close relationship between inflation and real estate asset prices, we run these tests with a proxy for inflation (CPI). Non-inclusion of the CPI almost always results in rank that is one less than rank with the CPI. This illustrates the close relationship between inflation and real estate assets. Results from these tests are shown in Exhibit 4.

The results for systems support the earlier evidence. In general, the test statistics are higher in the second period compared with the first period. Evidence of this is present in almost all pairings. Consider cointegration in the system composed of Regions and bond. Five common cointegrating vectors are detected in both periods, but the strength of the binding has increased dramatically in the second period. For example, for

**Exhibit 4**
**Johansen's Integration / Segmentation Tests—**
**Real Estate and Financial Assets (with CPI)**

| | | 1983–1989 | | 1990–1996 | |
|---|---|---|---|---|---|
| | *r* = | $L_{MAX}$ | Trace | $L_{MAX}$ | Trace |
| Regions with S&P | 0 | 58.18* | 143.40* | 61.77* | 174.30* |
| | 1 | 33.29* | 85.22* | 45.43* | 112.53* |
| | 2 | 18.34* | 51.93* | 32.25* | 67.09* |
| | 3 | 16.97* | 33.58* | 23.57* | 34.84* |
| | 4 | 12.24* | 16.61* | 12.27* | 18.33* |
| | 5 | 2.38 | 2.38 | 0.01 | 0.01 |
| Regions with Bond | 0 | 40.58* | 129.43* | 61.02* | 169.87* |
| | 1 | 38.27* | 88.85* | 42.31* | 108.85* |
| | 2 | 19.94 | 50.57* | 31.16* | 66.54* |
| | 3 | 13.81* | 30.64* | 19.46* | 35.38* |
| | 4 | 10.74* | 16.83* | 13.81* | 17.92* |
| | 5 | 2.09 | 2.09 | 2.11 | 2.11 |
| Asymptotic Critical Values | | | | | |
| | 0 | 24.63 | 89.37 | | |
| | 1 | 20.90 | 64.74 | | |
| | 2 | 17.15 | 43.84 | | |
| | 3 | 13.39 | 26.70 | | |
| | 4 | 10.60 | 13.31 | | |
| | 5 | 2.71 | 2.71 | | |
| Regions | | | | | |
| East with S&P | 0 | 28.66* | 40.95* | 36.23* | 50.57* |
| | 1 | 10.52 | 12.29 | 11.50* | 14.34* |
| | 2 | 1.77 | 1.77 | 2.34 | 2.34 |
| East with Bond | 0 | 19.28* | 29.95* | 21.59* | 40.26* |
| | 1 | 8.71 | 10.68 | 13.42* | 18.68* |
| | 2 | 1.97 | 1.97 | 2.26 | 2.26 |
| Midwest with S&P | 0 | 11.19 | 14.13 | 23.26* | 38.57* |
| | 1 | 2.26 | 2.95 | 13.29* | 15.31* |
| | 2 | 0.32 | 0.32 | 2.02 | 2.02 |
| Midwest with Bond | 0 | 12.52 | 15.50 | 17.62* | 28.59* |
| | 1 | 2.97 | 2.98 | 8.02 | 10.96 |
| | 2 | 1.27 | 1.27 | 1.94 | 1.94 |
| South with S&P | 0 | 13.05 | 17.05 | 25.04* | 38.24* |
| | 1 | 2.71 | 4.00 | 11.64* | 13.40* |
| | 2 | 0.52 | 0.52 | 1.56 | 1.56 |
| South with Bond | 0 | 12.33 | 18.31 | 19.28* | 29.95* |
| | 1 | 2.88 | 3.97 | 8.71 | 10.68 |
| | 2 | 0.95 | 0.95 | 1.97 | 1.97 |
| Asymptotic Critical Values | | | | | |
| | 0 | 13.39 | 26.70 | | |
| | 1 | 10.60 | 13.31 | | |
| | 2 | 2.71 | 2.71 | | |

**Exhibit 4** (continued)
**Johansen's Integration / Segmentation Tests—**
**Real Estate and Financial Assets (with CPI)**

| | | 1983–1989 | | 1990–1996 | |
|---|---|---|---|---|---|
| | $r$ = | $L_{MAX}$ | Trace | $L_{MAX}$ | Trace |
| West with S&P | 0 | 11.66 | 23.86 | 39.64* | 50.75* |
| | 1 | 10.24 | 12.20 | 13.43* | 15.10* |
| | 2 | 1.36 | 1.36 | 1.67 | 1.67 |
| West with Bond | 0 | 19.88* | 29.42* | 16.94* | 31.54* |
| | 1 | 7.93 | 9.54 | 12.23* | 14.60* |
| | 2 | 1.61 | 1.61 | 2.41 | 2.41 |
| Divisions | | | | | |
| Northeast with S&P | 0 | 27.22* | 40.80* | 43.65* | 52.56* |
| | 1 | 8.38 | 8.90 | 11.11* | 13.48* |
| | 2 | 1.25 | 1.25 | 1.96 | 1.96 |
| Northeast with Bond | 0 | 22.46* | 35.40* | 39.19* | 52.07* |
| | 1 | 9.67 | 12.94 | 11.81* | 13.87* |
| | 2 | 2.27 | 2.27 | 1.06 | 1.06 |
| Mideast with S&P | 0 | 25.26* | 36.52* | 29.82* | 45.25* |
| | 1 | 9.75 | 11.26 | 13.32* | 15.43* |
| | 2 | 1.51 | 1.51 | 2.11 | 2.11 |
| Mideast with Bond | 0 | 21.20* | 31.22* | 18.38* | 36.86* |
| | 1 | 8.65 | 10.02 | 12.66* | 18.28* |
| | 2 | 1.37 | 1.37 | 2.62 | 2.62 |
| East. N. Central with S&P | 0 | 10.81 | 17.40 | 21.64* | 37.27* |
| | 1 | 4.16 | 6.59 | 13.63* | 15.63* |
| | 2 | 0.68 | 0.68 | 2.00 | 2.00 |
| East N. Central with Bond | 0 | 13.21 | 20.87 | 15.62* | 25.53* |
| | 1 | 5.21 | 7.06 | 7.66 | 9.91 |
| | 2 | 1.29 | 1.29 | 2.26 | 2.26 |
| West N. Central with S&P | 0 | 17.77* | 27.91* | 32.95* | 46.47* |
| | 1 | 5.09 | 5.15 | 11.97* | 13.52* |
| | 2 | 2.17 | 2.17 | 1.54 | 1.54 |
| West N. Central with Bond | 0 | 22.07* | 29.04* | 24.35* | 37.32* |
| | 1 | 6.89 | 6.97 | 9.37 | 12.96 |
| | 2 | 0.86 | 0.86 | 2.59 | 2.96 |
| Asymptotic Critical Values | | | | | |
| | 0 | 13.39 | 26.70 | | |
| | 1 | 10.60 | 13.31 | | |
| | 2 | 2.71 | 2.71 | | |
| Southeast with S&P | 0 | 10.23 | 14.46 | 24.78* | 36.79* |
| | 1 | 3.84 | 4.23 | 9.99 | 12.01 |
| | 2 | 1.59 | 1.59 | 2.02 | 2.02 |
| Southeast with Bond | 0 | 12.93 | 19.30 | 20.15* | 31.02* |
| | 1 | 4.14 | 4.38 | 9.68 | 10.86 |
| | 2 | 0.16 | 0.16 | 1.19 | 1.19 |

**Exhibit 4** (continued)
**Johansen's Integration / Segmentation Tests—**
**Real Estate and Financial Assets (with CPI)**

| | $r =$ | 1983–1989 $L_{MAX}$ | 1983–1989 Trace | 1990–1996 $L_{MAX}$ | 1990–1996 Trace |
|---|---|---|---|---|---|
| Southwest with S&P | 0 | 15.56* | 27.91* | 23.84* | 38.88* |
| | 1 | 6.45 | 6.45 | 13.84* | 15.04* |
| | 2 | 0.52 | 0.52 | 1.20 | 1.20 |
| Southwest with Bond | 0 | 13.56* | 26.71* | 18.18* | 28.87* |
| | 1 | 6.55 | 8.05 | 7.85 | 10.68 |
| | 2 | 1.31 | 1.31 | 2.43 | 2.43 |
| Pacific with S&P | 0 | 11.72 | 22.16 | 37.50* | 46.59* |
| | 1 | 8.55 | 10.44 | 7.48 | 9.09 |
| | 2 | 2.17 | 2.17 | 1.62 | 1.62 |
| Pacific with Bond | 0 | 18.42* | 27.13* | 19.63* | 33.99* |
| | 1 | 7.93 | 8.71 | 12.84* | 14.36* |
| | 2 | 0.79 | 0.79 | 1.52 | 1.52 |
| Mountain with S&P | 0 | 20.25* | 32.60* | 37.98* | 53.97* |
| | 1 | 9.31 | 12.38 | 14.67* | 15.99* |
| | 2 | 2.07 | 2.07 | 1.32 | 1.32 |
| Mountain with Bond | 0 | 23.70* | 33.30* | 23.31* | 39.32* |
| | 1 | 8.04 | 9.60 | 11.17* | 16.00* |
| | 2 | 1.56 | 1.56 | 2.63 | 2.63 |
| Asymptotic Critical Values | | | | | |
| | 0 | 13.39 | 26.70 | | |
| | 1 | 10.60 | 13.31 | | |
| | 2 | 2.71 | 2.71 | | |

Where $r$ refers to the number of cointegrating vectors. If $r = 0$ hypothesis is rejected whereas $r = 1$ is not, then the system is tied by one cointegrating vector. In general, it is possible to examine sequentially the hypotheses that $r \leq 1$, $r \leq 2$, etc. If the null hypothesis cannot be rejected for, say $r \leq r_0$ but it has been rejected for $r \leq r_0 - 1$, the conclusion would be that the number of cointegrating vectors are equal to $r_0$.
*Significant at the 10% level.

$r = 1$ the Trace statistic is 129.43 in the first period. In the second period, the value is 169.87. The many factors outlined before may have lead to this. Webb, Miles and Guilkey (1992) find that the unsystematic risk (or variances) in transaction-driven individual property returns can be diversified away if a portfolio of real estate assets is created. On an aggregate basis, if unsystematic risk in real estate that is impacted by the local factors, is mostly eliminated, it would be systematic risk that would be common to both real and financial assets. This creates linkages between financial and real assets.

For regions, we find stronger binding with financial assets in the second period. Consider the Midwest region. With the S&P, the $L_{max}$ statistic for $r = 0$ is 23.26 and significant at the 10% level of significance during the second period. This is not so in the first period. The $L_{max}$ statistic for $r = 0$ is 11.19 and insignificant. In many instances, such as East with bond, in the second period two cointegrating vectors are detected. Only two exceptions are found, Midwest with bond and South with bond, where one cointegrating vector is detected. The results for divisions and financial assets are very similar. Stronger binding is detected in the second period between real estate divisions and financial assets. Some insight into the weaker relationship between Midwest and bond and South and bond may be discerned by examining the results in the second period for divisions and financial assets. Only one cointegrating vector is found for the pairing between bond and East North Central and West North Central divisions. With the bond-Southeast and bond-Southwest pairings, we also find one cointegrating vector. Local influences may be stronger in these regions even in the 1990s. In fact, we also find evidence of one cointegrating with Southeast and the S&P. Such information may be useful to investors seeking to diversify holdings.

## Summary and Conclusion

This study investigates the long-term stochastic properties of several real estate assets. The analysis uses the geographical categorization followed by NCREIF. The literature points to numerous problems associated with the NCREIF indices. Some of the limitations identified in the literature are serial correlation, appraisal smoothing of the data, quarterly appraisals and problems of seasonality. However, these appraisal-based data may be unbiased in the long run (Webb, Miles and Guilkey, 1992). We use cointegration methods that mitigate a large portion of these problems, and examine long-term relationships. We also examine the long-term relationships between real estate assets and financial assets.

NCREIF categorizes real estate assets into four regions: East, Midwest, South and North. The regions are further classified into eight divisions. The divisions include the Northeast and Mideast Division in the East Region, East North Central Division and West North Central Division in the Midwest Region, Southeast Division and Southwest Division in the South Region, and Pacific Division and Mountain Division in the West Region. Financial assets are represented by the S&P 500 (equity) and the Lehman Govt./Corp. Index (Treasury bonds). The nature of data collected by NCREIF is ex post. Thus, we identify possible breakpoints in the data to shorten the windows examined and to ensure that the estimates obtained are meaningful. The literature suggests that demand lagged supply in the 1980s, and that credit restraints characterized the early 1990s (Fisher and Webb, 1992). These structural breakpoints are used to divide the data into two periods: 1983–1989 and 1990–1996. This provides a more stable (shorter) environment to study the relationships between assets.

The initial exploratory tests find that most assets in our sample exhibit nonstationarity. Thus, it seems logical that cointegrated processes should be the relevant methodology. To eliminate non-decaying autocorrelations, we identify lag lengths, by minimizing the AIC. The results for the Johansen test for the regions and divisions suggest rejection of the hypothesis of no cointegration ($r = 0$) for the two time periods. However, real estate in the different regions seem to exhibit a closer relationship in the second period. The results for divisions also suggest that cointegration between real estate assets may have strengthened between the divisions that make up a region in the second period. Specifically, stronger relationships in the second period are found in divisions that make up the East, Midwest and West. However, this is not so with divisions that comprise the South region. Very little evidence of any difference between the two periods is detected. The South region comprises the Southeast division and the Southwest division. The results suggest no cointegration, implying that these two divisions are segmented.

Strong linkages between real estate regions and financial assets are noted in the second period. In many instances, such as East with bond, two cointegrating vectors are detected. Only two exceptions are found, Midwest with bond and South with bond, where one cointegrating vector is detected. The results for regional divisions and financial assets are very similar. Stronger binding is detected in the second period between real estate divisions and financial assets. Some insight into the weaker relationship between Midwest and bond and South and bond may be discerned by examining the results in the second period for divisions and financial assets. Only one cointegrating vector is found for the pairing between bond and East North Central and West North Central divisions. With the bond-Southeast and bond-Southwest pairings, we also find one cointegrating vector.

The results from this study have important implications for investors. The presence of cointegration between assets provides investors with cross-hedging opportunities, especially if markets differ in liquidity. On the other hand, the lack of cointegration should interest those seeking diversification. Strong local variation may have largely contributed to real estate pricing in the 1980s. However, local effects may be diminishing in the 1990s, although not uniformly across the country. Local effects still appear to be strong in the South. Future research to identify the importance of specific pricing factors in the eight divisions may be warranted. Overall, the information derived from our analysis sheds light on linkages among real estate assets and between real estate assets and financial assets. Such an analysis provides a framework for creating diversified portfolios to minimize risks in the long run.

## Endnotes

[1] Miles (1990) reviews the importance of U.S. commercial real estate in the U.S. economy. He estimates real estate values to range from $800 billion to $5 trillion. Even if the average value of this range is taken, the cumulative worth of real estate is almost equal in value to stock market capitalization values. In an earlier study, conducted by MIT in 1987, estimates of real estate value amounted to 25% of corporate value.

[2] The following nomenclature is used throughout the text. A GROUP is composed of two or more assets (time series). For example, the Regions group is composed of time-series belonging to East, Midwest, South and West. A system is made up of two or more groups or a group and an asset (time series).

[3] For instance, Nelson and Plosser (1982) find evidence consistent with the notion that macroeconomic time series are better characterized as nonstationary processes that have no tendency to return to the deterministic path, rather than as stationary fluctuations around a

deterministic trend. Brenner and Kroner (1995) also report the prevalence of stochastic trends in financial time series such as stock prices, foreign exchange rates, forward prices and futures prices. They also note that the presence of trends limits the set of statistical models that can be used to implement financial theories.

[4] The ECM has become quite popular and the use of such models is becoming more prevalent. The basic idea behind the ECM is straightforward. Disequilibrium in one period is corrected in the next. For example, in a two-variable system a typical error correction model would relate the change in one variable to past equilibrium errors, as well as to past changes in both variables (Engle and Granger, 1987).

[5] Engle and Granger (1987), point out that a series will be integrated of order $d$ ($I(d)$) if, when differenced $d$ times, the series has a stationary, invertible, ARMA representation. A system consisting of two or more series is said to be cointegrated if the series making up the system are integrated of the same order, but some linear combination of the series in the system is integrated of a lower order.

[6] Manuals, such as handbooks on the econometrics packages like SHAZAM and RATS provide a quick background on cointegration tests and references to related theory. For more detailed discussions and background texts such as *New Directions in Econometric Practice* by Charemza and Deadman should be consulted.

[7] The use of the AIC to determine lag length is not universal. Other procedures are used such as the Schwarz Criterion. The Schwarz Criterion is more restrictive than the AIC and will always choose a model smaller than or equal to the AIC. Some determine the lag length by successively testing shorter lag lengths as restrictions against longer lag lengths until significance is obtained. This procedure, generally, uses likelihood ratio methods (Harris, McInish, Shoesmith and Wood, 1995).

[8] P&P impose only weak restrictions on the error sequence, thus, this test is considered robust to a variety of heterogeneously distributed and weakly dependent innovations since P&P correct the standard errors used to compute $t$ values using the Newey and West (1987) correction.

[9] For a detailed discussion of both tests, the reader is referred to Johansen and Juselius (1990) and Johansen (1988, 1991, 1994).

[10] The symbols used in the tables are: SP, TBO for the S&P 500 and the Lehman Govt/Corp Index.

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