E C O N O M I C S B U L L E T I N

An Empirical Note on Testing the Cointegration Relationship Between the Real Estate and Stock Markets in Taiwan

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Abstract

This note studies the long-run relationship between real estate and stock markets in the Taiwan context over the 1986Q3 to 2006Q4 period, using standard cointegration test of Johansen and Juselius (1990) and that of Engle-Granger (1987) as well as the fractional cointegration test of Geweke and Porter-Hudak (1983). The results from both types of cointegration tests strongly indicate that these two markets are not cointegrated with each other. With respect to risk diversification, it is obvious that investors and financial institutions should have included both assets in the same portfolio during that period.

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I. INTRODUCTION

To portfolio investors who want to diversify in the real estate and stock markets, having a full understanding the long-run relationship between these two markets is central. It is quite apparent, after all, that if the two markets have a long-run relationship, then jointly holding such assets in the same portfolio would likely offer very few gains in terms of risk reduction.

Previous empirical studies have employed cointegration techniques to investigate whether there exist such long-run benefits from international equity diversification (to name a few, see Kwan et al., 1995; Masih and Masih, 1997). Yet, exactly what have they empirically shown? According to both of these studies, asset prices from two different efficient markets cannot be cointegrated. To be more precise, the studies report that if a pair of asset prices is cointegrated, then one asset price can be forecast (i.e., it is Granger-caused) by the other asset price. Thus, these cointegration results indicate that, with regard to reducing risk, certainly few-if any- gains are obtained from such portfolio diversification.

In revisiting this issue, this note contributes to this line of research by exploring whether there are any long-run benefits from asset diversification for those who invest in Taiwan's real estate and stock markets. Unlike other studies, here we test for cointegration using both the standard cointegration test of Johansen and Juselius (1990) and that of Engle-Granger (1987) as well as the fractional cointegration test of Geweke and Porter-Hudak (1983). With the results from the three tests combined, we determine that these two asset markets are not, in fact, pair-wise cointegrated with each other. The finding of no cointegration can be interpreted as clear-cut evidence that there are no long-run linkages between these two asset markets and that, therefore, potential gains are actually present for investors who diversify in these two asset markets over this sample period. These results are invaluable for investors and financial institutions holding long-run investment portfolios in these two asset markets.

. The remainder of this note is organized as follows. Section II presents a review of some previous literature. Section III presents the data used. Section IV presents the methodologies used and discusses the findings. Finally, Section V concludes.

II. Review of the Literature

Identifying the relationship between real estate prices and stock prices has been a widely debated issue within academic circles and among practitioners, alike. Although current literature on the relationship between real estate and equity markets tends to show conflicting results, most of the empirical evidence seems to support the view that the two markets are segmented. Goodman (1981), Miles et al., (1990), Liu et al., (1990) and Geltner (1990), for example, argue for the existence of such segmentation within various real estate markets and stock markets. In direct contrast, Liu and Mei (1992), Ambrose et al. (1992) along with Gyourko and Keim (1992), report results that contradict that position, claiming that real estate and stock markets are in fact integrated. The predicament faced here, therefore, is whether the two markets are segmented or integrated. Our primary objective then is to ascertain whether any significant relationship does exist between these markets and, if so, to determine what implications it may have for active market traders. One fundamental motivation behind our study is that our findings can yield considerable insight for both investors and speculators that may facilitate forecasting future performance from one market to the other.

III. DATA

The data sets used here consist of quarterly time series on the real estate price index (lresp) and stock price index (lstkp) for the 1986Q3 to 2006Q4 period. To avoid the omission bias, we also incorporate real interest rate (liret) into our study. The real interest rate and stock price indexes are obtained from the *AREMOS database* of the Ministry of Education of Taiwan. The real estate price index is collected and compiled by Hsin-Yi Real Estate Inc. An examination of the individual data series makes it clear that logarithmic transformations are required to achieve stationarity in variance; therefore, all the data series are transformed to logarithmic form.

Descriptive statistics for both real estate and stock markets returns are reported in Table 1. We find that the sample means of the real estate price returns are positive (1.12%), whereas the stock price returns are negative (-0.091%). Both the skewness and kurtosis statistics indicate that the distributions of the returns of both markets are normal. The Jung-Box statistics for 4 lags applied to the returns and square returns indicate that no significant linear or non-linear dependencies exist in either market.

IV METHODOLOGY AND EMPIRICAL RESULTS

A. Unit Root Tests

Previous studies point out that the standard ADF test is not appropriate for variables that may have undergone structural changes. Perron (1989, 1990), for instance, shows that the presence of structural changes biases the standard ADF test towards nonrejection of the null of a unit root. Hence, it might very well be misleading to conclude that variables are nonstationary merely on the basis of results from a standard ADF test. Perron (1990) develops a procedure to test the hypothesis that a given series $\{Y_t\}$ has a unit root with an exogenous structural break, which occurs at time T_B. Zivot and Andrews (1992, hereafter ZA) disagree with this assumption of an exogenous break point and develop a unit root test procedure that allows an estimated break in the trend function under the alternative hypothesis. For this reason, in this study, it seems most reasonable to treat the structural break as endogenous and test the order of integration using the ZA procedure. ZA tests are represented by the following augmented regression equations:

Model A: $\Delta Y_t = \mu_1^A + \beta_1^A t + \mu_2^A DU_t + \alpha^A Y_{t-1} + \sum_{j=1}^{k}\theta_j \Delta Y_{t-j} + \varepsilon_t$;

Model B: $\Delta Y_t = \mu_1^B + \beta_1^B t + \gamma^B DT_t^* + \alpha^B Y_{t-1} + \sum_{j=1}^{k}\theta_j \Delta Y_{t-j} + \varepsilon_t$; and

Model C: $\Delta Y_t = \mu_1^C + \beta_1^C t + \mu_2^C DU_t + \gamma^C DT_t^* + \alpha^C Y_{t-1} + \sum_{j=1}^{k}\theta_j \Delta Y_{t-j} + \varepsilon_t$, (1)

where $DU_t = 1$ and $DT_t^* = t - T_B$ if $t > T_B$, and 0 otherwise. Here T_B refers to a possible break point. Model A allows for a change in the level of the series, Model B permits a change in the slope of the trend function, while Model C combines changes in the level and slope of the trend function of the series. The sequential ADF test procedure estimates a regression equation for every possible break point within the sample and calculates the t-statistics for the estimated coefficients. This tests the null hypothesis of a unit root against the alternative hypothesis of a trend stationarity with a one-time break (T_B) in the intercept and slope of the trend function at an unknown point in time. The null of a unit root is rejected if the coefficient of Y_{t-1} is significantly different from zero. The selected break point for each series is that particular T_B for which the t-statistic for the null is minimized.

Since the choice of lag length k may affect the test results, the lag length is selected in accordance with the procedure of Campbell and Perron (1991). Start with an upper bound $k_{\max}$ for k. If the last included lag is significant, then choose k = $k_{\max}$. If not, reduce k by 1 until the last lag becomes significant. We set $k_{\max}$ = 4 for our quarterly data series.

For comparison, we also incorporate the standard ADF and KPSS (Kwiatkowski et al., 1992) tests into our study. Panels A and B in Table 2 present the results of the non-stationary tests for the stock price index (lstkp), real estate price index (lresp), and real interest rate (liret) from the ADF and KPSS tests. We find each data series is nonstationary in levels but stationary in first differences, suggesting that all the data series are integrated of order one.

Table 3 reports the minimum t-statistics that correspond to Model C. The test results summarized in Table 3 support the existence of a unit root when breaks are allowed. The test results are identical to those from the standard ADF and KPSS tests reported in Table 2, again implying that all the data series are integrated of order one, I(1), even when breaks are allowed. The plausible breaks for the series occur at 1998Q1, 1997Q4, and 1997Q2, respectively, for the stock price index, real estate price index, and real interest rate. On the basis of these results, we proceed to test whether these three variables are cointegrated, and to do so, we used three cointegration tests.

B. Cointegration Tests

1. The Johansen Method

Following Johansen and Juselius (1990), we construct a p-dimensional (3 x 1) vector autoregressive model with Gaussian errors, which can be expressed by its first-differenced error correction form as:

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + \Gamma_2 \Delta Y_{t-2} + ... + \Gamma_{k-1} \Delta Y_{t-k+1} - \Pi Y_{t-1} + \mu + \varepsilon_t , \quad (2)$$

where Y_t are the data series studied; ε_t is i.i.d. N(0, Σ), $\Gamma_i = -I + A_1 + A_2 + ... + A_i$, for i=1,2,...,k-1; and $\Pi = I - A_1 - A_2 - ... - A_k$. The Π matrix conveys information about the long-run relationship between the Y_t variables, and the rank of Π is the number of linearly independent and stationary linear combinations of the variables studied. Thus, testing for cointegration involves testing for the rank of the Π matrix r by examining whether the eigenvalues of Π are significantly different from zero.

Johansen and Juselius (1990) propose two test statistics for testing the number of cointegrating vectors (or the rank of Π): the trace (T_r) and the maximum eigenvalue (L-max) statistics. The likelihood ratio statistic for the trace test is:

$$-2 \ln Q = -T \sum_{i=r+1}^{p=3} \ln(1 - \hat{\lambda}_i), \quad (3)$$

where $\hat{\lambda}_{r+1} ..., \hat{\lambda}_p$ are the estimated p-r smallest eigenvalues.

The null hypothesis is that there are at most r cointegrating vectors. That is, the number of cointegrating vectors is fewer than or equal to r, where r is 0, 1, or 2. In each case, the null hypothesis is tested against the general alternative.

Alternatively, the L-max statistic is:

$$-2 \ln Q = -T \ln(1 - \hat{\lambda}_{r+1}) \quad (4)$$

In this test, the null hypothesis of the r cointegrating vectors is tested against the alternative of the r+1 cointegrating vectors. Hence, the null hypothesis r=0 is tested against the alternative that r=1; r=1 is tested against the alternative r=2; and so forth.

It is well known that Johansen's cointegration tests are very sensitive to the choice of lag length. As a result, the number of lags required in the cointegration test is determined on the basis of the Schwartz Information Criterion (SIC).[1] A VAR model is first fit to the data to find an appropriate lag structure and based on the SIC, 1 lag seems the most justified for our VAR model. Table 4 presents the results from the Johansen and Juselius (1990) cointegration test. According to Cheung and Lai (1993a), the Trace test shows more robustness to both skewness and excess kurtosis in the residuals than does the L-max test; therefore, we only used Trace statistics in our study. As shown in Table 4, the Trace statistics indicates that the null hypothesis of no cointegration cannot be rejected. Clearly then, the results demonstrate that no long-run cointegrating relationship exists between real estate and stock markets over this sample period.[2]

2. The Engle-Granger Cointegration Test

The Engle-Granger cointegration test consists of a two-stage procedure. By design, cointegration tests using this approach are extremely sensitive in regard to which variable in the cointegration equation is used for normalizing in the cointegration equation; thus, we perform tests based on both normalizations using lstkp and lresp as dependent variables in the cointegration equations. Following Engle and Granger (1987), in the first stage, we estimate the cointegration equations as follows:

$$Y_{1t} = \alpha + \beta Y_{2t} + \gamma X_t + \varepsilon_{1t};$$
$$Y_{2t} = a + bY_{1t} + \kappa X_t + \varepsilon_{2t} \qquad (5)$$

Secondly, we implement the ADF test on residuals ε_{1t} and ε_{2t}. The Engle-Granger test results are reported in Panel A of Table 5, and it is most evident that the results of the ADF test never reject the null of no cointegration, even at the 10% significance level.

3. The Fractional Cointegration Test

Once rejection of cointegration is established by the above two widely-used testing procedures, we conduct the fractional cointegration test using the residuals of the cointegration equations. The implementation of this technique follows the procedures reported in a similar study that uses the standard cointegration technique of Cheung and Lai (1993b). We first perform the cointegration regression of equation (5) and then check the residuals for fractional integration. In this stage, the fractional integration test on the residual series requires an estimation of parameter *d* along with a test of its significance. In this paper, parameter *d* is our primary interest because evidence of fractional integration (*i.e.* $0 < d < 1$) in the error series confirms a

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[1] Using Monte Carlo simulations, Cheung and Lai (1993a) showed that for autoregressive processes standard selection criteria, like the Schwartz Information Criterion (SIC) and Akaike Information Criterion (AIC), can be useful for selecting the correct lag structure for Johansen's cointegration test. They find that the SIC performs slightly better than the AIC.

[2] In order to further verify our results, we also employ Gregory and Hansen's (1996, Residual-based tests for cointegration in models with regime shifts, Journal of Econometrics, 70, 99-126) method to test the cointegrating relationship among these three variables, and the results are similar to those in our study indicating there exists no long-term equilibrium relationship between real estate and stock markets. Due to space constraints, we do not report Gregory and Hansen's testing results in our paper. Those results are available upon request.

long-run relationship between the two series. That is, a value of *d* between zero and one in the residual series from the cointegration regression provides evidence of a long-run equilibrium relationship.

Since the periodogram is used as an estimator of the spectral density, *d* may be approximated by regression. Cheung and Lai (1993b) show that, in the case of fractional cointegration, the OLS estimate is also consistent. In the present study, the estimation is calculated using the spectral regression technique developed by Geweke and Porter-Hudak (GPH) (1983).[3] The GPH estimation procedure relies on an OLS regression on:

$$\ln[I(W_j)] = \mu - d\ln[4\sin^2(W_j/2] + \eta_j, \quad \forall_j = 1,2,...n \tag{6}$$

for $W = 2\pi j/T, (\forall j = 1,2,...,T-1), n = g(T) < T$, where $I(W_j)$ is the periodogram of Z at frequency W_j defined by:

$$I(W) = \frac{1}{2\pi T}\left|\sum_{t=1}^{T} e^{itW}(Z_t - \bar{Z})\right|^2 \tag{7}$$

Since the periodogram is used as an estimator of the spectral density, a choice of a truncation parameter must be made for the number of low Fourier frequencies, n, to be used in the spectral regression. Geweke and Porter-Hudak (1983) recommend using $n = T^{0.5}$, where T is the number of observations. To avoid making subjective choices that may be either too high or too low, we follow Cheung and Lai (1993b) and choose a range of values with the power of ranging from 0.5 to 0.6. The estimated results of fractional cointegration by d(GPH) are reported in Panel B of Table 5. In each case of the varying power parameters, the hypothesis tests of *d* = 1 are not rejected at least at the 10% level, but those of *d* = 0 are rejected at least at the same 10% level. These results do not support any possible fractional cointegration between the real estate and stock markets. They further confirm that no long-run relationship is found between these two markets, thereby duplicating our conclusions from the standard cointegration tests of Johansen and Juselius (1990) and Engle and Granger (1987). The lack of a long-run relationship suggests that, in terms of risk reduction, there are indeed long-run benefits from jointly holding these two assets in a portfolio.

V. CONCLUSIONS

In this note we study the relationship between the real estate and stock markets in the Taiwan context over the 1986Q3 to 2006Q4 period, using both the standard cointegration tests of Johansen and Juselius (1990) and Engle and Granger (1987) along with the fractional cointegration test of Geweke and Porter-Hudak (1983). The results from all three cointegration tests conclusively show that these two markets are not cointegrated with each other. In terms of risk diversification, the two assets should have been included in the same portfolio in Taiwan during the 1986Q3 to 2006Q4 period. Consequently, these findings ought to be made readily available to individual investors and financial institutions holding long-term investment portfolios in these two asset markets for their likely implications today.

Of great value will be our future research, which will examine the robustness of the results of the present study by modifying the models of the linear cointegration

[3] For other methods, see Robinson's (1995) semi-parametric procedure (based on the frequency-domain) and Sowell's (1992) maximum likelihood procedure (based on the time-domain).

test along the lines of Okunev and Wilson (1997) and by applying this non-linear cointegration test to our data.

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Table 1. Summary Statistics of Real Estate and Stock Markets Returns

	Real Estate Price Returns	Stock Price Returns
Mean	1.12%	-0.091%
SD	17.55%	4.751%
Maximum	56.84%	5.766%
Minimum	-25.497%	-5.121%
Skewness	0.826	0.121
Kurtosis	4.599	2.973
J-B N Test	3.247	1.693
L-B (Q=4)	5.589	3.114
L-B (Q=4)-Square	6.568	5.537

Note: SD denotes standard error. The standard errors of the skewness and kurtosis are $(6/T)^{0.5}$ and $(24/T)^{0.5}$, respectively. J-B N Test denotes the Jarque-Bera normality test. L-B (Q=k) represents the Ljung-Box test for autocorrelation up to k lags. .

Table 2. ADF and KPSS Unit Root Tests

	Panel A: ADF		Panel B: KPSS (η_{μ})	
	level	difference	level	difference
lstkp	-1.132 (1)	-4.139* (1)	1.121* [1]	0.198 [1]
lresp	-0.798 (1)	-5.781* (1)	1.186* [1]	0.051 [1]
liret	-1.585 (1)	-5.122* (1)	0.921* [1]	0.036 [1]

Note: 1. Number in parentheses indicates the selected lag order of the ADF model. Lags are chosen based on Campbell and Perron’s (1991) method.
2. * indicates significance at the 5% level.
3. Critical values for the KPSS test are taken from Kwiatkowski et al. (1992).

Table 3. Zivot-Andrews Unit Root Tests for One Break

	Model	Break	$t(\hat{\lambda}_{\inf})$
lstkp	C	1998Q1	-2.212
lresp	C	1997Q4	-3.135
liret	C	1997Q2	-3.422

Note: 1.Model specification, (i.e., which model, A, B, or C, is most appropriate) is determined by first running each data series on Model C, with the possibility of both a slope and a level break. Model C is chosen if both dummy variables are significant. If only the slope dummy variable is significant, Model B is estimated. If the level dummy is significant, Model A is estimated.

2. Critical values are taken from Zivot and Andrew (1992). The 10% and 5% critical values are -4.58 and -4.80, respectively, for Model A, -4.11 and -4.42, respectively, for Model B, and -4.82 and -5.08, respectively, for Model C.

Table 4. Cointegration Tests Based on the Johansen and Juselius (1990) Approach (VAR lag = 1)

	Trace test	5% critical value	10% critical value
$H_0: r = 0$	24.89	29.68	26.79
$H_0: r \leq 1$	7.99	15.41	13.33
$H_0: r \leq 2$	0.56	3.76	2.69

Note: 1. Critical values are taken from Osterwald-Lenum (1992).

2. r denotes the number of cointegrating vectors.

Table 5. Engle-Granger Cointegration and GPH Fractional Cointegration Tests

Panel A. Engle-Granger Cointegration Test

Dependent Variable	Constant	lstkp	lresp	liret	R-squared	D-W
lstkp	17.211		-3.334	-0.261	0.403	0.652
lresp	6.221	-0.195		0.143	0.423	0.618

Normalized	ADF	Critical values
lstkp	-2.433	-3.649(1%), -2.95(5%), -2.616(10%)
lresp	-2.672	

Panel B. GPH Fractional Cointegration Test

	Residuals normalized on lstkp			Residuals normalized on lresp		
Power order		Null hypothesis			Null hypothesis	
λ	d(GPH)	d = 1	d = 0	d(GPH)	d = 1	d = 0
0.50	0.864	-0.283	1.856*	1.082	0.158	2.339*
0.55	0.787	-0.477	1.868*	1.466	1.047	3.228*
0. 60	0.851	-0.415	2.463*	1.089	0.241	3.024*

Note: λ denotes the power associated with the Fourier transforms ($n = T^{\lambda}$). The null hypothesis of d = 1 is tested against a one-sided alternative of d < 1; d = 0 is tested against a two-sided alternative of $d \neq 0$. The * denotes significance at the 10% level.