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A Note on Testing for the Periodically Collapsing Bubbles in Japanese REIT Markets

Shyh-Wei Chen,[a*] An-Chi Wu[b]

a. Department of International Business, Tunghai University
b. Department of Money and Banking, National Chengchi University.

A B S T R A C T

This simple note tests for the presence of Evans' (1991) periodically collapsing bubbles of three real estate investment trust (REIT) classifications in Japan by employing the momentum threshold autoregressive (MTAR) model and the MTAR model with smooth transition in trend (i.e., the LNV-MTAR model). The results of the conventional linear unit root test show evidence of rational bubbles in Japanese REIT markets. However, the results of the MTAR and LNV-MTAR test show that periodically collapsing bubbles do not hold in Japan REIT markets. An important implication of this study is that if we neglect the nonlinear properties inherent in the data, then we are inclined to wrongly agree with the existence of speculative bubble based only on the conventional linear approaches.

Keywords: Present value model, Periodically collapsing bubble, Unit root, MTAR

JEL classification: G12, C22

* Corresponding author.
E-mail addresses: shyhwei.chen@gmail.com

1. Introduction

Academic researchers like economists have shown considerable interest in the present value model of stock prices which argues that the stock price is the present discounted value of the future expected dividend (e.g., Campbell and Shiller, 1987; Campbell et al., 1997; Cochran, 2001). However, asset prices that are in excess of what is viewed as the asset's fundamental value have been interpreted as speculative bubbles. A class of speculative bubbles known as rational bubbles, do not violate the rational expectations hypothesis and are consistent with the efficient markets hypothesis.

Theoretically, non-fundamental stock price increases and crashes can be integrated into present value models by dropping the transversality condition which imposes a unique solution on the stock prices. Empirical studies, however, have often reported substantial deviations between actual stock prices and theoretical stock prices derived from the linear present value model. For example, many studies find that U.S. stock prices are more volatile than those determined by the present value model. A number of factors have been put forth to account for this substantial deviation, including stochastic speculative bubbles (Blanchard and Watson, 1982; Evans, 1991; West, 1987); noise traders models (Kirman, 1991, 1993; Shleifer, 2000), fads (Shiller, 1981), varying discount rates (Campbell and Shiller, 1988a, 1988b), and the intrinsic bubble (Driffill and Sola, 1998).[1]

The extant empirical evidence of the existence of the rational bubble (e.g., Campbell and Shiller, 1987; Campbell et al., 1997) has been extensively presented in the unit root and cointegration framework. The cointegration test examines the relationship between securities prices and the vector of fundamental factors over the long term (see Bohl, 2003; Brooks and Katsaris, 2003; McMillan, 2007). Those reporting unit root behaviors in the price-dividend relationship, which in turn provides implicit support for the rational bubbles hypothesis (e.g. see Froot and Obstfeld, 1991; Balke and Wohar, 2002; Bohl and Siklos, 2004), and those arguing that the price-dividend ratio exhibits fractional integration such that while it is characterized by long memory, the series is ultimately mean reverting (Caporale and Gil-Alana, 2004; Cunado et al., 2005; Koustas and Serletis, 2005).

The periodically collapsing bubble is one kind of the rational bubble. Traditionally, from the viewpoint of econometrics, if the residuals of the regression of securities prices on any set of fundamentals are stationary, I(0), then this can be regarded as evidence against the existence of a bubble. In addition, if securities prices and fundamental factors exhibit a long-run relationship as evidenced by any number of cointegrating vectors, they serve as evidence against the existence of bubbles in the securities prices. However, Evans (1991) argues that this standard approach will not be able to detect a class of periodically collapsing rational bubbles. For example, the sudden collapse of a bubble may be mistaken by standard cointegration tests for mean reversion, resulting in a bias towards the rejection of the null hypothesis of no cointegration. Intuitively, if the stock price exhibits the phenomenon of the periodically collapsing bubble, then the stock price will increase without bound and then collapses. Next, it rebounds to the peak and collapses again. Hence, the stock price is looking like a stationary process rather than a random walk process.

Many researchers, for instance, Payne and Waters (2005, 2007), Jirasakuldech et al. (2006),

[1] Camerer (1989) considers the possibility that asset prices might deviate from intrinsic values based on market fundamentals. Three broad categories of theory are surveyed in Camerer's article, including of (a) growing bubbles (b) fads and (c) information bubbles. Redaers are referred to his paper for details.

Hui, and Yue (2006), Waters and Payne (2007), Anderson, Brooks and Tsolacos (2011), Paskelian, Hassan and Whittaker (2011), Paskelian and Vishwakarma (2011), Zhou and Anderson (2013), Vishwakarma (2012, 2013), Brauers, Thomas and Zietz (2014), Xie and Chen (2015), Engsted, Hviid and Pedersen (2016) and Escobari and Jafarinejad (2016) have devoted their efforts to test for the periodically collapsing bubbles in the real estate investment trust markets. To the best of the author's knowledge, most of these studies pay attention to the developed countries such as the US and OECD REIT markets, a few of them to the emerging markets, for example, India (Vishwakarma, 2012, 2013) and China (Hui and Yue, 2006; Paskelian and Vishwakarma, 2011). No one has ever tested for the periodically collapsing bubble for the Japanese REIT markets.[2] This paper fills the gap. The aim of this study is to investigate the issue regarding the periodically collapsing bubbles in Japanese real estate investment trust (REIT) markets.

The previous studies have so far typically provided an inconclusive answer to the periodically collapsing bubbles. However, a REIT price may face a bubble problem for a number of periods, but in the long run the REIT price is determined by its market fundamentals, i.e., the present discounted value of the future expected dividend. This paper takes this 'possibility' into account and examines whether there is a periodically collapsing bubble in Japanese real estate investment trust markets. In doing so, we adopt the momentum threshold (hereafter MTAR) unit root test, proposed by Enders and Granger (1998) and Enders and Siklos (2001), in this study. The MTAR model allows for the possibility of a regime shift between two different trend paths over time.

The MTAR model is attractive because it is powerful in testing for periodically collapsing bubbles. As explained in Bohl (2003), the MTAR model can be used to analyze bubble driven run-ups in stock prices followed by a crash in a cointegration framework by asymmetric adjustment. This technique offers a more potent insight in the stock price behavior than can possibly be obtained using conventional linear cointegration tests. The Monte Carlo simulation findings of Bohl (2003) show that the MTAR approach provides a sufficiently powerful test to detect periodically collapsing bubble behavior when the actual data generating process is given by the bubble model put forward by Evans.

Recently, Phillips, Wu, and Yu (2011, hereafter PWY) and Phillips, Shi, and Yu (2015, hereafter PSY) have proposed new bubble detection strategies based on recursive and rolling ADF unit root tests (sup-ADF) that enable us to detect bubbles in the data and to date-stamp their occurrence. These types of tests use a right tail variation of the Augmented Dickey-Fuller unit root test wherein the null hypothesis is of a unit root and the alternative is of a mildly explosive process.[3] However, as noted by Adammer and Bohl (2015, p. 69), this approach cannot answer the question of dependencies between different prices and fundamentals since the sup-ADF test investigates whether prices are temporarily mildly explosive. On the contrary, the MTAR approach avoids this deficiency by estimating a small number of parameters in the regression.

In addition, in order to take the possibility of non-linear trends into consideration, we also use the logistic smooth transition momentum threshold (hereafter LNV-MTAR) unit root test,

[2] A related study on the issue of bubble detection for the Japanese real estate investment market is Pierce (2006). However, he did not test for the periodically collapsing bubbles for the Japanese REIT markets.

[3] For applications of the PWY and PSY to detect bubbles in the different asset markets, readers are referred to, for example, Gutierrez (2011), Yiu, Yu and Jin (2013), Engsted, Hviid and Pedersen (2016), Fantazzini (2016) and Escobari and Jafarinejad (2016).

championed by Cook and Vougas (2009), in this paper.[4] This approach permits a structural break to occur gradually over time instead of instantaneously. In the context of economic time series, this has considerable intuitive appeal. Generally speaking, changes in economic aggregates are influenced by the changes in behavior of a very large number of agents. It is highly unlikely that all individual agents will react simultaneously to a given economic stimulus; while some may be able to (and want to) react instantaneously, others will be prone to different degrees of institutional inertia (dependent, for instance, on the efficiency of the markets in which they have to operate) and so will adjust to different time lags (Leybourne and Mizen, 1999, p 804).

The remainder of this paper is organized as follows. Section 2 reviews the theoretical foundation of the relation between a non-linear asset price and dividend. Section 3 outlines the statistical methods used for testing for nonlinearity and unit roots. Section 4 discusses the data used and the empirical results and compares our results with the extant literature. Finally, section 5 concludes.

2. Present Value Model and Periodically Collapsing Bubble

Define the net simple return on a stock as

$$R_{t+1} = \frac{P_{t+1} - P_t + D_{t+1}}{P_t} = \frac{P_{t+1} + D_{t+1}}{P_t} - 1 \tag{1}$$

where *Rt*+1 denotes the return on the stock held from time t to $t + 1$ and *Dt+1* is the dividend in period $t + 1$. The subscript $t + 1$ denotes the fact that the return only becomes known in period $t + 1$.

The presence of time-varying expected stock returns has led to a non-linear relation between prices and returns. Campbell and Shiller (1988) suggest a log-linear approximation of Eq. (1) and we obtain

$$p_t = \frac{\alpha}{1-\lambda} + \sum_{j=0}^{\infty} \lambda^j \left[(1-\lambda) d_{t+1+j} - r_{t+1+j}\right] \tag{2}$$

where the lower case letters *p*, *d* and *r* denote the logarithm of prices, dividends and the discount rate, respectively. The symbols l and a denote linearization parameters which are $\lambda = 1/[\exp(\overline{\mathrm{d} - \mathrm{p}})]$ and $\alpha = -\log(\lambda) - (1-\lambda)\log(1/\lambda - 1)$. Finally, taking the mathematical expectation of (10) based on information available at time *t, and* rearranging in terms of the log dividend-price ratio yields

$$d_t - p_t = -\frac{\alpha}{1-\lambda} + E_t\left[\sum_{j=0}^{\infty} \lambda^j \left[-\Delta d_{t+1+j} + r_{t+1+j}\right]\right] \tag{3}$$

According to (3), if asset prices (p_t) and real dividends (d_t) follow integrated processes of order one, and no bubbles are present, the log asset price and the log dividends are cointegrated with the cointegrating vector (1,–1) and the log dividend-price ratio ($d_t - p_t$) is a stationary process under no rational bubble restriction. On the contrary, the presence of a unit-root of the log dividend-price ratio is consistent with rational bubbles in asset markets.

Evans (1991) questions the approach undertaken by Diba and Grossman (1988) in that a class of bubbles, known as periodically collapsing bubbles, may very well exist that would not be detected by simple cointegration techniques. Recognizing the issue raised by Evans (1991),

[4] Xie and Chen (2015) adopt the same methodology to examine the periodically collapsing bubble for the US REIT markets.

Waters and Payne (2007) modify this model to allow for the possibility of positive and negative periodically collapsing bubbles as follows:

$$B_{t+1} = (1+r)B_t\nu_{t+1}, \text{if } |B_t| \le \alpha, \tag{4}$$

$$B_{t+1} = [\delta + (1+r)]\pi^{-1}\theta_{t+1}(B_t - (1+r)^{-1}\delta)\nu_{t+1}, \text{if } |B_t| > \alpha, \tag{5}$$

where the parameters in the equations (4) and (5) satisfy δ, $\alpha > 0$ and $0 < \delta < (1+r)\,\alpha$. The stochastic process ν_t is an i.i.d., has conditional expectation $E_t\nu_{t+1} = 1$ and is always positive, so that $\nu_t > 0$, which ensures that a bubble will not switch sign. The term θ_t represents a Bernoulli process that takes the value 1 with probability p and the value 0 with probability $1-\pi$. Equation (4) represents the phase when the bubble grows at a mean rate $(1+r)$ but Equation (5) shows that if the bubble exceeds the threshold a, it explodes at mean rate $(1+r)\pi^{-1}$. However, this phase does not last indefinitely as the bubble collapses with probability $1-\pi$ in each period. The nonlinearity of the process in Equation (4) and (5) creates difficulties in detecting such bubbles via standard cointegration tests between prices and dividends.

3. Methodology

3.1 MTAR Unit Root Test

The well-known Dickey-Fuller test and its extensions assume a unit root as the null hypothesis and a symmetric adjustment process under the alternative. These tests are mis-specified if the adjustment dynamics are non-linear or asymmetric. A formal way to quantify an asymmetric adjustment process as a generalization of the Dickey-Fuller test is given by the MTAR model proposed by Enders and Granger (1998) and Enders and Siklos (2001). Consider the following regression:

$$\Delta(d - p)_t = \mathrm{I}_t\rho_1(d - p)_{t-1} + (1 - I_t)\rho_2(d - p)_{t-1} + \varepsilon_t \tag{6}$$

where the indicator variable is defined as:

$$I_t = \begin{cases} 1, \text{if } \Delta(d-p)_{t-1} \ge \tau, \\ 0, \text{if } \Delta(d-p)_{t-1} < \tau, \end{cases} \tag{7}$$

and *t* denotes the value of the threshold and is derived by minimizing the residual sum of squares. The MTAR model allows the speed and direction of adjustment, represented by ρ_1 and ρ_2, depending on the previous period's change in$(d-p)_{t-1}$. This model is especially valuable when the adjustment is believed to exhibit more momentum in one direction than the other, as in the case of collapsing bubbles (Evans, 1991).

If the system is convergent, $\Delta(d - p)_t = \tau$ is the long-run equilibrium value. In case $\Delta(d - p)_t$ is above its long-run equilibrium value, the adjustment is$\rho_1(d - p)_{t-1}$, and if it is below its equilibrium value, the adjustment is $\rho_2(d - p)_{t-1}$. The Dickey-Fuller test is a special case of the MTAR model (6) and (7) in case of a symmetry in the error correction process $\rho_1 = \rho_2$.

The MTAR model sets up the null hypothesis of a unit root in the log dividend-price ratio, that is, H_0: $\rho_1 = \rho_2 = 0$. The distribution of this statistic is non-standard and, therefore, the critical values provided in Enders and Granger (1998), and Enders and Siklos (2001), are used. We denote the statistic testing the null hypothesis of a unit root (or no cointegration) as F_C. If this null hypothesis is rejected, then the null hypothesis of symmetric adjustment, H_0: $\rho_1 = \rho_2$, can be tested using the usual F-statistics denoted as F_A. In case the null hypothesis H_0: $\rho_1 = \rho_2$ is not rejected, a linear and symmetric adjustment in the log dividend-price ratio is favored.

3.2 LNV-MTAR Unit Root Test

Cook and Vougas (2009) combine the ideas of Enders and Granger (1998) and Leybourne et al. (1998) and develop a test for the null hypothesis of a unit root, under the alternative hypothesis allows for a stationary asymmetric adjustment around a smooth transition between deterministic linear trends.

Leybourne et al. (1998) consider three models:

Model A $y_t = \alpha_1 + \alpha_2 S_t(\gamma, c) + \nu_t,$ (8)

Model B $y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, c) + \nu_t,$ (9)

Model C $y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, c) + \beta_2 t S_t(\gamma, c) + \nu_t,$ (10)

where ν_t is a zero mean I(0) process, and $S_t(\gamma, c)$ is the logistic smooth transition function:

$$S_t(\gamma, c) = [1 + exp\{-\gamma(t - cT)\}]^{-1} \quad (11)$$

and the parameter c determines the timing of the transition midpoint. Since $\gamma > 0$, we have $S_{-\infty}(\gamma, c) = 0$, $S_{+\infty}(\gamma, c) = 1$, and $S_{cT}(\gamma, c) = 0.5$. The speed of transition is determined by the parameter γ. If ν_t is a zero-mean I(0) process, then in Model A y_t is stationary around a mean which changes from an initial value α_1 to a final value $\alpha_1 + \alpha_2$. Model B is similar, with the intercept changing from α_1 to $\alpha_1 + \alpha_2$, but it allows for a fixed slope term. In Model C, in addition to the change in the intercept from α_1 to $\alpha_1 + \alpha_2$, the slope also changes simultaneously, and with the same speed of transition, from $\beta_1\ to\ \beta_1 + \beta_2$.

Cook and Vougas (2009) combine Eqs (8)–(10), (12) and (13) and propose a logistic smooth transition-momentum TAR (LNV-MTAR) model as follows:

$$\Delta\hat{\nu}_t = M_t\hat{\rho}_1\hat{\nu}_{t-1} + (1 - M_t)\hat{\rho}_2\hat{\nu}_{t-1} + \sum_{i=1}^{k}\hat{\delta}_i\,\Delta\hat{\nu}_{t-i} + \hat{\eta}_t, \quad (12)$$

where M_t is the Heaviside indicator function,

$$M_t = \begin{cases} 1, if\ \ \Delta\,\hat{\nu}_{t-1} \geq\ 0, \\ 0, if\ \ \Delta\,\hat{\nu}_{t-1} <\ 0, \end{cases} \quad (13)$$

$\hat{\nu}_t$ is the residual from the first step by using the non-linear least squares for equation (12). If the null hypothesis of H0: $\rho_1 = \rho_2 = 0$ cannot be rejected in Eq. (13), then $\hat{\nu}_t$ and therefore y_t contains a unit root. The statistics are referred to as F_α^* , $F_{\alpha(\beta)}^*$ and $F_{\alpha\beta}^*$ corresponding to Models A to C, respectively. If the null hypothesis of H0: $\rho_1 = \rho_2 = 0$ is rejected and $\rho_1 = \rho_2 < 0$ holds, then $\hat{\nu}_t$ (y_t) is a stationary LNV-MTAR process with symmetry adjustment. If H0: $\rho_1 = \rho_2 = 0$ is rejected and $\rho_1 < 0$, $\rho_2 < 0$, $\rho_1 \neq \rho_2$ hold, then $\hat{\nu}_t$ (y_t) is a stationary LNV-MTAR process displaying asymmetric adjustment. Critical values must be tabulated via Monte Carlo simulations.

4. Data and Results

4.1 Data Description and Basic Statistics

The sample period was determined primarily based on the availability of the data. Monthly data on the price indices and dividend yields for the three broad classifications of REITs in Japan, the Composite, Office and Residential REITs covering the period from 2001:09 to 2012:07 are downloaded from the Sumitomo Mitsui Trust Research Institute at the following website: http://www.smtri.jp/en/JREIT_Index/index.html.

Some descriptive statistics of changes in the dividend-price series are outlined in Table 1. First, the coefficients of skewness of all of the series are positive, implying that returns are flatter to the right compared to the normal distribution. The coefficients of excess kurtosis for the raw returns are much higher than 0, indicating that the empirical distributions of these samples have fat tails. The coefficients of skewness and excess kurtosis reveal non-normality in the data. This is confirmed by the Jarque-Bera normality test as shown in Table 1. Second, the Ljung-Box Q-statistics, LB(24), indicate significant autocorrelations for all of the series. We also report a standard ARCH test for the raw returns. The test results indicate that an insignificant ARCH effect exists for the REIT of Japan.

Table 1: Summary Statistics

	Japan REITs		
	Composite	Office	Residential
Mean	0.001	−0.001	0.013
S.D.	0.11	0.113	0.168
SK	0.938	0.663	2.31
EK	3.412	1.627	13.619
JB	82.142**	23.887**	775.688**
LB(24)	38.294**	37.698**	23.583**
ARCH(4)	1.406	0.623	0.173

(1) ** denotes significance at the 5% level. (2) Mean and S.D. refer to the mean and standard deviation, respectively. (3) SK is the skewness coefficient. (4) EK is the excess kurtosis coefficient. (5) JB is the Jarque-Bera statistic. (6) LB(24) is the Ljung-Box Q statistic calculated with twenty-four lags. (7) ARCH(4) is the ARCH test calculated with four lags on raw returns.

As a preliminary analysis, we apply a battery of linear unit root tests to determine the order of integration of the dividend-price ratio. We consider the Augmented Dickey-Fuller (ADF) test, as well as the ADF-GLS test of Elliott et al. (1996) in this study. Vougas (2007) highlights the usefulness of the Schmidt and Phillips (1992) (SP hereafter) unit root test in practice. Therefore, we also employ it in this study. These authors propose some modifications of existing linear unit root tests in order to improve their power and size. For the ADF and ADF-GLS tests, an auxiliary regression is run with an intercept and a time trend. To select the lag length (k) we use the 't-sig' approach proposed by Hall (1994). That is, the number of lags is chosen for which the last included lag has a marginal significance level less than the 10% level.

Table 2 reports a battery of unit root tests for three broad classifications of REITs in Japan, namely, the Composite, Office and Residential. Based on the results from Table 2, it is shown that the null hypothesis of a unit root cannot be rejected at the 5% significance level for the ADF, SP(1), SP(2) and ADF-GLS statistics and therefore favors the rational bubble hypothesis.[5]

As Perron (1989) pointed out, in the presence of a structural break, the power to reject a unit root decreases if the stationary alternative is true and the structural break is ignored. To address this, we use Zivot and Andrews' (1992) sequential one trend break model and Lumsdaine and Papell's (1997) two trend breaks model to investigate the order of the empirical variables. We use the 't-sig' approach proposed by Hall (1994) to select the lag length (k). We

[5] The terms SP(1) and SP(2) denote the Schmidt-Phillips τ tests with a linear and quadratic trend, respectively.

set k max = 12 and use the approximate 10% asymptotic critical value of 1.60 to determine the significance of the t-statistic on the last lag. We use the 'trimming region' [0.10T, 0.90T] and select the break point endogenously by choosing the value of the break that maximizes the ADF t-statistic. We report the results in the bottom panel of Table 2. The results suggest that, for the Japanese REIT markets, the null hypothesis of a unit root cannot be rejected at the 5% significance level, indicating that the dividend yields are non-stationary in their respective levels. These findings fully echo those obtained from the linear unit roots.

Table 2: Results of the linear unit root tests—Japan

	Linear trend		
	ADF	SP(1)	DF-GLS
Composite	−2.314	-1.205	−1.113
Office	−1.609	-1.233	−1.070
Residential	−2.352	−1.355	−1.263
	Quadratic trend and breaks test		
	SP(2)	ZA, Model C	LP, Model C
Composite	−1.430	−4.203	−6.099
Office	−1.581	−4.465	−6.603
Residential	−1.550	−3.252	−6.257

(1) *, **, *** denote significance at the 10%, 5% and 1%, respectively. (2) ADF, SP(1) and DF-GLS denote the augmented Dickey-Fuller test, Schmidt-Phillips t test with linear trend and Elliott et al. (1996) DF-GLS test, respectively. (3) SP(2), ZA and LP denote the Schmidt-Phillips t test with quadratic trend, Zivot and Andrews (1992) and Lumsdaine and Papell (1997) tests, respectively. (4) The 5% critical values for the ADF, SP(1) and DF-GLS tests are −3.43, −3.04 and −2.89, respectively. (5) The 5% critical values for the SP(2), ZA and LP tests are −3.55, −5.08 and −6.75, respectively.

4.2 Results from the MTAR and LNV-MTAR approaches

Following Bohl and Siklos (2004), we report the results for the demeaned, as well as demeaned and detrended data on $\Delta(d-p)_t$ based on the following reason: if there is a time trend in the data and the regression equation does not contain a trend term, then the test has low power. On the other hand, if the regression equation contains a trend term, but a trend does not exist in the data, then the null hypothesis is too often rejected. The first difference of the log dividend-price ratio is demeaned by regressing $\Delta(d-p)_t$ on a constant, C, and, alternatively, demeaned and detrended, C, T, by regressing $\Delta(d-p)_t$ on a constant, as well as a linear trend prior to estimation in the MTAR regression equation. Hence, we allow for a constant term and a linear trend as attractors. We perform the tests with a linear time trend included due to its possible impact on the properties of the tests.

The threshold, τ, is consistently estimated via Chan's (1993) method. This involves sorting the estimated residuals in ascending order, excluding 15% of the largest and smallest values, and selecting from the remaining 70% the threshold parameter which yields the lowest residual sum of squares (Enders and Siklos, 2001). We employ the 't-sig' approach proposed by Hall (1994) to select the lag length (k). We set k_{max} = 12 and use the approximate 10% asymptotic critical value of 1.60 to determine the significance of the t-statistic on the last lag.

The results of the MTAR test for Japan's REIT dividend-price ratios are reported in Table 3. It is shown that the null hypothesis of a unit root can neither be rejected at the 5 percent significance level for the demeaned data nor for the demeaned and detrended data. The empirical evidence of the MTAR statistics favors the rational bubble in Japan's REIT markets. Since the F_C statistics are not rejected for Japan's REIT dividend-price ratios, we skip the discussion of tests for symmetric adjustment.

Table 3: Results of the MTAR unit root test— Japan

	demeaned data		
	Composite	Office	Residential
Attractor	-8.677	−8.406	−6.874
F_C	1.899	2.415	2.676
F_A	2.179	1.363	0.711
	[0.142]	[0.245]	[0.401]
ρ_1	0.008	−0.062**	-0.067**
	(0.020)	(0.029)	(0.031)
ρ_2	−0.038	−0.017	-0.027
	(0.019)	(0.025)	(0.034)
	demeaned and detrended data		
	Composite	Office	Residential
Attractor	$-8.555+0.009t$	$-8.734 + 0.005t$	$-7.636 +0.013t$
F_C	3.335	3.268	3.12
F_A	1.473	1.828	0.123
	[0.227]	[0.178]	[0.726]
ρ_1	-0.064	-0.067	-0.086
	(0.027)	(0.027)	(0.042)
ρ_2	−0.021	-0.018	−0.063
	(0.023)	(0.023)	(0.044)

(1) *, **, *** denote significance at the 10%, 5% and 1%, respectively. (2) F_C and F_A denote the F-statistics for the null hypothesis of a unit root $H_0: \rho_1 = \rho_2 = 0$ and the null hypothesis of symmetry $H_0: \rho_1 = \rho_2$, respectively. (3) The 10%, 5% and 1% critical values for the F_C statistic of demeaned data are 4.05, 4.95 and 6.91, respectively. (4) The 10%, 5% and 1% critical values for the F_C statistic of demeaned and detrended data are 5.60, 6.57 and 8.74, respectively. (5) The numbers in parenthesis are standard errors. (6) The numbers in square parenthesis are p-values.

A prime culprit is the ignorance of the structural break as noted by Jirasakuldech et al.

(2006) and Payne and Waters (2007). They stress the importance of recognizing the possibility of a structural shift in the REIT prices and dividends in testing for the null hypothesis of a unit root. We take this possibility into consideration by employing Cook and Vougas's (2009) LNV-MTAR approach. The results of applying the LNV-MTAR test of Model B as well as Model C of Japan's REIT dividend-price ratios are reported in the top and bottom panels in Table 4, respectively. The F_C ($F^*_{\alpha(\beta)}$ and $F^*_{\alpha\beta}$) statistics are significant at he 5% level, and reject the null hypothesis of a unit root in the log dividend-price ratio, irrespective of the chosen model. This finding, again, can be interpreted as evidence in favor of a cointegrating relationship between p_t and d_t with a [1,−1] cointegrating vector. That is, the log dividend-price ratio is a non-linear stationary process. Hence, our empirical evidence generally supports the long-run validity of the present value model with time-varying expected returns for the US REIT markets. However, the F_A statistics are insignificant at the 5 percent significance level, indicating that the log dividend-price ratio does not exhibit differing speeds of adjustment toward the long-run equilibrium and, therefore, does not favor periodically collapsing bubbles.

Figures 1 to 3 present the time series plots of the dividend-price ratios (black line) and the estimated logistic smooth transition functions (blue line) for Japan, respectively. Intuitively[1], if the true data generating process follows the logistic smooth transition nonlinear process, then the estimated logistic smooth transition trend is close to the raw data. As such, it is highly possible to reject the null hypothesis of non-stationarity. Taking Japan's Composite REIT as an example (Figure 1), the estimated logistic smooth transition trend of Model C is quite close to those of the raw data. These plots echo the rejections of the null hypothesis of a unit root by the $F^*_{\alpha(\beta)}$ and $F^*_{\alpha\beta}$ statistics as shown in Table 4.

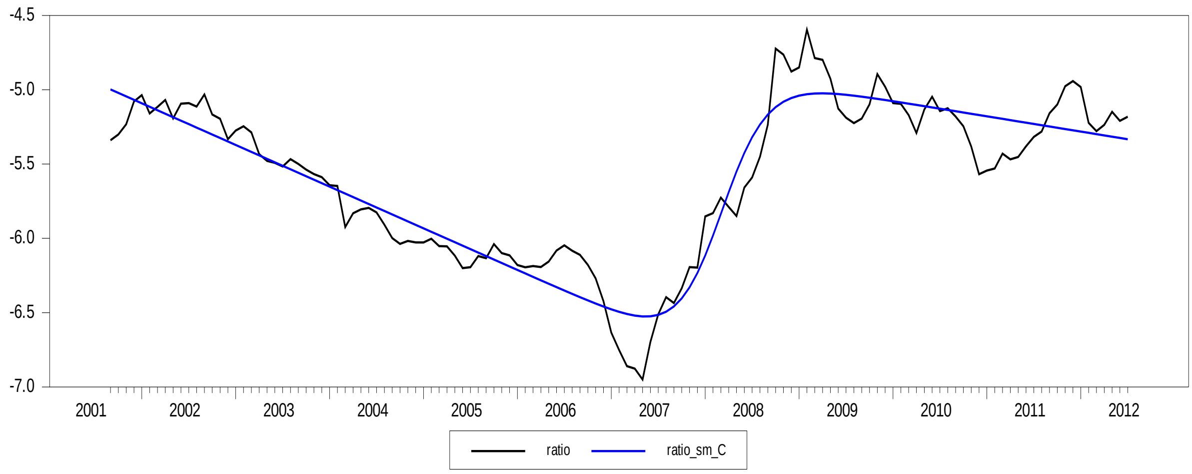

Figure 1: Scatter plot of the logarithm dividend-price ratio (black line) and fitted smooth transition function (blue line) of Model C for Japan Composite REIT.

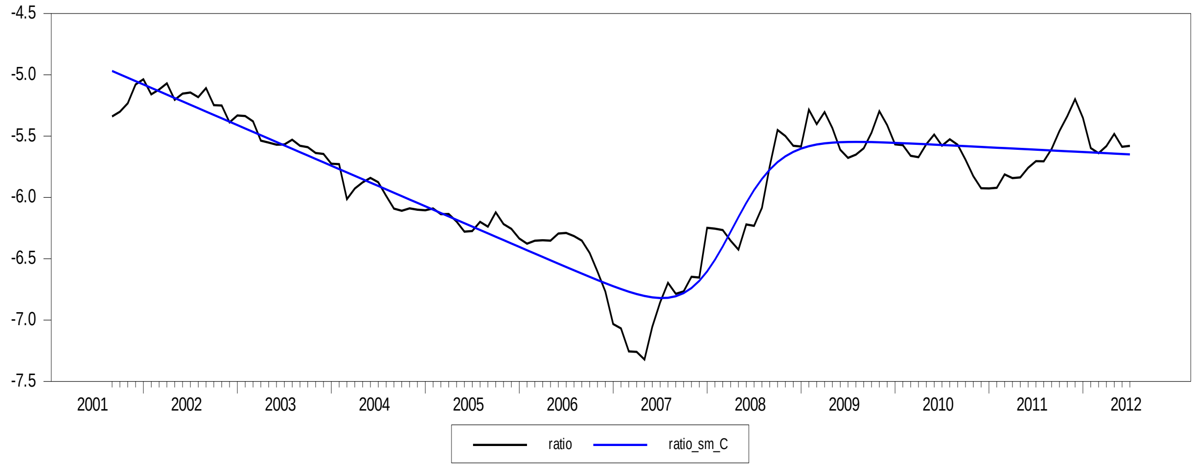

Figure 2: Scatter plot of the logarithm dividend-price ratio (black line) and fitted smooth transition function (blue line) of Model C for Japan Office REIT.

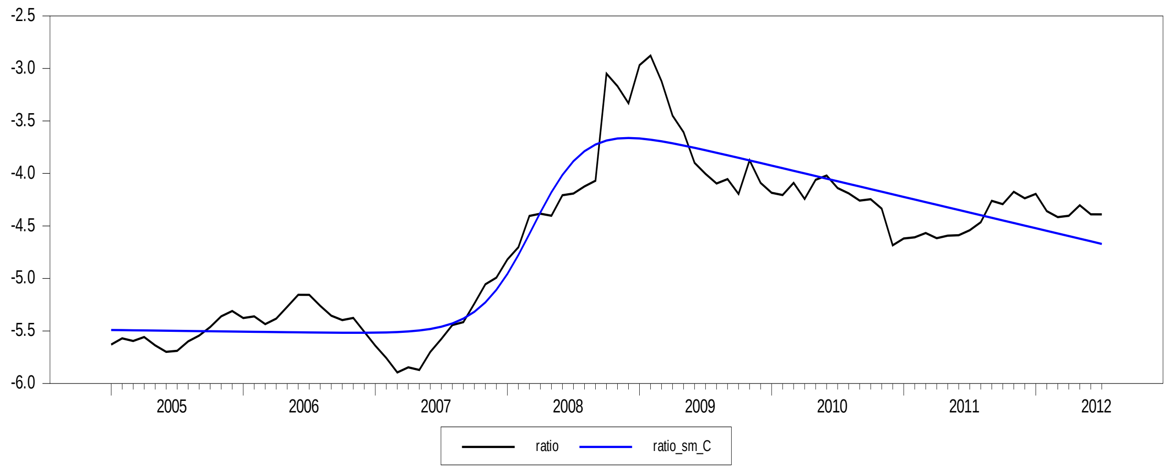

Figure 3: Scatter plot of the logarithm dividend-price ratio (black line) and fitted smooth transition function (blue line) of Model C for Japan Residential REIT.

5. Concluding Remarks

This paper examines the mean-reversion patterns of the dividend-price ratios of Japan REIT markets in order to test for the bubble behavior. A variety of unit root tests ranging from univariate estimators to non-linear testing principles have been employed in an effort to obtain inferences that are robust to problems associated with non-stationarity. This study makes use of an idea from Bohl and Siklos (2004) and adopts the MTAR unit root test, which helps detect a non-linear dividend-price relationship without specifying a threshold in advance. We also employ Cook and Voguas' (2009) approach, i.e., the LNV-MTAR root, which under the alternative hypothesis allows for a stationary asymmetric adjustment around a smooth transition between deterministic linear trends.

This study reaches the following key conclusions. First, by using a battery of univariate unit root tests, we have obtained evidence in favor of non-stationary dividend-price ratio series in Japanese REIT markets, which is consistent with the rational bubble hypothesis. Second, empirical evidence from the MTAR and LNV-MTAR test (i.e., the MTAR model with a smooth transition to characterize the structural break) shows that the periodically collapsing bubble does not hold in Japanese REIT markets. An important implication of this study is that if we neglect the nonlinear properties inherent in the data, then we are inclined to wrongly agree with the existence of speculative bubble based only on the conventional linear approaches.

Table 4: Results of the LNV-MTAR unit root test— Japan

	Model B		
	Composite	Office	Residential
$F^*_{\alpha(\beta)}$	7.311**	6.261**	9.585**
F_A	1.59	0.368	0.018
	[0.209]	[0.544]	[0.891]
ρ_1	−0.241**	-0.185**	−0.310**
	(0.070)	(0.061)	(0.086)
ρ_2	−0.127**	−0.134**	−0.293**
	(0.062)	(0.063)	(0.098)
	Model C		
	Composite	Office	Residential
$F^*_{\alpha\beta}$	11.424**	14.291**	11.579**
F_A	0.304	0.131	1.974
	[0.581]	[0.717]	[0.163]
ρ_1	−0.293**	−0.325**	−0.273**
	(0.076)	(0.074)	(0.093)
ρ_2	−0.237**	−0.288**	−0.449**
	(0.073)	(0.081)	(0.104)

(1) *, **, *** denote significance at the 10%, 5% and 1%, respectively. (2) $F^*_{\alpha(\beta)}$ and F_A denote the F-statistics for the null hypothesis of a unit root $H_0: \rho_1 = \rho_2 = 0$ and the null hypothesis of symmetry $H_0: \rho_1 = \rho_2$, respectively. (3) The critical values for the LNV-MTAR statistics are obtained from Cook and Vougas (2009). (4) The numbers in parenthesis are standard errors. (5) The numbers in square parenthesis are p-values.

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