REITs and Inflation: A Long-Run Perspective

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Abstract. **We examine whether REITs provide an inflation hedge in the long run. We also investigate whether the apparent lack of a positive relationship between general prices and REIT returns in prior studies arises from the impact that stock market movements have on REITs. As in most prior research, regression analysis provides no evidence that REIT returns are positively related to temporary or permanent components of inflation measures. We rule out the possibility that a stock market-induced proxy effect is the cause for the apparent lack of relationship between REITs and inflation. On the other hand, we find some evidence that REITs provide a long-run inflation hedge. Johansen (1988) tests for cointegration isolate cointegrating vectors between alternate REIT indices and the CPI over the 1972–95 interval. However, the more standard residual-based cointegration techniques failed to provide similar evidence.**

Introduction

Several researchers have suggested that real estate investment trusts (REITs) tend to behave like other equities with respect to their inflation-hedging characteristics. It is now well documented that stock returns in the United States and several other countries are either unrelated or negatively related to inflation, inconsistent with the Fisher (1930) hypothesis (*e.g.*, Bodie, 1976; Jaffe and Mandelker, 1976; Fama and Schwert, 1977; Fama, 1981; Geske and Roll, 1983; Mandelker and Tandon, 1985; and Stulz, 1986). Most studies on the relationship between REIT returns and inflation arrive at similar conclusions (Murphy and Kleiman, 1989; Chan, Hendershott and Sanders, 1990; Park, Mullineaux and Chew, 1990; and Yobaccio, Rubens and Ketcham, 1995). Only a few studies, such as Gyourko and Linneman (1988) and Chen and Tzang (1988), indicate that REITs possess some inflation-hedging properties. Gyourko and Linneman document that REITs provide a partial hedge against the inflation rate derived from the Consumer Price Index (CPI) adjusted for the Home Purchase Price component. However, the authors also find that REITs act as a perverse hedge against unexpected inflation. Chen and Tzang find that REITs have some ability to hedge the expected component of inflation.[1]

The evidence from unsecuritized real estate has been far more favorable. For instance, Hartzell, Heckman and Miles (1987) document that a portfolio of commercial real estate provided an effective hedge against inflation over the 1973–83 interval.[2] An earlier study by Fama and Schwert (1977) suggested that residential real estate was

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a complete hedge against the expected and unexpected components of inflation. More recently, Rubens, Bond and Webb (1989) find that residential, commercial and farmland real estate provide at least partial hedges against inflation. The authors also find improvements in the hedging effectiveness of portfolios once real estate is included. Similarly encouraging results are obtained in Brueggeman, Chen and Thibodeau (1984), Ibbotson and Siegel (1984) and Miles and Mahoney (1997). In sum, the evidence on the inflation-hedging potential of real estate is very different across studies that employ unsecuritized real estate and those that employ securitized real estate.[3]

These differences in the evidence for unsecuritized real estate and REITs give rise to two interesting, albeit related, questions. First, is it only smoothing bias that causes the divergence in the evidence, or could it also be the result of the behavior of REITs themselves? It is now well known that REITs have a substantial stock market component, and as widely documented, stock returns have tended to be negatively related to the rate of inflation. Second, is there a longer-run relationship between inflation and REITs that standard econometric techniques fail to capture? Real estate markets are prone to long boom-and-bust cycles that are known to be out of sync with other markets (Grenadier, 1995). Out of sync cycles and the market component in REITs may be obfuscating a long-run real estate-inflation relationship. This line of reasoning is bolstered by prior evidence that there exists a temporally unstable relationship between expected inflation and REIT returns (Chan, Hendershott and Sanders, 1990).

We address the above two questions in this article. We also extend prior studies by conducting standard regression tests on the short-run relationship between REIT returns and the permanent and temporary components of inflation. However, the inflation decomposition techniques employed here differ from prior efforts in that we allow for a stochastic trend component in the inflation proxies. Prior findings of a weak and even perverse relationship between REIT returns and inflation continue to be supported by the standard regression analysis. Further, we find no evidence that the apparent lack of relationship between REITs and inflation is attributable to the stock market component in REITs.

We conduct Johansen (Johansen and Juselius, 1990) cointegration tests to examine the long-run relationship between REITs and inflation. We find evidence of cointegration when employing the CPI as the proxy for inflation. We find much weaker evidence of cointegration between the REIT index levels and the three-month or the one-year Treasury bill rate, the other proxies for inflation. Moreover, there is no evidence of cointegration between REITs and inflation, however measured, when we employ the more traditional residual-based cointegration methods. We suggest that the mixed evidence from the cointegration tests are symptomatic of *fractional* cointegration. In this scenario, REITs provide only a partial hedge against inflation in the long run.

Data

The investigation employs monthly and quarterly indices for all REITs (*AREIT*), equity REITs (*EREIT*), mortgage REITs (*MREIT*) and hybrid REITs (*HYREIT*), along with inflation proxies over the interval January 1972 through December 1995.[4] Following prior research, we employ the CPI and Treasury bill rates to derive proxies for inflation (see Park, Mullineaux and Chew, 1990; and Yobaccio, Rubens and Ketcham, 1995).

Exhibit 1 provides correlations and unit root tests for the series examined. Panel A presents the Pearsons coefficients for correlation between monthly REIT returns and alternate measures of inflation. The REIT returns and inflation rates are given by $((index_{i,t} - index_{i,t-1})/index_{i,t-1})*100)$, where the REIT and CPI indices are benchmarked 1972/01 = 100. The trailing one-year inflation rate (annualized) is employed to provide indications on the relationship of REIT returns and a slower-adjusting measure of inflation. The correlations suggest that the return and inflation rate proxies are negatively related or, at best, unrelated. The coefficients for the CPI rate are consistently negative, notably so for *EREIT* returns.

Panel B reports the Phillips-Perron (Phillips and Perron, 1986; and Phillips, 1987) *t*-Statistics that test for the null of nonstationarity, with the alternate that the series has

Exhibit 1
Pearsons Correlations and Unit Root Tests (1972/01–1995/12)

	AREIT	*EREIT*	*MREIT*	*HYREIT*	CPI	T-bill Rate
Panel A: Correlations: REIT Returns and Inflation						
CPI Inflation Rate	−0.08	−0.11	−0.08	−0.03		
Trailing 1-Yr Inflation Rate	0.00	−0.00	0.00	0.02		
3-Month T-bill Rate	−0.03	−0.06	−0.01	−0.01		
1-Year T-bill Rate	−0.01	−0.05	0.01	0.00		
Panel B: Phillips-Perron Unit Root Test Statistics						
No Trend	−1.79	−1.78	−1.84	−1.79	−1.63	−1.87
Trend	−2.51	−2.56	−1.58	−1.76	−1.07	−2.33
First Difference						
No Trend	−14.15*	−15.30*	−13.57*	−14.67*	−7.15*	−10.42*
Trend	−14.33*	−15.62*	−13.59*	−14.76*	−8.89*	−10.37*

The monthly REIT returns and the CPI inflation rates in Panel A are given by $(index_{i,t} - index_{i,t-1})/index_{i,t-1})*100$ where $index_i$ = 100 in 1970/01. The level REIT indices in Panel B are AREITs, EREITs, MREITs and HYREITs (1972/01 = 100 for the REIT indices). The CPI in Panel B is the CPI (1972/01 = 100) and the T-bill Rate represents the one-year T-bill rate. The critical values for the Phillips-Perron tests in Panel B are: Trend: −3.13 (10%), −3.41 (5%), −3.96 (1%); No Trend: −2.57 (10%), −2.86 (5%) and −3.43 (1%).
*Significance at the .01 level.

no unit root or that the series is I(0).[5] We cannot reject the null for any of the series considered. On the other hand, the statistics suggest the series are stationary in their first differences. The *t*-Statistics consistently reject the null that the differenced series are nonstationary at the .01 level. In other words, the level series may be appropriately described as integrated of order 1 or I(1), and therefore, are appropriate for the deployment in standard cointegration tests.[6] The characteristics of I(1) and I(0) processes are discussed further later in the article.

Regression Results

REIT Returns and Inflation

Prior studies that employ regression analysis to test the inflation hedging properties of real estate or REITs construct expected and unexpected components of inflation series by fitting inflation series to univariate Box-Jenkins (ARIMA-type) models. In this study, we decompose the inflation series into their permanent and temporary components using the technique suggested by Hodrick and Prescott (1980). The technique defines the permanent component $\bar{x}$ of a variable x as the one that minimizes the function:

$$\sum_{t=0}^{T} (x_t - \bar{x}_t)^2 + \theta \sum_{t=0}^{T-1} [(\bar{x}_{t+1} - \bar{x}_t) - (\bar{x}_t - \bar{x}_{t-1})], \qquad \text{for } \theta > 0. \tag{1}$$

Here we alternately select $\theta = 25, 50,...,125$ to arrive at various decompositions (higher order θs provide smoother trend components (see Kydland and Prescott, 1989). The technique allows for a stochastic "permanent" component while deriving the temporary (stationary) component. To extent that trends in inflation are associated with the general economic trends, using the Hodrick and Prescott method may be more appropriate.[7] Moreover, it is now widely recognized that all decompositions are statistical, and there are infinite number of ways to plausibly decompose a series into permanent and transitory components. An advantage of the Hodrick and Prescott method is that it relies on a minimum number of assumptions and hence is more easily defensible.

The decompositions are employed in the regression:

$$R_t = \alpha_0 + \alpha_1 P(I_t) + \alpha_2 T(I_t) + \varepsilon_t, \tag{2}$$

where R_t represents REIT returns for month t, and $P(\,)$ and $T(\,)$ are the permanent and temporary components of the inflation rate proxy obtained from Equation (1). Preliminary solutions to the regressions provided Durbin-Watson statistics consistent with autocorrelated error terms. Subsequently, the *t*-Statistics from the regressions were adjusted employing the Newey and West (1987) variance estimator for models with autocorrelated disturbances.

Exhibit 2 reports the results from regressing the REIT returns on the permanent and temporary components of the CPI-inflation rate and on the components of the changes

Exhibit 2
Regression Results (1972/02–1995/12)

	AREIT		*EREIT*		*MREIT*		*HYREIT*	
	$\theta = 25$	$\theta = 100$	$\theta = 25$	$\theta = 100$	$\theta = 25$	$\theta = 100$	$\theta = 25$	$\theta = 100$
Panel A: Temporary								
α_0	1.60***	1.49***	1.63***	1.52***	1.48**	1.40**	1.57**	1.44**
	(2.9)	(2.6)	(3.2)	(2.9)	(2.2)	(2.1)	(2.5)	(2.3)
$P(CPI_t)$	−0.14	−0.12	−0.10	−0.08	−0.16	−0.15	−0.13	−0.11
	(−0.2)	(−1.0)	(−0.9)	(−0.7)	(−1.2)	(−1.1)	(−1.0)	(−0.8)
$T(CPI_t)$	−0.02	−0.56	−0.14	−0.16*	−0.01	−0.05	0.10	0.05
	(−0.1)	(−0.5)	(−1.3)	(−1.7)	(−0.1)	(−0.3)	(0.8)	(0.4)
Adj. R^2	.00	.00	.01	.01	.00	.00	.00	.00
Panel B: Permanent								
α_0	0.95	0.79	1.23*	1.07	0.66	0.54	0.93	0.74
	(1.4)	(1.2)	(1.9)	(1.6)	(0.9)	(0.7)	(1.1)	(0.7)
$P(T\text{-}bill_t)$	−0.01	0.01	−0.02	0.01	−0.02	0.00	−0.01	0.01
	(0.2)	(0.1)	(−0.2)	(0.1)	(−0.2)	(<0.1)	(−0.1)	(0.1)
$T(T\text{-}bill_t)$	−0.21	−0.43	−1.03**	−1.05***	0.70	0.31	0.17	−0.18
	(−0.4)	(−0.9)	(−2.4)	(−2.9)	(1.0)	(0.5)	(0.2)	(0.4)
Adj. R^2	.00	.00	.01	.03	.00	.00	.00	.00

The results are from the regression:

$$R_t = \alpha_0 + \alpha_1 P(I_t) + \alpha_2 T(I_t) + \varepsilon_t,$$

where R_t represents REIT returns for month *t*, and $P(\ )$ and $T(\ )$ represent, respectively, the permanent and temporary components of the CPI-inflation rate (or change in the 1-year T-bill rate) obtained from the Hodrick and Prescott model (1980). The *t*-Statistics, in ( ), are adjusted employing the Newey and West (1987) variance estimator for models with autocorrelated disturbances. θ = 25, 100 represent alternate constructions of the permanent and temporary components of inflation (T-bill) rate. Higher levels of θ smooth the (stochastic) permanent component.
*Significant at the .10 level.
**Significant at the .05 level.
***Significant at the .01 level.

in the one-year T-bill rate.[8] For the sake of brevity, we only present results from the regressions that employ $\theta = 25$ and $\theta = 100$ in the Hodrick and Prescott decomposition. The deployment of the alternate decompositions do not change the implications of the results.[9] There is no evidence that REIT returns are positively associated with either the permanent or temporary component of CPI-inflation (Panel A). There is some difference in the results across the definitions of trend. For instance, the results suggest that *EREIT*s are a perverse hedge against temporary inflation (at the .10 level) when employing $\theta = 100$ alone. The *T*(T-bill rate) coefficients in Panel B further indicate that EREITs' returns are negatively related to the temporary

Exhibit 3
Regression Results: Hedged REIT Returns (1976/01–1995/12)

	α_0	$P(CPI_t)$	$T(CPI_t)$	$P(T\text{-}bill_t)$	$T(T\text{-}bill_t)$	Adj. R^2
Panel A: Temporary						
$\theta = 25$	0.76*	0.05	−0.20**			0.02
	(1.9)	(0.7)	(−2.4)			
$\theta = 100$	0.68	0.07	−0.19**			0.02
	(1.6)	(0.8)	(−2.6)			
Panel B: Permanent						
$\theta = 25$	0.76			0.03	−0.67**	0.01
	(1.4)			(0.6)	(−2.6)	
$\theta = 100$	0.70			0.04	−0.60**	0.01
	(1.2)			(0.7)	(−2.6)	

The results are from the regression:

$$H_{R,t} = \alpha_0 + \alpha_1 P(I_t) + \alpha_2 T(I_t) + \varepsilon_t,$$

where $H_{R,t}$ represents the hedged *AREIT* returns for month t, and $P(\ )$ and $T(\ )$ represent, respectively, the permanent and temporary components of the CPI-inflation rate (or change in the 1-year T-bill rate) obtained from the Hodrick and Prescott (1980) model. The hedged REIT returns are obtained from rolling regressions of *AREIT* returns on the S&P 500 Index (see Liang, Chatrath and McIntosh, 1996). The t-Statistics, in ( ), are adjusted employing the Newey and West (1987) variance estimator for models with autocorrelated disturbances. $\theta = 25$, 100 represent alternate constructions of the permanent and temporary components of the inflation proxies. Higher levels of θ smooth the (stochastic) permanent component.
*Significant at the .10 level.
**Significant at the .05 level.

component of inflation. The other REIT indices do not exhibit this behavior. Nonetheless, the regression results in Exhibit 1 fail to provide any indication that REITs are effective in hedging against inflation.

Hedged REIT Returns and Inflation

As indicated earlier, it would be of some interest to investigate whether the stock market component in REITs is to blame for the apparent lack of a positive relationship between REIT returns and inflation. To assess the role of the stock market's influence in the evidence in Exhibit 2, we first construct hedged REIT indices that are purged of the stock market component. We follow the methodology in Liang, Chatrath and McIntosh (1996) in the construction of hedged indices. First, a hedge-ratio series is obtained from the 48-month rolling regression model:

$$r_{R,t} = \alpha_0 + h_t * r_{SP,t} + \varepsilon_t, \tag{3}$$

where $r_{R,t}$ and $r_{SP,t}$ represent total *AREIT* returns and total S&P 500 returns, and ε_t is the regression error term.[10]

The hedged REIT return index is then given by:

$$H_{R,t} = r_{R,t} - h * r_{SP,t}, \tag{4}$$

where $H_{R,t}$ is the hedged REIT return from 1976/01 onward.

Exhibit 3 presents the results from the regressions where the hedged *AREIT* returns are the dependent variable.[11] The results suggest that the hedged REIT returns are negatively associated with the temporary component of CPI-inflation (or the change in T-bill rate). Note that these results in contrast to the insignificant coefficients documented for the *AREIT* regressions in Exhibit 2. In sum, the results continue to indicate that REITs are not effective inflation hedges in the short run. In other words, the stock market component in REITs does not seem to be responsible for the weak evidence regarding the short-term inflation-hedging ability of REITs.

The regression analysis techniques employed to derive the results in Exhibits 2 and 3, however, involve differencing of the data series. Granger and Newbold (1974), and several others since, suggest that differencing implies loss of information in the data, and could seriously bias the conventional tests toward rejecting the null hypothesis of no relation. Unless the difference operator is also applied to the error process, the traditional regression techniques may fail to provide evidence of a relationship when one does exist (also see Johansen and Juselius, 1990).

Cointegration Results

In this section, we provide results from the Johansen tests for cointegration between the REITs indexes and the CPI (and alternately, the one-year T-bill rate). Cointegration itself is a sufficient condition for the existence of a common attractor for a set of series. In other words, cointegrated pairs will not persistently wander (drift) far apart over long intervals, even though they individually own trend components or have long memory processes (*e.g.*, Engle and Granger, 1987). In the context of this study, the finding of cointegration would imply that REITs and general price levels move together when evaluated over long horizons. In other words, cointegration would imply that REITs are effective inflation-hedges in the long run. The Johansen maximum likelihood tests for estimating the number of cointegrating vectors have been shown by simulation studies to be more powerful and robust than alternate cointegration tests (Johansen, 1988). A brief summary of the testing procedures with respect to REITs and inflation is provided here.

Consider the unrestricted, bivariate, near-VAR model for monthly (or quarterly) series on REITs and the CPI:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \Phi D_t + \varepsilon_t, \tag{5}$$

Exhibit 4
Johansen Cointegration Statistics Model Without Linear Trend

	AREIT		*EREIT*		*MREIT*		*HYREIT*	
	M	Q	M	Q	M	Q	M	Q
Panel A: Trace Statistics								
CPI								
$\tau = 0$	31.95***	26.98***	34.01***	31.99***	22.13**	17.57	18.79*	15.37
$\tau \leq 1$	4.47	4.32	4.74	5.52	4.62	4.26	5.13	4.55
T-bill Rate								
$\tau = 0$	6.86	6.55	7.37	6.67	6.71	6.49	6.69	6.83
$\tau \leq 1$	0.50	0.42	0.47	0.56	0.89	0.91	0.81	0.66
Panel B: Maximal Eigenvalues								
CPI								
$\tau = 0$	27.47***	22.66***	29.27***	26.47***	17.50**	13.31	13.66	10.88
$\tau \leq 1$	4.48	4.32	4.74	5.52	4.62	4.26	5.13	4.55
T-bill Rate								
$\tau = 0$	6.36	6.13	6.90	6.10	5.83	5.58	5.88	6.17
$\tau \leq 1$	0.50	0.42	0.47	0.56	0.89	0.91	0.81	0.66

The results (trace tests and maximal eigenvalues) are from the Johansen and Juselius (1990) procedures for cointegrating regressions for $k = 4$. The statistics presented are for bivariate regression (REITs versus CPI, or REITs versus the 1-year T-bill rate). $\tau = 0$ ($\tau \leq 1$) represents the trace statistic for at most 1 (2) cointegrating vector(s). Critical values were obtained from Johansen and Juselius (1990, Table A3: 209).
*Significant at the .10 level.
**Significant at the .05 level.
***Significant at the .01 level.

where at t, X is the (2×1) vector of the log of (REITs, CPI), D is a centered seasonal dummy variable, μ, is a (2×1) vector of constants and ε is a (2×1) vector of white noise. Johansen and Juselius (1990) show that the coefficient (2×2) matrix Π contains the essential information about the long-term relationship between the two dependent variables.[12] Specifically, the number of equilibrating relationships is equal to the rank of Π: if (1) Π is full rank (rank = 2 in our model); (2) the vector X_t is stationary so that there are as many cointegrating vectors as there are dependent variables; (3) rank(Π) = 1, then X_t is only partially stationary, but the dependent variables *are* cointegrated; (4) rank(Π) = 0, *i.e.*, Π is a null matrix, X_t is not stationary, the dependent variables are not cointegrated; and (5) Equation (5) corresponds to a traditional differenced vector time series model with no error correction mechanism.

A cointegrating relationship between REITs and the CPI from the Johansen procedures will imply that there are $2 \times r$ matrices α and β, such that $\Pi = \alpha\beta'$ where rank(Π) = r. The matrix β represents the cointegrating vectors and α is a matrix of weights. Therefore, the cointegration model shown in Equation (5) produces a set of nonlinear

Exhibit 5
Johansen Cointegration Statistics Model With Linear Trend

	AREIT		*EREIT*		*MREIT*		*HYREIT*	
	M	Q	M	Q	M	Q	M	Q
Panel A: Trace Statistics								
CPI								
$\tau = 0$	15.92*	19.58**	12.42	17.44*	16.31*	19.82**	14.01	19.38**
$\tau \leq 1$	0.55	1.77	0.45	1.53	0.44	1.85	0.54	2.29
T-bill Rate								
$\tau = 0$	12.66	15.78*	11.09	14.67	13.56	16.29*	12.01	15.21
$\tau \leq 1$	1.53	3.12	1.09	2.30	1.54	3.66	1.66	3.15
Panel B: Maximal Eigenvalues								
CPI								
$\tau = 0$	16.47**	14.49*	22.12***	14.24*	19.91***	15.43**	16.23**	15.39**
$\tau \leq 1$	3.02	1.15	4.74	3.46	3.26	0.60	3.63	2.65
T-bill Rate								
$\tau = 0$	9.96	7.61	10.64	7.23	11.18	7.50	10.38	7.91
$\tau \leq 1$	1.67	1.31	2.04	3.08	0.56	0.22	0.80	0.99

The results (trace tests and maximal eigenvalues) are from the Johansen and Juselius (1990) procedures for cointegrating regressions for $k = 4$. The statistics presented are for bivariate regression (REITs versus CPI, or REITs versus the 1-year T-bill rate). $\tau = 0$ ($\tau \leq 1$) represents the trace statistic for at most 1 (2) cointegrating vector(s). Critical values were obtained from Johansen and Juselius (1990, Table A2: 208).
*Significant at the .10 level.
**Significant at the .05 level.
***Significant at the .01 level.

cross equation contraints on the parameters of the unrestricted model. Tests for parameter restrictions on β are accomplished by forming likelihood ratio statistics from the estimation of restricted and unrestricted models. These statistics, trace statistics and the maximal eigenvalues, are distributed chi-square with degrees of freedom equal to the number of parameter restrictions in the null hypothesis. Johansen and Juselius (1990) indicate the potential sensitivity of these test statistics to the assumption of a linear trend in the cointegrating regression.[13]

The Johansen tests are reported in Exhibit 4 (model without trend) and Exhibit 5 (model with trend). The residuals from Equation (5) are consistently found to be void of autocorrelation for $k = 4$ and this lag structure is employed throughout. The statistics corresponding to the null, $\tau = 0$, represents a test for the alternate of at least one cointegrating vector. Similarly, the statistic corresponding to the null, $\tau \leq 1$, represents a test for two cointegrating vectors. The tests for $\tau = 0$ and $\tau \leq 1$ are ordered and thus are highly dependent on one another.

Exhibit 6
Summary of Results

	AREIT	*EREIT*	*MREIT*	*HYREIT*
Panel A: Regression Results				
Evidence on the ability of REITs to hedge the temporary component of inflation	No	No	No	No
Evidence on the ability of REITs to hedge the permanent component of inflation	No	No	No	No
Panel B: Cointegration Results				
Evidence of cointegration from Dickey-Fuller tests employing CPI	No	No	No	No
Evidence of cointegration from Dickey-Fuller tests employing T-bill rate	No	No	No	No
Evidence of cointegration from Johansen tests employing CPI	Yes	Yes	Weak	Weak
Evidence of cointegration from Johansen tests employing T-bill rate	No	No	No	No

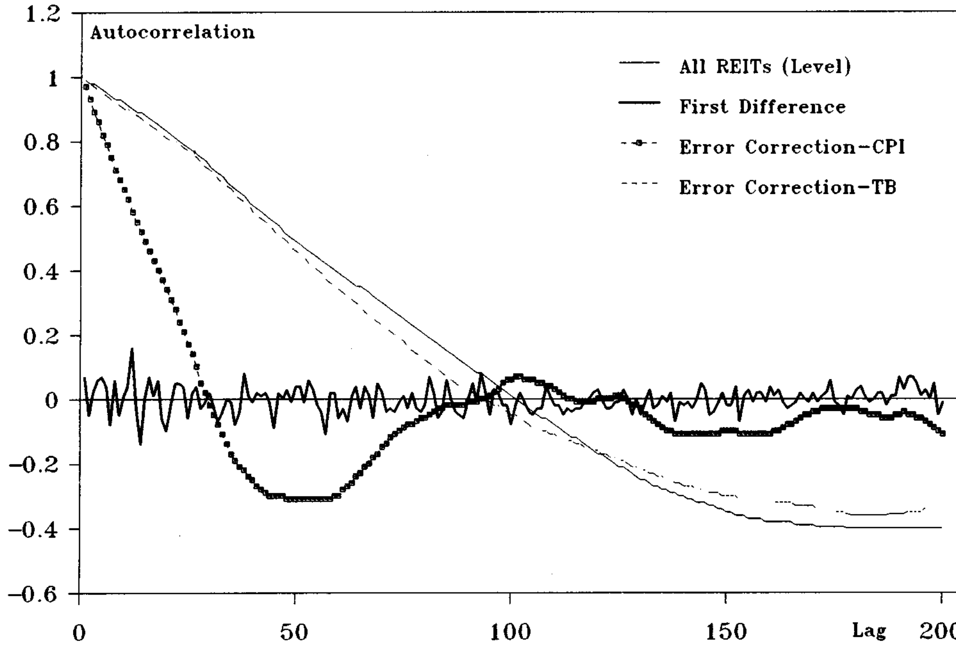
In Panel A of Exhibit 4, the trace and maximal eigenvalue statistics consistently suggest one cointegration vector between *AREIT*s and the CPI, and between *EREIT*s and the CPI. For instance, for the *AREIT*-CPI pair in Panel A, the trace statistic for at most one cointegrating vector is 31.95, and the 99% critical value given in Johansen and Juselius (1990, Table A3) is 24.99. The order of cointegration of one simplifies the interpretation of the cointegration vectors as a long run relationship in the levels of the processes (Johansen and Juselius, 1990:206).

However, the results are weaker for the other REIT series. For instance, the statistics for the quarterly *MREIT* and *HYREIT* series with quarterly seasonal dummies are not significant at the .10 level. Moreover, the tests fail to provide any evidence of a long-run relationship between the REIT indices and the one year T-bill rate. In Panel B, the trace statistics and the maximal eigenvalues fail to reject the null of at most one cointegrating vectors.

Exhibit 5 reports results from the Johansen procedures from the model with trend. The test statistics and their significance levels are conspicuously different from those in Exhibit 4. For instance, in contrast to the results in Exhibit 4, the trace statistic for the monthly *AREIT*s-CPI series is barely statistically significant, and that for the *EREIT*-CPI series is not significant. Interestingly, Exhibit 5 provides some evidence of cointegration between REITs and the one-year T-bill rate. In sum, the cointegration results are fairly sensitive to the trend term in Equation (5).

While the Johansen tests provide some encouraging results (of cointegration), it should be noted that the more standard, residual-based Dickey-Fuller and Phillips tests for

Exhibit 7
Correlogram for the AREITs Index and REITs-CPI, REITS-Treasury Bill Rate Cointegrating Vector (1972/01–1995/12)



The autocorrelations are for the *AREITs* index, the first difference of the *AREITs* index, and for the error correction terms $z_{1,t}$ and $z_{2,t}$, obtained from OLS regressions of the monthly levels of the *AREITs* index on a constant and monthly levels of the CPI (one-year Treasury bill rate).

cointegration failed to provide any indication of cointegration between either the REITs and CPI, or between the REITs and the one year T-bill rate. For the sake of brevity we do not present these test results.[14] However, to highlight the differences across the procedures, we provide in Exhibit 6 a summary of the evidence from the regression and cointegration tests presented so far.

The inconsistencies among the cointegration results do raise other questions regarding the degree of comovements between the REIT and CPI series. The residual based tests for cointegration (Dickey-Fuller and Phillips tests) are predicated on the assumption that the cointegrating vector is I(0). Among others, Baillie and Bollerslev (1994) consider the possibility of a form of cointegration that exists with the cointegrating vector being I(d), $d < 1$ (fractional cointegration).[15] Such a form of cointegration may possess very long memory, according to which autocorrelations will display long term cycles, inconsistent with a strict I(1) series which exhibit slow, monotonic decay. As we find mixed evidence of cointegration, we consider the possibility of fractional cointegration.

We plot the autocorrelations for error correction terms, $z_{1,t}$ = (AREITs $index_t - \beta$(CPI index$_t$)) and $z_{2,t}$ = (*AREIT*s index $- \beta$(T-bill Rate$_t$)), where $z_{1,t}$ ($z_{2,t}$) is obtained from the monthly ordinary least squares regression of the *AREIT*s index on a constant term and the CPI (one-year T-bill rate). These plots are presented in Exhibit 7. For comparative purposes, we also plot the functions for the *AREIT* series and for the first difference in the *AREIT* series.

The autocorrelation function for the *AREIT* series has a slow decay process that is typical of an I(1) series. On the other hand, the first differenced function is obviously I(0). The $z_{1,t}$ autocorrelations has a slow decay relative to the function representing the first differenced series. However, these autocorrelations exhibit long-term cycles, similar to function demonstrated by Baillie and Bollerslev (1994) in their study of exchange rates. The $z_{1,t}$ function is also in marked contrast to the I(1) *AREIT* series, making a case for further investigation into the possibility of fractional cointegration between REITs and general price levels. The $z_{2,t}$ function makes a weaker case for cointegration or fractional cointegration between the *AREIT* index and T-bill rate, consistent with the results in Exhibits 4 and 5.

Conclusion

This article examines the inflation-hedging characteristics of REITs over the 1972/01–1995/12 interval. Regression estimations in the more traditional mold fail to provide evidence that REITs are effective inflation hedges. We also rule out the possibility that a stock-market induced ‘proxy effect’ causes the apparent lack of regression relationship between REITs and inflation.

We find some evidence of a long-run relationship between the CPI and alternate REIT indices when employing Johansen cointegration tests. However, while the Johansen tests provide some encouraging results with respect to the long-run relationship between the CPI and the REIT indices, the overall evidence of cointegration between REITs and inflation is tenuous: very weak evidence of cointegration is noted between REITs and the T-bill rate when employing the Johansen procedures, and more traditional tests for cointegration (*e.g.*, Phillips and Dickey-Fuller) fail to indicate any evidence of cointegration between REITs and inflation, however measured. We make a case for an investigation into the possibility of fractional cointegration between REITs and general price levels. Fractional cointegration would imply that REITs provide only a partial hedge of inflation in the long run.

It is widely recognized that any asset class in a portfolio should not provide a negative real return over the long run. However, portfolio managers seeking to hedge inflation risk must seek out assets that not only provide positive real returns on average, but also those that are positively related to general price levels. Convincing evidence of a long-term general price component in the REIT indices would thus have had important implications to such portfolio managers. The implications of a weaker form of cointegration between real estate and general prices remains to be examined.

Notes

[1] For a review on the evidence of the REITs-inflation relationship and the models employed to examine this relationship, see Yobbaccio, Rubens and Ketcham (1995).

[2] The authors also provide an excellent review of literature on the inflation-hedging capacity of commercial real estate.

[3] It should be noted, however, that the inflation hedging effectiveness evidenced in the group of studies that employ unsecuritized real estate data has often been attributed to appraisal-smoothed biases common to such data series (Geltner, 1989; and Giaccotto and Clapp, 1992).

[4] Monthly REIT indices are from the 1996 *NAREIT Fact Book,* Washington, DC. Other data series are from the Pinnacle Data Corporation, Webster, NY.

[5] The Phillips-Perron test statistics are obtained from employing the equations (without trend and with trend (t)):

$$\Delta Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \varepsilon_t,$$

$$\Delta Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \alpha_2 t + \varepsilon_t.$$

The test statistics (on α_1) are transformed to remove the effects of serial correlation on the statistics' asymptotic distribution as proposed by Phillips and Perron (1986) and Perron (1988: 308–9). The critical values are: with trend: −3.13 (10%), −3.41 (5%), −3.96 (1%), without trend: −2.57 (10%), −2.86 (5%), −3.43 (1%).

[6] For a review of the concepts of integration/cointegration, see Engle and Granger (1987).

[7] See, for example, Kydland and Prescott (1990) and Backus and Kehoe (1992).

[8] The deployment of the three-month T-bill rate provided very similar results.

[9] The results from the alternate decompositions are available from the author.

[10] To clarify, each regression of forty-eight sets of observations produced a regression coefficient that was employed as the hedge ratio pertaining to the last month in that subsample.

[11] Similar results are obtained for the other REITs in the sample.

[12] Johansen and Juselius term Π the 'long-run impact matrix.'

[13] The sensitivity of the Johansen procedures across models with and without trend has also been noted by Baillie and Bollerslev (1994) and Diebold, Gardeazabal and Yilmaz (1994).

[14] The results from the Phillips and Dickey-Fuller cointegration techniques are available from the author.

[15] Standard unit root tests (*e.g.*, Phillips-Perron) are known to have very low power against fractional alternatives (Diebold and Rudebusch, 1991).

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