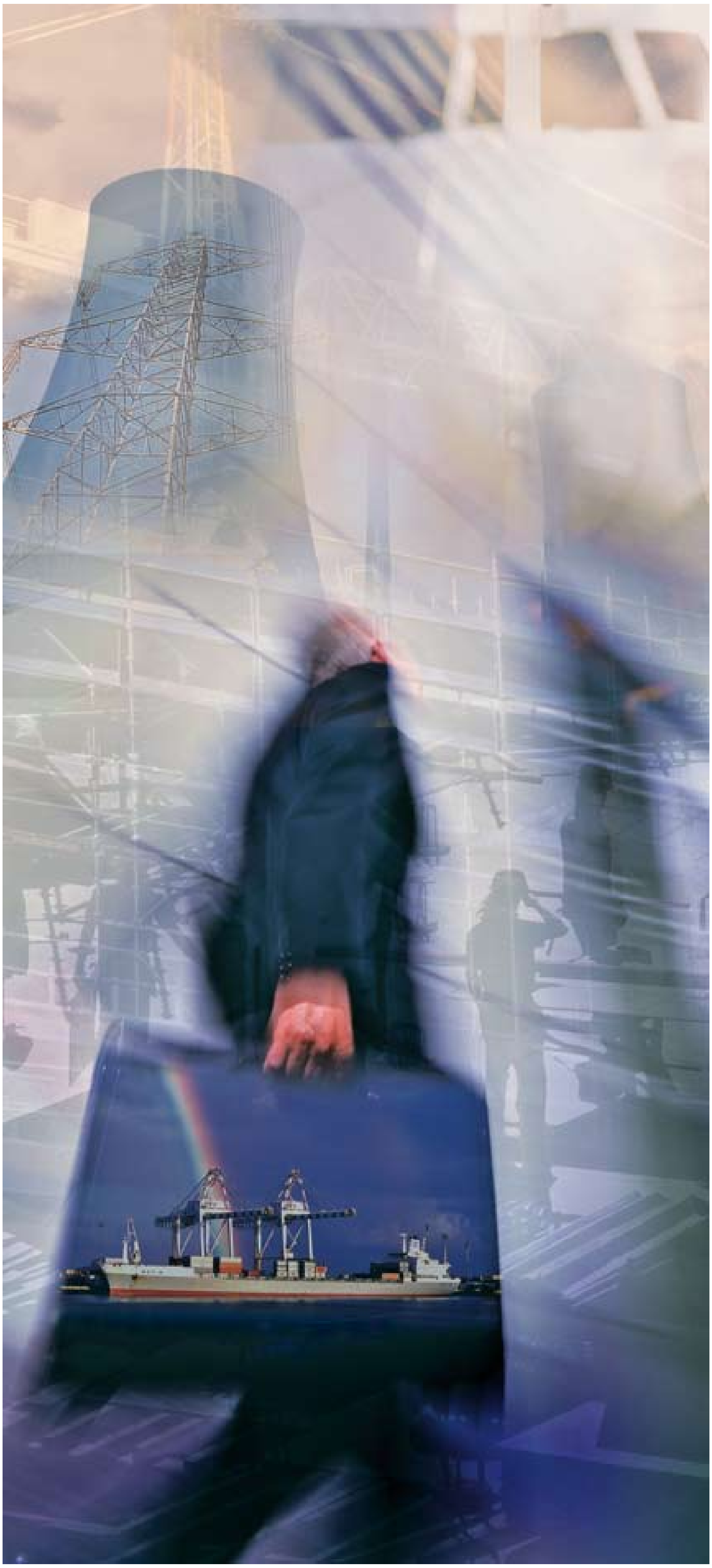


Queensland

The Economic Society of Australia Inc.

Proceedings of the 37th Australian Conference of Economists

Papers delivered at ACE 08

30th September to 4th October 2008
Gold Coast Queensland Australia

ISBN 978-0-9591806-4-0

Welcome

The Economic Society of Australia warmly welcomes you to the Gold Coast, Queensland, Australia for the 37th Australian Conference of Economists.

The Society was formed 83 years ago in 1925. At the time, the Society was opposed to declarations of policy and instead focused on open discussions and encouraging economic debate. Nothing has changed today, with the Society and the conference being at the forefront of encouraging debate.

This year we have a large number of papers dealing with Infrastructure, Central Banking and Trade.

Matters of the greatest global importance invariably boil down to be economic problems. Recent times have seen an explosion of infrastructure spending, after world-wide population growth has seen demand outpace aging supply. The world has become more globalised than at any time since World War I but the benefits of this (and the impact on our climate) has been questioned by some.

At the time of preparing for this conference we could not have known that it would have been held during the largest credit crisis since the Great Depression. The general public and politicians both look to central banks for the answers.

We are also very pleased to see a wide selection of papers ranging from applied economics to welfare economics. An A – Z of economics (well, almost).

Another feature of this conference is that we have gone out of our way to bring together economists from all walks of life, in particular from academia, government and the private sector. We are grateful to all of our sponsors, who are as diverse as the speakers.

The Organising Committee

James Dick
Khorshed Alam (Programme Chair)
Michael Knox
Greg Hall
Allan Layton
Rimu Nelson
Gudrun Meyer-Boehm
Jay Bandaralage
Paula Knight

Published November 2008

GPO Box 1170
Brisbane Queensland Australia
ecosocqld@optushome.com.au

Our Gold Sponsors

Keynote Sponsors

Special Session Sponsors

Unless we have specifically been requested to do otherwise, all the papers presented at the conference are published in the proceedings in full. A small number of papers will have versions that have also been made available for special editions of Journals, Economic Analysis and Policy, and the Economic Record. Authors will retain the right to seek additional publication for papers presented at the conference so long as it differs in some meaningful way from those published here.

The Paper following forms part of - *Proceedings of the 37th Australian Conference of Economists*
ISBN 978-0-9591806-4-0

An Empirical Analysis of the Mean-Reversion Property of Real Interest Rates

Jae H. Kim*

Department of Econometrics and Business Statistics
Monash University, Caulfield East
VIC 3145, Australia

Philip Inyeob Ji

Department of Accounting and Finance
Monash University, Clayton, VIC 3800
Australia

Abstract

This paper examines the mean-reversion property of real interest rates of a number of industrialized and East Asian countries. Many past studies have found the evidence of mean-aversion, which has been a puzzling outcome. We have employed panel unit root tests and half-life estimation based on the bias-corrected bootstrap, which allow for the presence of cross-sectional dependence. We have found that the real interest rates are overall mean-reverting. It is also found that the real interest rates of industrialized countries show a lower rate of mean-reversion than those of East Asian markets. For East Asian emerging markets, we have found evidence that the speed of mean-reversion has become slower after the Asian crisis. Our results strongly suggest that the speed of mean-reversion is closely related with the degree of market friction caused by government interventions. There is a strong tendency that the rates form more advanced and liberalized markets show slower speed of mean-reversion.

Keywords: Bootstrap, Cross-sectional dependence, Half-life, Highest Density Region, Panel unit root tests, Stationarity.

JEL Classifications: F21, G15, C63

---

* Corresponding author. Tel: (61 3) 99031596 Fax: (61 3) 99032007
Email: Jae.Kim@BusEco.monash.edu.au

An earlier version of this paper was presented at the Annual Q-Group/Monash Melbourne Colloquium on Empirical and Quantitative Finance, 2006. We would like to thank the participants for helpful comments. We would like to thank David Rapach for making his data set available.

1. Introduction

The real interest rate, the nominal interest rate adjusted with inflation, represents the growth rate of purchasing power derived from an investment. It plays a crucial role in a wide range of prominent economics and finance models, including the neoclassical growth models (Solow, 1956), Black-Scholes option pricing model (Black and Scholes, 1973), consumption-based capital asset pricing models (Lucas, 1978; and Hansen and Singleton, 1982), and Fisher equation (Fisher, 1930). These models assume that, *ceteris paribus*, the real interest rate is constant over time. For example, the Fisher equation postulates that a rise in (expected) inflation rate will eventually cause an equal rise in the nominal interest rate. This indicates that, in practice, the real interest rate should be mean-reverting or stationary, where a deviation from its long run level has only a transitory effect.

The accumulated empirical evidence on the time series properties of real interest rate, however, has provided rather mixed and puzzling outcomes. Most past studies tested stationarity or otherwise of real interest rate using (univariate) unit root tests. Using annual and monthly US data, Rose (1988) found that the nominal interest rate is I(1) while inflation rate is I(0), and that real interest rate is I(1).[1] Rapach and Weber (2004), using quarterly data for 16 industrialized countries, found that both nominal interest and inflation rates are I(1). This result implies

[1] A time series with a unit root is non-stationary and exhibit a mean-averting behavior over time with a shock to the time series having a permanent effect. Such a time series is said to be integrated of order 1, denoted as I(1). A stationary I(0) time series with no unit roots shows mean-reverting behavior, where the time series shows a strong tendency to revert back to its long- run mean after a shock is given.

that the Fisher equation represents a spurious relationship between the nominal interest rate and inflation. On the other hand, Lai (2004, 2008) provided evidence that real interest rates are mean-reverting for a number of industrialized and developing countries, based on the univariate unit root tests with mean-shift alternatives. In contrast, Phillips (2005) examined the long memory property of real interest rate using US historical data, and concluded that it is fractionally non-stationary, i.e., it is of I(d) with $0.5 \leq d < 1$.

Rapach and Wohar (2004) took a different approach by calculating the half-life of real interest rates. The half-life is a popular measure of mean-reversion, defined as the number of periods (or years) required for the response of a time series to a unit shock to dissipate by one half. It has been used widely in international economics to test for the empirical validity of parity conditions such as the purchasing power parity (see, for example, Rogoff; 1996). The half-life of an I(0) time series is positive and finite, while that of an I(1) time series is infinite. Rapach and Wohar (2004) constructed confidence intervals for the half-life of real interest rates based on the grid bootstrap of Hansen (1999), which is a method applicable to a univariate time series. These authors found that the intervals for all 13 countries have infinite upper bounds. This indicates that the real interest rates are not statistically distinguishable from an I(1) time series, which is evidence of mean-aversion.

When this discussion reaches the equity premium puzzle and the risk free rate puzzle, spurred by Mehra and Prescott (1985) and Weil (1989), stationarity of real interest rate faces an even bigger challenge. Explanation of these puzzles has become one of the central issues in macroeconomics and finance, and continuing research efforts are being devoted to it up to date. This stream of literature found equity returns to be mean-reverting, while real interest rate mean-averting. For instance, Fama and French (1988) and Poterba and Summers (1988) found negative auto-correlation in equity returns over long horizons. Siegel and Thaler (1997) quoted these findings to suggest that the equity premium puzzle is even more challenging than what had been previously thought because not only the size of the equity premium but also now its sign is a mystery: "It is not that the risk of equities is not great enough to explain their high rate of return; rather, for long-term investors, fixed income securities have been riskier in real terms. By this reasoning, the equity premium should be negative!"

In this paper, we re-examine the mean-reversion property of real interest rates for a number of industrialized and East Asian countries. The former represent highly liberalized capital markets; while the latter a mixture of liberalized markets and emerging markets. For the former, we examine tax-adjusted quarterly real bond yield from thirteen countries, previously investigated by Rapach and Wohar (2004). For the latter, we use short-term monthly real interest rates of seven East Asian capital markets, paying attention to the effect of the Asian currency crisis in 1997, which exerted considerable impact on most of capital markets in this region. In the methodological content, this paper is distinct from the past studies on two points. First, we combine the results from panel unit root testing and half-life estimation in this paper. Second, our testing and estimation methods take account of cross-sectional dependence among the real interest rates, which is a point totally overlooked by the above-mentioned past studies.[2]

We use the panel unit root (IPS) test developed by Im et al. (2003) and the inverse normal test of Choi (2001) to test whether a set of real interest rates are stationary. We consider the subsampling versions of these tests proposed by Choi and Chue (2007), which allow for cross-sectional dependence of the time series in the panel. We conduct half-life estimation based on the bias-corrected bootstrap, recently proposed by Kim et al. (2007), where the highest density region (HDR) method of Hyndman (1996) is used to construct confidence intervals. This HDR method provides a more sensible way of constructing confidence intervals for half-life than the conventional methods, given the atypical distributional properties of the half-life estimator. Kim et al. (2007) provided Monte Carlo evidence that their bias-corrected bootstrap HDR method performs much better than the conventional methods in small samples. In this paper, we extend the methods of Kim et al. (2007), so that the half-life estimates are calculated using a vector autoregression, in order to take explicit account of cross-sectional dependence of the time series.

[2] There has been voluminous literature on real interest linkages across international capital markets (see, for example, Chin and Frankel, 1995, Glick and Hutchison, 1990; Goldberg et al, 2003; Goodwin and Grennes, 1994; Phylaktis, 1999). Current standing on the degree of real interest linkages in the literature is that high income countries such as European markets and the US appear to be integrated more than Asian markets in the context of cointegration, although there is little empirical support for real rate equalisation. For East Asian countries, Phylaktis (1999) found evidence to suggest that financial liberalisation increased integration.

The main finding of the paper is that international real interest rates are mean-reverting for both industrialized and East Asian countries. We have found that the East Asian real interest rates overall show a faster convergence to long-run mean than those of industrialized countries. For the East Asian emerging markets in particular (excluding Japan and Singapore), we have found evidence that the speed of mean-reversion has become substantially lower after the crisis. These results suggest that the markets with a higher degree of market friction caused by government interventions show faster mean-reversion of the real interest rates. In the next section, we present the details of the methodologies used in this study. Section 3 provides the data details, and Section 4 provides the empirical results. Section 5 concludes the paper.

2. Methodologies

An attraction of panel unit root testing is that, by pooling the observations from different cross-sectional units, the test can enjoy a larger sample size, which can give rise to a higher power. However, as Maddala and Kim (1998; p.138) noted, it may be of limited value in practice because it does not reveal the speed of convergence or mean-reversion associated with individual time series. To measure the speed of mean-reversion, the half-life is also estimated in this paper.

2.1. Panel Unit Root tests

Im-Pesaran-Shin (IPS) test

Im et al. (2003) considered a model of the form

$$\Delta Y_{it} = \mu_i + \phi_i Y_{it-1} + \sum_{j=1}^{p_i} \varphi_{ij} \Delta Y_{it-j} + e_{it}, \qquad (1)$$

where i (= 1, …., N) indicates a cross-sectional unit, t (= 1, …, T) is a time index and e_{it} ~ IID(0,σ^2). They specified the null and alternative hypotheses of the form

$$\begin{aligned} &H_0 : \phi_1 = \phi_2 = ... = \phi_N = 0 \\ &H_1 : \phi_1 < 0, \phi_2 < 0, ..., \phi_{N_0} < 0 \qquad (N_0 \leq N) \end{aligned} . \qquad (2)$$

The null hypothesis indicates that all time series in each cross-sectional unit are non-stationary with a unit root. Under the alternative hypothesis, at least N_0 time series are stationary. The test statistic is constructed from the t-test statistics calculated from individual cross-sectional units. Let $\hat{\tau}_i$ denote the augmented Dicky-Fuller (ADF) t-statistic to test for $\phi_i = 0$. Im et al. (2003) have shown that

$$\frac{\sqrt{N}(\hat{\tau}_N - E(\hat{\tau}_i))}{\sqrt{Var(\hat{\tau}_i)}} \Rightarrow N(0,1),$$

where $\hat{\tau}_N = \frac{1}{N}\sum_{i=1}^{N} \hat{\tau}_i$. That is, the average of $\hat{\tau}_i$ statistics over all cross-sectional units converges to the standard normal distribution, when standardized. The values of $E(\hat{\tau}_i)$ and $Var(\hat{\tau}_i)$ are tabulated in Im et al. (2003).

Inverse normal test

Choi (2001) suggested a test for the null and alternative hypotheses given in (2) based on the p-values of individual statistics. Let π_i denote the p-value of the individual t-statistic $\hat{\tau}_i$. Choi (2001) proposed the inverse normal test

$$Z_{NT} = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Phi^{-1}(\pi_i),$$

where $\Phi(\cdot)$ is the standard normal cumulative distribution function. Under the null hypothesis given in (1), as T $\to\infty$, Z_{NT} converges to $\frac{1}{\sqrt{N}}\sum_{i=1}^{N} z_i$ where z_i follows the standard normal distribution.

Subsampling for cross-sectional dependence

The IPS and Fisher ADF tests above assume cross-sectional independence. This can be unrealistic in practice since it is highly likely that the time series in the panel are contemporaneously correlated. That is, the vector of error term (e_{1t}, …, e_{NT}) has the covariance matrix with non-zero off-diagonal elements. To allow for this dependence, Choi and Chue (2007) proposed the subsampling versions of the IPS and Fisher ADF tests. The subsampling method (Romano and Wolf, 2001) approximates the sampling distribution of a statistic, by calculating the test statistic of interest from overlapping subsamples of the observed data.

Let $Z_t = (Y_{1t}, \ldots, Y_{Nt})$ be the vector of panel time series at time t. Let the subsample of size be denoted as $W_s = (Z_s, Z_{s+1}, \ldots, Z_{s+b-1})$ ($1 \leq s \leq T\text{-}b+1$). Let T be the statistic of interest and T_s be the statistic calculated from the subsample W_s. Calculate the statistic for all possible subsamples give the collection of the statistics $\{T_s\}_{s=1}^{T-b+1}$, which can be used as an approximation to the sampling distribution of T. Since a subsample preserves cross-sectional dependence, $\{T_s\}_{s=1}^{T-b+1}$ provides a small sample distribution of T, allowing for cross-sectional dependency. Choi and Chue (2007) proposed the subsampling versions of IPS and

Fisher ADF tests. They showed that, under certain regularity conditions, these subsampling tests are consistent and possess valid asymptotic properties (Choi and Chue, 2007; Theorems 1 and 2). To implement the tests for the null hypothesis given in (2), we use the critical values obtained from $\{T_s\}_{s=1}^{T-b+1}$.

In implementing the subsampling tests, the choice of the subsample length b is important. Choi and Chue (2007) proposed two data-dependent methods: one based on simulation-based calibration rule and the other on minimum volatility rule. Both methods are used in this paper, although their details are not given here for simplicity.

2.2 Half-life estimation

The half-life is widely used as a measure of persistence or mean-reversion of economic time series, particularly in the context of testing for the validity of parity conditions in international economics. For example, mean-reversion of real exchange rates is a key condition for the empirical validity of purchasing power parity (Rogoff, 1996). In the simple case of the univariate AR(1) model, the half-life is calculated as log(0.5)/log(α), where α is the AR(1) slope coefficient. In the higher order or higher dimensional case, a closed form solution does not exists, but the half-life can be calculated from the impulse response function.

In this paper, we calculate point and interval estimates of half-life using the bias-corrected bootstrap proposed by Kim et al. (2007), who reported Monte Carlo evidence that their method performs sunstantially better than the existing methods of half-life estimation. We extend their bias-corrected bootstrap based on the univariate AR model to the vector autoregression (VAR), and estimate the half-life of a time series taking account of cross-sectional dependence.

VAR model and estimation

We consider the N-dimensional vector autoregressive (VAR) model of the form

$$Z_t = \nu + B_1 Z_{t-1} + \ldots + B_p Z_{t-p} + u_t, \tag{3}$$

where $Z_t = (Y_{1t}, \ldots, Y_{Nt})$ is the $N\times1$ vector of variables at time t, ν is the $N \times 1$ vector of intercepts, and B_i's are the $N\times N$ matrices of coefficients. Note that u_t is the $N\times1$ vector of innovations with $E(u_t) = 0$ and $E(u_t u_t') = \Sigma_u$ .The unknown VAR order p in (1) is estimated using the Bayesian information criterion (BIC).

The impulse responses are calculated from the coefficients of the MA(∞) representation of the VAR model. Given T realizations ($Z_1, \ldots, Z_T$) of (2), the unknown coefficients are estimated using the least-squares (LS) method. The LS estimators for $B = (\nu, B_1, \ldots, B_p)$ is denoted as $\hat{B} = (\hat{\nu}, \hat{B}_1, \ldots, \hat{B}_p)$, and the associated vector of residuals as $\{\hat{u}_t\}_{t=p+1}^{n}$. The impulse responses are defined as Φ_i where Φ_i's are the coefficients of the MA(∞) representation of (1). A typical element of Φ_i is denoted as $\theta_{kl,i}$, and it is interpreted as the response of the variable k to a one-time impulse in variable l, i periods ago. The plot of $\theta_{kl,i}$ against i is called the impulse response function of the variable k to a one-time impulse in variable l. Using $\hat{B}$, the estimator for impulse response $\hat{\theta}_{kl,i}$ for $\theta_{kl,i}$, can be calculated.

As mentioned earlier, the half-life of a time series is defined as the number of periods required for the response of a time series, to its own shock, to be halved. As such, it can readily be obtained from the impulse response function of a time series. In the VAR case, the half-life of the k^{th} time series in the system, denoted as h_k can be calculated from the impulse response function to its own shock, namely $\theta_{kk,i}$, where $k = 1, \ldots, K$. The half-life estimator for h_k, $\hat{h}_k$, can be obtained from $\hat{\theta}_{kk,i}$.

Bias-corrected bootstrap for point and interval estimation

The bootstrap is a computer-intensive method of approximating the sampling distribution of a statistic. It has been applied widely in econometrics and is often found to provide a superior alternative to the conventional methods in small samples (see, Li and Maddala, 1996; Berkowitz and Kilian, 2000; and MacKinnon, 2002). Note that impulse response estimates and half-life estimates are necessarily biased in small samples, due to small sample biases present in the VAR parameter estimators (see Tjostheim and Paulsen, 1983; Nicholls and Pope, 1988; and Pope, 1990). The biases are particularly severe when the VAR model has unit roots or near unit roots; when the VAR dimension N is larger; or when the sample size is smaller. It is highly likely that these biases adversely affect the small sample properties of the confidence intervals.

To obtain confidence intervals with improved small sample properties, Kilian (1998a, 1998b) proposed the use of the bias-corrected bootstrap (or bootstrap-after-bootstrap). It is a bootstrap method of constructing confidence intervals, in which the biases associated with parameter estimators are adjusted in two stages of the bootstrap. Although it was originally proposed for statistical inference of impulse response, the bias-corrected bootstrap can easily be adapted to half-life estimation. The bias-corrected bootstrap of Kilian (1998a, 1998b) involves two stages of bias-correction for VAR parameter estimates. Here we follow Kilian (1998b) in using Pope's (1990; p.253) asymptotic bias formula to obtain bias-corrected parameter estimators.

The bias-corrected confidence interval for $\theta_{kl,i}$ can be obtained as below:

In Stage 1, Pope's (1990) formula is applied to $\hat{B} = (\hat{v}, \hat{B}_1, ..., \hat{B}_p)$ to obtain the bias-corrected estimator $\hat{B}^c = (\hat{v}^c, \hat{B}_1^c, ..., \hat{B}_p^c)$ for B. It is possible that $\hat{B}$ satisfies the condition of stationarity, while $\hat{B}^c$ does not. In this case, Kilian (1998a, 1998b) suggested an adjustment to $\hat{B}^c$ so that it implies stationarity. This adjustment is called the stationarity correction, and its details can be found in Kilian (1998a, 1998b).

In Stage 2, generate a pseudo data set following the recursion

$$Y_t^* = \hat{v}^c + \hat{B}_1^c Y_{t-1}^* + ... + \hat{B}_p^c Y_{t-p}^* + u_t^*, \quad (3)$$

using the first p values of the original data as starting values and u_t^*'s are generated as random resampling of $\hat{u}_t$'s with replacement.

In Stage 3, using $\{Y_t^*\}_{t=1}^n$, the VAR coefficient matrices are re-estimated and denoted as $\hat{B}^* = (\hat{v}^*, \hat{B}_1^*, ..., \hat{B}_p^*)$. Pope's (1990) bias formula is again applied to $\hat{B}^*$ in order to obtain a bias-corrected version $\hat{B}^{*c} = (\hat{v}^{*c}, \hat{B}_1^{*c}, ..., \hat{B}_p^{*c})$ of $\hat{B}^*$. The stationarity correction is again applied to $\hat{B}^{*c}$ if necessary.

Repeat Stages 2 and 3 sufficiently many times, say m, to generate bootstrap replicates of $\left\{\hat{B}^{*c}(j)\right\}_{j=1}^m$, and the bootstrap replicates of half-life $\left\{\hat{h}_k^*(j)\right\}_{j=1}^m$ are obtained from $\left\{\hat{\theta}_{kk,i}^*(j)\right\}_{j=1}^m$ in Stage 3. The 100(1-2α)% bias-corrected bootstrap confidence intervals for h_k can be constructed as the interval $[\hat{h}_k^*(\alpha), \hat{h}_k^*(1\text{-}\alpha)]$, where $\hat{h}_k^*(q)$ is the qth percentile from the distribution of m bootstrap replicates $\left\{\hat{h}_k^*(j)\right\}_{j=1}^m$. To obtain a tight and informative confidence interval from $\left\{\hat{h}_k^*(j)\right\}_{j=1}^m$, we follow Kim et al. (2007) to use the HDR method of Hyndman (1996).[3]

HDR point and interval estimators for half-life

[3] Note that we have set m = 2000 throughout the paper.

Let $f(x)$ be the density function for a random variable X. The 100(1-θ)% HDR is defined (Hyndman, 1996) as the subset $R(f_\theta)$ of the sample space of X such that $R(f_\theta) = \{x: f(x) \geq f_\theta\}$, where f_θ is the largest constant such that $\Pr[X \in R(f_\theta)] \geq 1 - \theta$. $R(f_\theta)$ represents the smallest region with a given probability content. In short, the HDR method produces confidence intervals concentrated around the modes of the distribution. In the present context, X is the half-life estimator of a time series and its density can be estimated from the bootstrap replicates of the half-life $\left\{\hat{h}_k^*(j)\right\}_{j=1}^{m}$. We estimate the density $f(x)$ using a kernel estimator with the Gaussian kernel, with bandwidth selected using the Sheather-Jones (1991) rule. From the estimated density, the mode of the distribution is used as the bias-corrected point estimator for the half-life, along with the interval concentrated around the mode of the distribution with the probability content 100(1-θ)% . In the multi-modal case, the global mode and the associated interval are used as point and interval estimates of half-life.

Examples of the bias-corrected HDR point and interval estimates of the half-life are illustrated in Figure 1, where the density estimates of $\left\{\hat{h}_k^*(j)\right\}_{j=1}^{m}$ for the Thai real interest for Periods I and II are plotted. Note that X-axis is half-life replicates in years. Both graphs clearly indicate that the sampling distribution of half-life estimates heavily skewed, with extreme values on the right end of the distribution. These features are the reflection of the atypical distributional properties of half-life estimators mentioned earlier. The global mode is the bias-corrected HDR point estimate, while the horizontal lines corresponding to the f_θ values where $\theta =$ 0.05, and the associated intervals indicate 95% HDR confidence intervals. In Period I, the point estimate of half-life is 0.15 years, and 95% confidence interval is [0.08, 0.27] in years. In Period II, the distribution has much higher variability with higher proportion of extreme values. The point estimate 0.99 in Period II is a lot bigger than in Period I, and 95% confidence interval [0.25, 7.94] is also substantially wider. This indicates that the speed of convergence of Thai real interest rate has become slower after the Asian crisis. We provide more discussions about this point in Section 4.

3. Data

We re-evaluate the data set of Rapach and Wohar (2004), which are tax-adjusted real bond yields for 13 industrialized countries (Australia, Belgium, Canada, Denmark, France, Ireland, Italy, the Netherlands, New Zealand, Norway, Switzerland, the UK and the US), quarterly from 1960:4 to 1998:3.[4] For East Asian markets, we have selected short-term real interest rates for Japan, two NIE countries (Singapore and Korea) and four ASEAN members of developing countries (Indonesia, Malaysia, Philippines and Thailand).[5] These countries represent major emerging markets in East Asia, except for Japan which is included as a major player in this region. For East Asian rates, we have used monthly data from 1987:1 to 1996:12 (pre-crisis period; Period I), and from1999:1 to 2007:05 (post-crisis period; Period II). The starting date reflects the

[4] Rapach and Wohar (2004) provide more detailed description with the data source.
[5] The choice of country is largely determined by data availability.

timing of deregulation where most Asian countries started to liberate their financial markets. All observations from 1997 and 1998 were excluded to eliminate noisy and unstable observations. Most of East Asian countries hit by the crisis were already in financial distress in the fist half of 1997 and they only started to display a sign of recovery in September 1998.

For Japan and Indonesia, we use the call rate, while the interbank rate has been used for Singapore and Malaysia. Money market rate is used for Korea, Thailand and the Philippines. To generate real interest rate, seasonally adjusted ex-post inflation rate is subtracted from short term nominal interest rates. To calculate the rate of inflation, the consumer price index is seasonally adjusted with the X-12 method in Eviews (version 6). All nominal interest rates are then deflated by the ex-post inflation rate in order to produce the ex post real interest rate series. All data are obtained from *International Financial Statistic* Database.

4. Empirical Results and Discussion

Table 1 reports the panel unit root test results for Rapach-Wohar data. The orders p_i's in (1) are selected using BIC, following Choi and Chue (2007). The details of the selected orders are not reported for simplicity. Both IPS and the inverse normal tests rejects the null hypothesis of unit root for all capital markets, based on the 5% critical values obtained from the subsampling method. The critical values are not so sensitive to the different methods of subsample length selection. Table 2 reports the IPS and inverse normal tests for the East Asian capital markets.

For both Periods I and II, the two tests rejects the null hypothesis of unit root, based on the 5% critical values obtained from subsampling. Again, the critical values are not so sensitive to the different methods of subsample length selection.

The half-life estimates are obtained from VAR model given in (3). The VAR orders are selected using BIC. For the Rapach-Wohar data, 13-dimensional VAR(1); while, for the East Asian data, 7-dimensional VAR models are fitted for both Period I and II. Table 3 reports the case for the 13 developed capital markets. The grid bootstrap estimates of Rapach and Wohar (2004) are reproduced from their Table 1. The HDR point estimates are overall smaller than the grid bootstrap estimates: the former have the mean value of 0.97, while the latter 3.47. As for the interval estimates, the grid bootstrap intervals have infinite upper bounds for all markets, indicative of unit-root property of the real interest rates. In stark contrast, the bootstrap bias-corrected HDR interval estimates are tight with finite upper bounds. The mean values of the lower and upper bounds of the bootstrap bias-corrected HDR 95% confidence intervals are [0.44, 7.28], indicating that, on average, the half-life of real interest rates is between 0.44 and 7.28 years with 95% confidence. The main cause of this dramatic difference is our method of half-life estimation, which is different from that of Rapach and Wohar (2004) in the treatment of cross-sectional dependence and construction of bootstrap confidence intervals.

For East Asia, in Period I, the point estimates are in the range of 0.07 to 0.92 years with the mean value of 0.36 years. The 95% confidence intervals are tight, indicating that all rates are mean-reverting. The mean values of the lower and upper bounds of the 95% confidence intervals are 0.13 and 2.41. In Period II, similarly to Period I, all rates show small point estimates and tight confidence intervals. Overall, mean-reversion has become only slightly slower for most countries after the Asian crisis: the mean point estimate of 0.39 years in Period II is nearly same as that of Period I, but the 95% confidence intervals in Period II have median lower and upper bounds of 0.14 and 2.89, which are wider than those of Period I which are 0.14 and 0.89.

Speed of Convergence, Financial Frictions, and Risk Premia

A notable observation from the results presented above is the difference in half-life estimates among capital markets in differing stages of maturity. For the thirteen industrialized countries, the mean point estimate is 0.68 with the mean 95% confidence interval of [0.44, 7.28]. In contrast, as mentioned above, the mean point estimate of the East Asian markets is 0.36 with the mean 95% confidence interval of [0.13, 2.41] before the Asian crisis; the mean point estimate 0.39 with the mean confidence interval of [0.14, 2.89] after the crisis.[6] This indicates that real interest rates for industrialized countries overall show slower

[6] We have measured univariate half-life but the results are omitted to conserve space. Multivariate half-life estimates are remarkably smaller than univariate for industrialized countries. For East Asian countries the difference between univariate and multivariate estimates is not as dramatic and reserves interpretation. Given the existing evidence on real interest rate linkages in major financial markets, it seems that the interdependence across industrialized countries is captured in the difference between multivariate half-life and univariate estimates. In this sense multivariate half-life estimates provide more accurate measures for speed of convergence for western economies.

convergence than those for East Asian markets. This suggests a tendency that the real interest rates of more matured and liberalized markets show lower degree of mean-reversion.

Our conjecture for a possible cause of such tendency is financial frictions caused by government interventions. It has been long and well documented that East Asian markets experienced a high degree and a large scope of government interventions (see Masuyama et al, 1999). The effects of the intervention on the market still remain as a long-standing issue especially in relation to the East Asian financial crisis (see, for details, Corsetti, Pesenti and Roubini, 1998a, 1998b; and Wade, 1998a, 1998b). A plausible form of financial frictions for East Asian capital markets is financial repressions which include government controls on interest rates, entry, credit allocation, foreign exchange and taxation.[7] Differing views to the effects of financial repressions have been proposed since 1970s. The McKinnon-Shaw approach suggested that financial repression inhibits financial development, while Stiglitz (1994) and Amsden (1989, 2001) argued that some departures from strict market-oriented credit allocation were beneficial (see McKinnon, 1973; Shaw, 1973; Kapur, 1976; Fry, 1978).

Regardless of pros and cons of financial repression as a policy debate, the half-life estimates presented in this paper seem closely related to the degree of the financial friction for East Asia. Relatively large half-life estimates for liberalised markets in

[7] These restrictions continued to influence the markets mostly heavily at least until the early 1990s. See, for details, Greenwood (1986).

the Rapach-Wohar data imply that market forces drive real interest rates to the long-run mean level at a slower pace. The estimates for the East Asian markets are much smaller, which suggests that the rates deviate in a narrower range due to the financial repression.

The financial crisis dichotomy in mean-reversion for the East Asian markets provide stronger support for our conjecture in relation to the presence of financial friction in these markets when only emerging markets are considered. Excluding the two highly liberated markets, Japan and Singapore[8], the other members of the East Asian countries show pre-crisis mean half-life of 0.25, with 95% mean interval of [0.11, 1.18]. In the post-crisis period, however, their half-life estimates are much larger overall, with mean point estimate 0.43 with the 95% mean confidence interval of [0.15, 3.54]. We argue that the lower speed of mean-reversion in the post-crisis period is largely due to heavy deregulation-oriented remedy that East Asian emerging markets have commonly implemented. Korea, Indonesia, the Philippines and Thailand adopted either floating or managed floating exchange rate regime and drastically abolished capital controls over the crisis. In particular, Korea, Indonesia, the Philippines and Thailand were the major recipients of the IMF bailout packages with obligations to accept the IMF recommendations for structural reforms, and monetary and fiscal policies.[9] A key post-crisis change that these countries had to accommodate is the removal of

[8] Although it is difficult to make an exact list of emerging financial markets, market indices such as Morgan Stanley Emerging Markets Index include Korea, Malaysia, Philippines and Thailand as emerging markets. We follow this convention in the paper.
[9] Unlike the others who received support loan package, the Philippines was approved a major extension of existing credit.

control on foreign capital as part of market liberalisation. Our reasoning is that such market deregulation contributes to a decrease in the speed of mean-reversion of real interest rates. If markets are deregulated and unleashed from regulatory monetary authorities, external shocks caused by market forces to data generating process should drive the rates to larger deviations from and slower adjustments to their long-run levels than in a regulated environment. This is evidenced by the fact that half-life point estimates for these recipient countries have all increased over the crisis, with an exception of Indonesia. On the contrary, the point estimate for Malaysia, that refused the IMF intervention, has declined by almost half.

The lower rate of mean-reversion implies that real return from fixed income investment has a larger degree of unpredictable component, and that the investors may be subject to a higher degree of risk, particularly in the long term. Another inference that can be made from the comparison of half-life estimates between the industrialized and East Asian countries is that the participants in less repressed industrialized markets are taking larger risk premium on risk-free assets. A similar finding has been documented in Campbell and Viceira (2002, p78) on US real interest rates, who measured the degree of persistence for real interest rates and inflation and found that, since the early 1980s, aggressive monetary controls on inflation raised risk premia and volatility of returns on fixed income assets. Why long-horizon investors are willing to take such premium in developed countries is another interesting food for thoughts but left as a future research work.

5. Concluding remarks

This paper examined the mean-reversion property of the real interest rates from a number of industrialized and East Asian countries. Most of the previous studies have found the real interest rates to be non-stationary and mean-averting. This is a puzzling outcome since the real interest rate is expected to be stationary or mean-reverting over time. We note that these past studies used the unit root testing and half-life estimation methods which are univariate in nature, totally ignoring the presence of cross-sectional dependence among the real interest rates. Distinct from the past studies, we have used panel unit root tests and half-life estimation which take explicit account of cross-sectional dependence among the real interest rates. We have employed subsampling panel unit root tests Choi and Chue (2007) and estimated the half-life of real interest rates using the vector autoregressive models. The latter is based on the bias-corrected bootstrap HDR method, which is an extension of the univariate method recently proposed by Kim et al. (2007).

The major finding of the paper is that real interest rates are overall mean-reverting and stationary for both industrialized and East Asian countries. We also have found a strong tendency that the real interest rates of more matured and liberalized capital markets show slower rate of mean-reversion than those of East Asian markets. In addition, mean-reversion of real interest rates for East Asian emerging markets has become substantially slower after the Asian crisis. We suggest that government intervention in capital control causes financial frictions which prevent market forces from exerting full effects on the real interest rates. The rates deviate in a narrower range and revert to the long-run value at a faster rate in the presence of such friction. We note that this point needs a further investigation in a future research.

Table 1: Unit Root Test Results for Rapach and Wohar data

	Test Statistic	5% Subsampling Critical Values	
		MV	SC
IPS	-2.62	-2.43	-2.28
Inverse Normal	-6.53	-5.24	-5.49

MV: minimum volatility rule for subsample length selection
SC: stochastic calibration rule for subsample length selection

Table 2: Unit Root Test Results for East Asian data

	Test Statistic	5% Subsampling Critical Values	
		MV	SC
Period I			
IPS	-3.54	-2.67	-2.46
Inverse Normal	-5.02	-4.34	-4.92
Period II			
IPS	-3.17	-2.70	-2.41
Inverse Normal	-5.23	-4.29	-4.91

MV: minimum volatility rule for subsample length selection
SC: stochastic calibration rule for subsample length selection
Period I: 1987:01 to 1996:12; Period II: 1999:01 to 2007:05

Table 3: Half-Life Estimates for Rapach and Wohar data (in years)

	Point		95% Interval			
	RW	HDR	RW		HDR	
Australia	4.29	1.24	2.08	∞	0.52	29.33
Belgium	4.38	1.12	2.49	∞	0.45	6.90
Canada	3.18	0.52	1.56	∞	0.30	1.16
Denmark	0.92	0.50	0.88	∞	0.32	0.95
France	8.73	2.00	4.43	∞	0.87	17.87
Ireland	3.11	0.38	1.42	∞	0.20	0.75
Italy	4.15	0.70	1.60	∞	0.33	2.35
Netherlands	0.99	0.67	0.93	∞	0.37	1.38
New Zealand	4.57	1.01	1.57	∞	0.45	2.65
Norway	3.86	0.67	1.28	∞	0.29	1.79
Switzerland	1.52	1.95	1.19	∞	0.62	8.15
UK	2.36	0.60	1.51	∞	0.38	0.92
US	3.04	1.31	2.45	∞	0.67	20.50
Mean	3.47	0.97	1.80	∞	0.44	7.28

RW: the grid bootstrap estimates of Rapach and Wohar (2004)
HDR: Bias-corrected HDR Bootstrap estimates

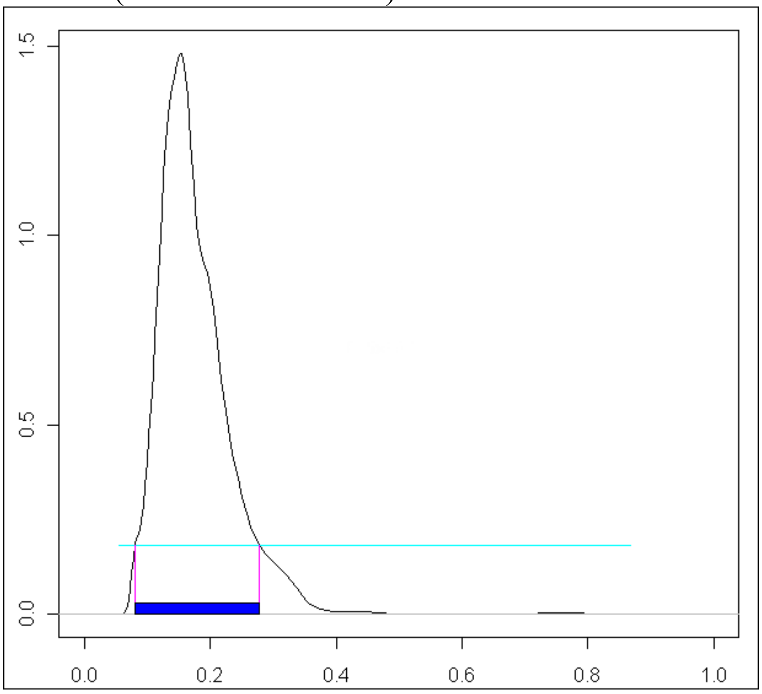
Table 4: Half-Life Estimates for East Asian Markets (in years)

	Period I			**Period II**		
	Point	**95% Interval**		**Point**	**95% Interval**	
JPN	0.92	0.24	9.27	0.30	0.12	0.99
KOR	0.31	0.13	1.24	0.32	0.14	5.00
SNG	0.31	0.14	1.72	0.31	0.11	1.56
IND	0.16	0.09	0.43	0.12	0.08	0.16
MAL	0.57	0.18	3.83	0.24	0.10	1.46
PHP	0.07	0.05	0.10	0.48	0.17	3.12
THI	0.15	0.08	0.27	0.99	0.25	7.94
Mean	0.36	0.13	2.41	0.39	0.14	2.89
Mean*	0.25	0.11	1.18	0.43	0.15	3.54

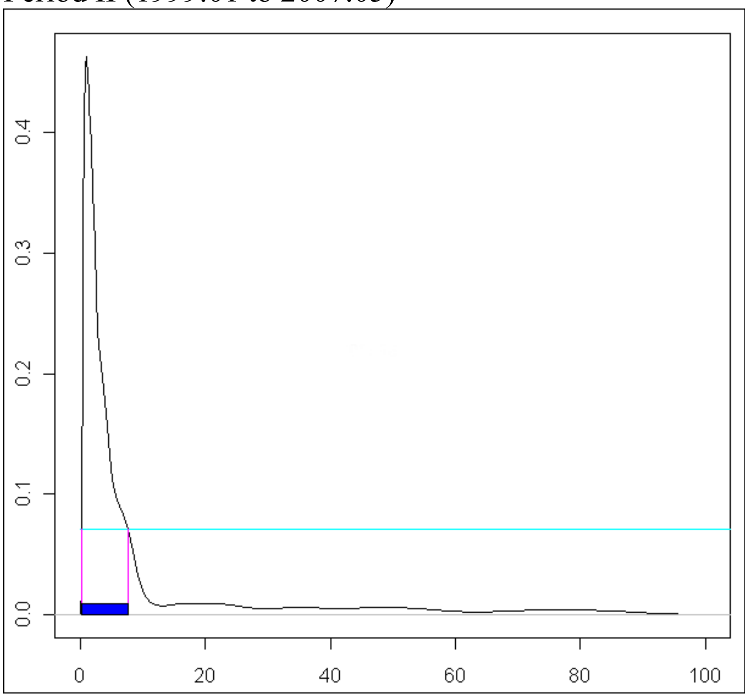
Period I: 1987:01 to 1996:12; Period II: 1999:01 to 2007:05
The entries are bias-corrected bootstrap HDR estimates
Mean indicates the average for all countries, while Mean* indicates the average excluding Japan and Singapore

Figure 1: Density Estimates for the Bias-Corrected Bootstrap Replicates of Half-Life: the case of Thailand

Period I (1987:01 – 1996M12)



Period II (1999:01 to 2007:05)



The X-axes of both graphs indicate half-life in years

References

Amsden, A.H. 2001, *The rise of "the rest": Challenges to the West from late-industrializing economies,* Oxford University Press, Oxford and New York.

Amsden, A.H. 1989, *Asia's next giant: South Korea and late industrialization,* Oxford University Press, New York and Oxford.

Black, F. & Scholes, M.S. 1973, "The Pricing of Options and Corporate Liabilities", *Journal of Political Economy,* vol. 81, no. 3, pp. 637-654.

Campbell, J. Y. and Viceira, L. M. 2002 , *Strategic Asset Allocation: Portfolio Choice for Long-Term Investors*, Oxford University Press, New York and Oxford

Caporale, G. M., Cerrato, M. and Spagnolo, N. 2005. Measuring half-lives: using a non-parametric bootstrap approach. Applied Financial Economics Letters 1, 1–4.

Choi, I. 2001, "Unit Root Tests for Panel Data", *Journal of International Money and Finance,* vol. 20, no. 2, pp. 249-272.

Choi, I. & Chue, T.K. 2007, "Subsampling Hypothesis Tests for Nonstationary Panels with Applications to Exchange Rates and Stock Prices", *Journal of Applied Econometrics,* vol. 22, no. 2, pp. 233-264.

Corsetti, G., Pesenti, P. & Roubini, N. 1998a, *What Caused the Asian Currency and Financial Crisis? Part I: A Macroeconomic Overview*, National Bureau of Economic Research, Inc, NBER Working Papers.

Corsetti, G., Pesenti, P. & Roubini, N. 1998b, *What Caused the Asian Currency and Financial Crisis? Part II: The Policy Debate*, National Bureau of Economic Research, Inc, NBER Working Papers.

Efron, B. and R. J. Tibshirani (1993). An Introduction to the Bootstrap. New York, Chapman & Hall.

Fama, E.F. & French, K.R. 1988, "Dividend Yields and Expected Stock Returns", *Journal of Financial Economics,* vol. 22, no. 1, pp. 3-25.

Fisher, I. 1930. *The Theory of Interest.* New York: Macmillan.

Fisher, R. A. 1932. *Statistical Methods for Research Workers*, 4th Edition, Edinburgh, Oliver and Boyd.

Fry, M.J. 1978, "Money and Capital or Financial Deepening in Economic Development?", *Journal of Money, Credit, and Banking,* vol. 10, no. 4, pp. 464-475.

Goncalves, S. & Kilian, L. 2004, "Bootstrapping Autoregressions with Conditional Heteroskedasticity of Unknown Form", *Journal of Econometrics,* vol. 123, no. 1, pp. 89-120.

Gospodinov, N. 2004, "Asymptotic Confidence Intervals for Impulse Responses of Near-Integrated Processes", *Econometrics Journal,* vol. 7, no. 2, pp. 505-527.

Hansen, B. E. (1999). "The Grid Bootstrap and the Autoregressive Model." *The Review of Economics and Statistics* **81**: 594-607.

Hansen, L.P. & Singleton, K.J. 1982, "Generalized Instrumental Variables Estimation of Nonlinear Rational Expectations Models", *Econometrica,* vol. 50, no. 5, pp. 1269-1286.

Hyndman, R.J. (1996) "Computing and graphing highest density regions", *American Statistician*, 50, 120-126.

Im, K.S., Pesaran, M.H. & Shin, Y. 2003, "Testing for Unit Roots in Heterogeneous Panels", *Journal of Econometrics,* vol. 115, no. 1, pp. 53-74.

Islam, S.S. & Mozumdar, A. 2007, "Financial Market Development and the Importance of Internal Cash: Evidence from International Data", *Journal of Banking and Finance,* vol. 31, no. 3, pp. 641-658.

Kapur, B.K. 1976, "Alternative Stabilization Policies for Less-developed Economies", *Journal of Political Economy,* vol. 84, no. 4, pp. 777-795.

Kilian, L. 1998, "Small-Sample Confidence Intervals for Impulse Response Functions", *Review of Economics and Statistics,* vol. 80, no. 2, pp. 218-230.

Kim, J. H., Silvapulle, P. and Hyndman, R, 2007, "Half-Life Estimation based on the Bias-Corrected Bootstrap: A Highest Density Region Approach", *Computational Statistics and Data Analysis*, 51, 7, 3418-3432.

Lai, K.S., 2004. On structural shifts and stationarity of the ex ante real interest rate. *International Review of Economics and Finance*, 13 (2), 217-228.

Lai, K. S., 2008, The Puzzling Unit Root in the Real Interest Rate and Its Inconsistency with Intertemporal Consumption Behavior, *Journal of International Money and Finance,* 27, 140-155.

Lucas, R.E., Jr 1978, "Asset Prices in an Exchange Economy", *Econometrica,* vol. 46, no. 6, pp. 1429-1445.

Maddala, G. S. and Kim, I.-M., 1998, Unit roots, coinegration and structural change, Cambridge University Press.

Maddala, G.S. & Wu, S. 1999, "A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test", *Oxford Bulletin of Economics and Statistics,* vol. 61, pp. 631-652

Masuyama, S., Vandenbrink, D. & Yue, C. S. 1999, *East Asia's Financial Systems*, Institute of Southeast Asian Studies and Nomura Research Institute, Singapore and Tokyo

McKinnon, R.I. 1973, *Money and Capital in Economic Development, Brookings Institution*, Washington DC

Mehra, R. & Prescott, E.C. 1985, "The Equity Premium: A Puzzle", *Journal of Monetary Economics,* vol. 15, no. 2, pp. 145-161.

Murray, C.J. & Papell, D.H. 2002, "The Purchasing Power Parity Persistence Paradigm", *Journal of International Economics,* vol. 56, no. 1, pp. 1-19.

Murray, C.J. & Papell, D.H. 2005a, "Do Panels Help Solve the Purchasing Power Parity Puzzle?", *Journal of Business and Economic Statistics,* vol. 23, no. 4, pp. 410-415.

Murray, C.J. & Papell, D.H. 2005b, "The Purchasing Power Parity Puzzle Is Worse Than You Think", *Empirical Economics,* vol. 30, no. 3, pp. 783-790.

Phillips, P.C.B. 2005, "Econometric Analysis of Fisher's Equation", *American Journal of Economics and Sociology,* vol. 64, no. 1, pp. 125-168.

Poterba, J.M. & Summers, L.H. 1988, "Mean Reversion in Stock Prices", *Journal of Financial Economics,* vol. 22, no. 1, pp. 27-59.

Rapach, D.E. & Weber, C.E. 2004, "Are Real Interest Rates Really Nonstationary? New Evidence from Tests with Good Size and Power", *Journal of Macroeconomics,* vol. 26, no. 3, pp. 409-430.

Rapach, D.E. & Wohar, M.E. 2004, "The Persistence in International Real Interest Rates", *International Journal of Finance and Economics,* vol. 9, no. 4, pp. 339-346.

Rogoff, K. 1996, "The Purchasing Power Parity Puzzle", *Journal of Economic Literature,* vol. 34, no. 2, pp. 647-668.

Romano, J.P. & Wolf, M. 2001, "Subsampling Intervals in Autoregressive Models with Linear Time Trend", *Econometrica,* vol. 69, no. 5, pp. 1283-1314.

Rose, A.K. 1988, "Is the Real Interest Rate Stable?", *Journal of Finance,* vol. 43, no. 5, pp. 1095-1112.

Rossi, B. 2005, "Confidence Intervals for Half-Life Deviations from Purchasing Power Parity", *Journal of Business and Economic Statistics,* vol. 23, no. 4, pp. 432-442.

Shaw, E. 1973, *Financial Deepening in Economic Development*, Oxford University Press, New York.

Sheather, S. J., and M. C. Jones, 1991, “A Reliable Data-based bandwidth Selection Method for kernel Density Estimation”, *Journal of the Royal Statistical Society,B* 53, 683-690.

Siegel, J.J. & Thaler, R.H. 1997, "Anomalies: The Equity Premium Puzzle", *Journal of Economic Perspectives,* vol. 11, no. 1, pp. 191-200.

Solow, R.M. 1991 [1956], "A Contribution to the Theory of Economic Growth" in *Growth theory. Volume 1. Descriptive growth theories*, eds. R. Becker & E. Burmeister eds, Elgar, International Library of Critical Writings in Economics, no. 10 Aldershot, U.K. and Brookfield, Vt., pp. 3-32.

Stiglitz, J.E. 1994, "The Role of the State in Financial Markets" in *Proceedings of the World Bank Annual Conference on Development Economics, 1993: Supplement to The World Bank Economic Review and The World Bank Research Observer*, eds. M. Bruno & B. Pleskovic eds, World Bank, Washington, D.C., pp. 19-52.

Wade, R. 1998a, "From 'Miracle' to 'Cronysim': Explaining the Great Asian Slump", *Cambridge Journal of Economics,* vol. 22, no. 6, pp. 693-706.

Wade, R. 1998b, "The Asian Debt-and-Development Crisis of 1997-?: Causes and Consequences", *World Development,* vol. 26, no. 8, pp. 1535-1553.

Weil, P. 1989, "The Equity Premium Puzzle and the Risk-Free Rate Puzzle", *Journal of Monetary Economics,* vol. 24, no. 3, pp. 401-421.