THE UNIVERSITY OF MELBOURNE

Melbourne Institute Working Paper Series

Working Paper No. 14/05

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MELBOURNE INSTITUTE
of Applied Economic and Social Research

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Melbourne Institute Working Paper No. 14/05

ISSN 1328-4991 (Print)

ISSN 1447-5863 (Online)

ISBN 0 7340 3192 0

September 2005

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Abstract

This paper tests for the presence of nonlinear dynamics in selected Asian short rates and employs a regime varying unit root test to detect non-stationarity for distinct regimes. Nonlinearities in the form of Markov-switching dynamics are found in all short rates sample. The mean-reverting behaviour of interest rates is dependent on both the level and volatility of interest rates. The occasional random walk and mean-reverting dynamics of short rates are attributed to the macroeconomic fundamentals, exchange rate regimes and monetary policy objectives in these economies.

Keywords: Unit root; Markov-switching; Interest rates; Non-linearity
J.E.L. Reference Numbers: C12; G12; E44

1 Introduction

Testing for the stationarity properties of the short-term interest rate has been a major focus of empirical finance and macroeconomics because of its implications for asset pricing as well as monetary policy. Traditionally theoretical literature on the stochastic process of short-term interest rates adopts a linear and mean-reverting drift (see Marsh and Rosenfeld,1983; Chan et al.,1992; Cox et al.,1985; and Brennan and Schwartz 1980). However, the assumption of short rate stationarity in theoretical models has been challenged by empirical studies. Short-rate diffusion models estimated by Marsh and Rosenfeld (1983), Chan et al. (1992), Tse (1995), Aquila et al. (2003), and Brenner et al. (1996) *inter alia* documented evidence that short-term interest rates follow a random walk process. The lack of consensus on short-rate mean reversion between empirical and theoretical models could be a result of failing to account for possible nonlinear dynamics in the short rate process.

This paper contributes to this discussion by using a regime varying unit root test to provide evidence about the time-series properties of the short rate. The methodology is motivated by recent contributions in the term structure literature that have employed regime shifts and/or nonlinear drift specification to model structural breaks and nonlinearity in the spot interest rate (see Gray, 1996; Bekaert et al., 1997; Ait-Sahalia, 1996a,b and Bali and Wu, 2005). The class of interest rate models that accommodates regime changes in both its autoregressive parameters and variance shows that short rates are mean-reverting when their levels and volatility are high, but they are nonstationary during periods of low interest rates and low volatility (see Ang and Bekaert (1998), Gray (1996) and Bekaert et al. (1997), *inter alia*). The occasional random walk behaviour of the short rate implies that the standard ADF test may have low power in detecting stationarity in short rates. Thus incorporating this regime switching behaviour in the unit root test procedure may yield reliable inference about the stationarity properties of short rates.

Our approach proceeds in two stages. First, we conduct the test developed by Cheung and Erlandsson (2005) to detect the presence of Markov-switching dynamics. Cheung and Erlandsson (2005) argued that long swings found in realizations from a near unit root process may be mistaken for Markov-switching dynamics thus they recommend preliminary testing of the series. IOnce regime-switching dynamics are identified, the second stage employs a Markov-Switching Augmented Dickey Fuller (MSADF) test of Kanas and Genius (2005) to determine the stationarity of interest rates for distinct regimes. The MSADF test generalises the Augmented Dickey-Fuller (hereafter ADF) unit root test using Hamilton's (1989,1990) Markov-switching model and admits the ADF regression parameters and the variance of interest rate to take different values for different regimes. In so doing, the test improves its power to detect stationarity when the process possesses occasional random walk characteristics and experiences a shift in variance (see Kanas and Genius, 2005).

The second contribution of the present paper is in the choice of our sample which comprises selected East Asian economies short rates. Our results expand on the literature that is predominantly based upon U.S. short rates. The question whether short-rate mean reversion depends on its levels and volatility is particularly relevant for the East Asian economies given the dramatic rise and fall in interest rates in the recent South-East Asian financial crisis and the high inflation episodes experienced in economies like Indonesia and the Philippines. Furthermore, differences in the monetary policy objectives in these economies may also contribute to fundamentally different short rate dynamics. For example, Singapore, being a small and open economy with high import content of her domestic expenditures, adopts a monetary policy that is centred on management of the trade-weighted exchange rate rather than traditional monetary policy instruments such as money supply and interest rates. Hong Kong's monetary policy, on the other hand, is tied to maintaining the nominal exchange rate linked to the U.S. dollar.

The results of our analysis suggest that the sample of Asian short rates exhibit Markov-switching behaviour. When these regime switching dynamics are accounted for in the unit root test, there is evidence that Asian short rates revert to some long-run mean when the levels and volatility are high. However, during periods when interest rate levels and volatility are low they behave more like a random walk process. While these findings are consistent with the literature that employs regime switching short rate models, our results for Singapore and the Philippines show that the short-rate dynamics need not follow an occasionally random walk process. Singapore's short rate is nonstationary in both regimes which may be a result of her exchange rate centred monetary policy and stable macroeconomy which have kept interest rates at low levels with little volatility. In contrast, the Phillipines economy was fraught with instability caused by conflicting objectives in the monetary policy, and poor financial and macroeconomic management that resulted in several episodes of devaluation in the Peso. These events may have contributed to the high and volatile short rates in both regimes thereby causing it to mean-revert.

This paper is organised as follows. Section 2 gives a literature review on nonlinear dynamics in short-term interest rates. Section 3 provides a preliminary description of the data and motivates the need for modelling regime shifts. Section 4 tests for the presence of Markov-switching dynamics and unit roots, and interprets the results. Section 5 discusses the sources of (non)stationary in Asian short-term rates. Section 6 summarizes and concludes.

2 Literature Review

The Chan, Karolyi, Longstaff and Sanders (1992) (CKLS, hereafter) general nonlinear process for short-term interest rates, $\{r_t, t \geq 0\}$, written as

$$dr = (\mu + \lambda r)\, dt + \phi r^{\delta} dW \tag{1}$$

is widely estimated in the empirical literature. Here r represents the level of the short-term interest rate, W is a Brownian motion and μ, λ and δ are parameters. The linear drift component of short-term interest rates is captured by $\mu + \lambda r$ while the variance of unexpected changes in interest rates equals $\phi^2 r^{2\delta}$. The estimate $\lambda < 0$ if significant suggests that the short rate is mean-reverting. The CKLS model is popular because it nests many of the existing interest rate models by placing restrictions on δ. For example, when $\delta = 0$ then (1) reduces to the Vasicek (1977) model, while $\delta = 1/2$ yields the Cox, Ingersoll and Ross (1985) model, see Chan et al. (1992) for further details. The empirical results on the mean-reverting properties of short rates, however, are mixed. Using the generalised method of moments, CKLS report a non-significant λ estimate for the one-month U.S. Treasury bill yields. Brenner, Harjes and Kroner (1996), on the other hand, show that the one-month Treasury bills mean revert. By using robust generalised method of moments which improve on previous studies methodology, Aquila et al. (2003) show that the non mean-reverting characteristic of the short rate is a robust feature of the one-month U.S. Treasury bill yields. The results are largely unaffected even after controlling for a cluster of influential data points that is due to the Fed monetary experiment between 1979 and 1982.

Recent articles in the term structure literature have relied on regime switching techniques to model structural breaks and nonlinearity in the spot interest rate process. The earlier work of Sanders and Unal (1988) shows that the U.S. short rate model switches regime three times over the sample period from March 1959 to December 1985 with two of the switches coinciding with the 1979 and 1982 changes in the Federal Reserve monetary policy. These regime changes may potentially cause the parameter estimate of the mean reversion coefficient for a single regime short rate model to be unstable over the whole sample period. Hamilton (1988, 1989), Cecchetti et al. (1993) and Garcia and Perron (1996) employ a Markov switching model that accommodates possible stochastic changes in the regime of interest rates conditional mean. Their models assume interest rate levels as a function of an intercept that is allowed to shift between regimes. By accounting for regime changes, the Markov switching model better represents the univariate process for interest rates. Gray (1996) incorporates regime shifts in both the conditional mean and variance of the U.S. three month Treasury Bill rates. His empirical results confirm that these generalisations are statistically and economically significant. Moreover, he shows that the U.S. short rates display random walk behaviour when it is in low-volatility state and mean-revert when it is high-volatility state.

The evidence that mean-reversion occurs during high and volatile short rates have also been documented in studies that employ a nonlinear drift specification in short rate models. Ait-Sahalia (1996a,b) using a general specification test of a short rate diffusion model finds that the test rejects a linear drift in favour of nonlinear models. Using a nonparametric approach, Stanton (1997) and Jiang (1998) find similar nonlinearity in the drift function on the 3-month U.S. Treasury bills. Conley et al. (1997) also find evidence of stronger mean reversion for both very high and low levels of the federal fund rates. Bali and Wu (2005) show that the linear drift specification is strongly rejected for the U.S. overnight federal funds rate after controlling for the conditional variance specification using GARCH volatility and non-normal interest rate innovation.

Various economic intuitions to explain short rate adjustment to its long-term mean when it is high and volatile have been proposed. Ait-Sahalia (1996a) alluded to a speculative interpretation whereby market participants anticipate a response from the central bank by intervening in the money market and restoring the level of interest rates to its long-run level. Bali and Wu (2005) adopts a similar view about interest rates adjustment to its long-term mean when it reaches historical highs such as the period between 1979 and 1982. Often high interest rates are a result of the federal reserve taking drastic measure to combat high inflation and shortly after the short-term interest rate adjusts downward to bring about mean-reversion.

In summary, the substantive body of literature evidencing regime shifts and nonlinearity in interest rates suggest that one may have to exercise caution when drawing inference on short-rate stationarity using the t-test that is associated with the mean reversion coefficient of a single regime specification, λ. Additionally, standard unit root tests may be misleading when the process faces structural breaks (Perron,1989; Zivot and Andrews, 1992; Lumsdaine and Papell, 1997; Leybourne et al.,1998) or shifts in variance (Hamori and Tokihisa, 1997). Furthermore, Nelson et al. (2001) show that standard unit root tests have low power when applied to interest rate models with Markov-switching autoregressive parameters where they exhibit mean-averting behaviour in one regime and mean-reverting characteristic in the other.

3 Data and Empirical Facts

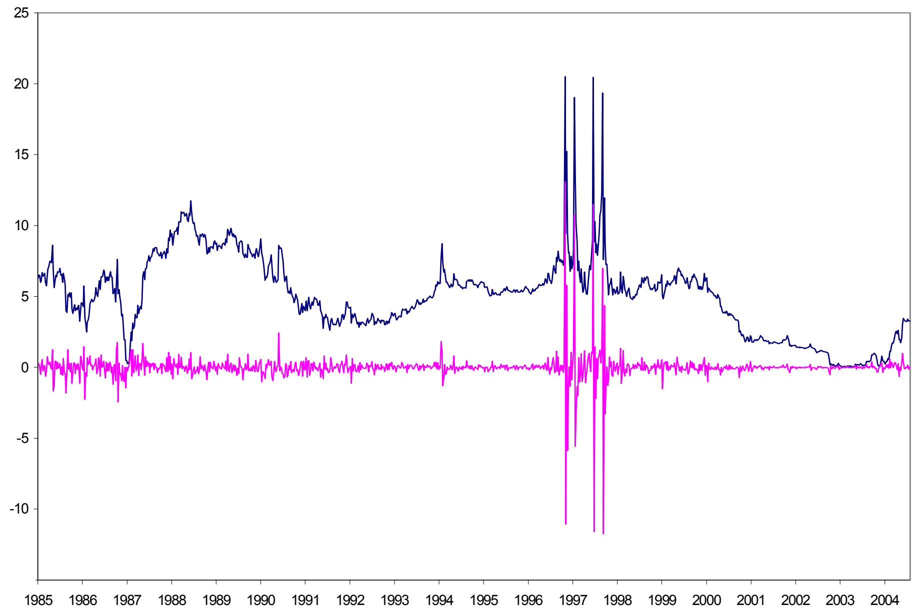
The data are a set of weekly 1-month interbank rates for Hong Kong (HK), Singapore (SI), Thailand (TH), Indonesia (ID) and the Philippines (PH). The sample covers the periods from 30 December 1985 to 25 July 2005 for HK, 8 January 1988 to 22 July 2005 for SG, 6 January 1992 to 30 June 2003 for TH, 30 November 1987 to 25 July 2005 for ID and 19 January 1990 to 29 July 2005 for PH. The data for HK, TH, ID and PH are obtained from DataStream while SI data are obtained from the Monetary Authority of Singapore[1]. Figure 1 plots the levels and the change in levels of the five Asian short rates. Visual inspection of the plots (see Figure 1) shows that there are variations in the level of interest rates across the sample period, but the dramatic rise in interest rate levels is most noticeable during the Asian financial crisis which started from the second half of 1997. In addition, in periods when the interest rate levels are high, short rates are observed to be more volatile. This is particularly true with the onset of the Asian financial crisis from July 1997 where short rate volatility (that is proxied by the change in short rate) rises substantially. The data plot also reveals that Thailand and the Philippines short rates are relatively more volatile when compared with the other short rates prior to the crisis.

Table 1 reports a wide range of descriptive statistics of the five short rate series for the full sample and the two subperiods, namely pre- and post-1997 crisis. Over the whole sample, the mean of all short rate change, apart from the Philippines, are negative. The impact of the 1997 crisis on the mean and variance of the series is apparent. The mean of short rate change changes in sign and magnitude between the pre-crisis and post-crisis periods. In the post-crisis sample, the volatility of short rate change in some of the East Asian economies, with the exception of Thailand and the Philippines, has experienced approximately a four-fold increase compared with the pre-crisis sample. Thailand and the Philippines short rate volatility fell by more than half in the post-crisis period. The short rate distributions are also heavily skewed and leptokurtic.

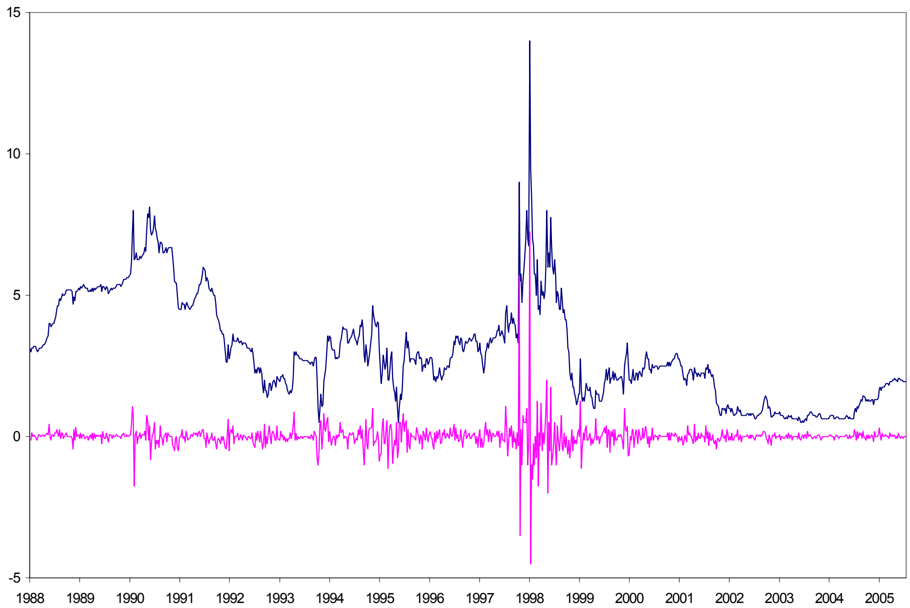
Based on the Ljung-Box statistics, there is evidence of fifth order serial correlation for all five series. Serial correlation is also found in all short rates within the two subsamples. The null hypothesis of homoskedasticity is rejected in all series for the whole sample and the two subsamples.

[1] https://secure.mas.gov.sg/frames/dataroom

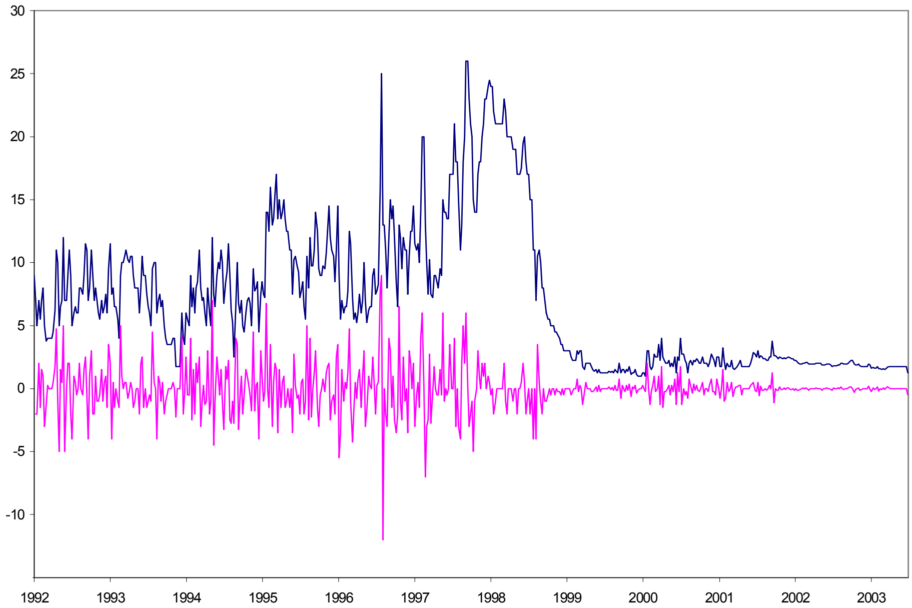
Figure 1: Plots of Asian Short Rates in Levels (upper) and Change (lower)



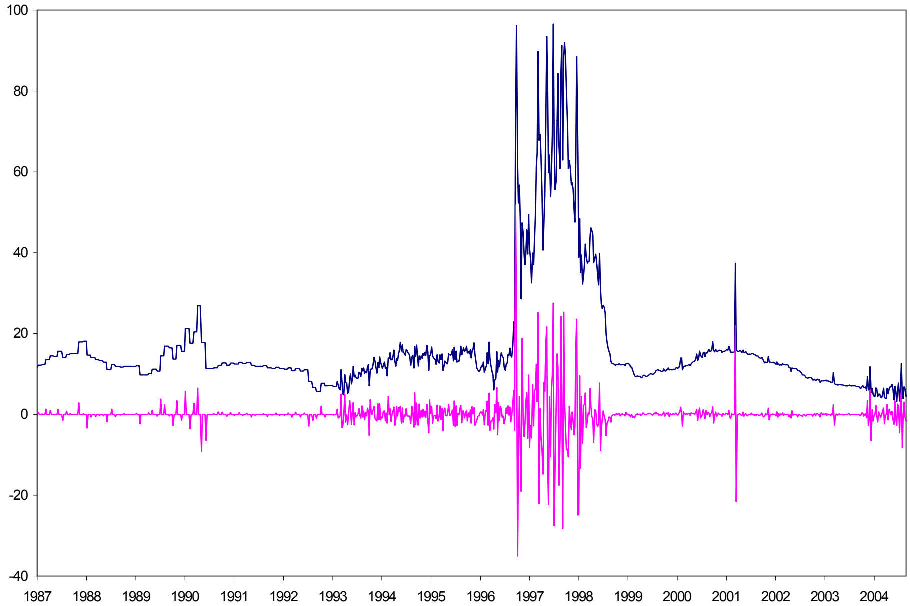
Hong Kong



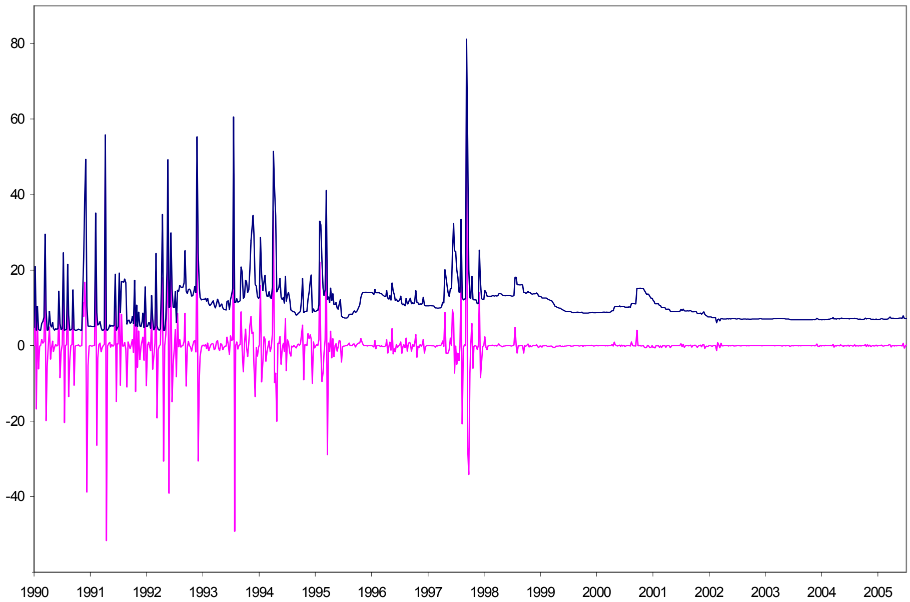
Singapore



Thailand



Indonesia



Philippines

Preliminary unit root tests are performed to determine the stationarity property of these short rate series. For robustness, two different types of unit root tests, namely the Augmented Dickey Fuller test and the Phillips and Perron test (1988) (PP test hereafter) are employed. The optimal lag length is chosen based on the Schwarz Information Criteria (SIC) to ensure that the residual of the unit root regression is free from serial correlation. The results are reported in Table 2. Based on the whole sample, the null of unit root fails to be rejected in all short rates except for Indonesia and the Philippines. The unit root tests results in the two subsamples suggest that short rate dynamics in three of the five economies have changed since the 1997 crisis. In particular, short rates in Thailand, Indonesia and the Philippines are no longer stationary after the crisis. It is noteworthy that the results of the unit root tests are only indicative of a change in the dynamics of short rates as they may be sensitive to the pre-determined break point which was chosen to be July 17, 1997. A better approach would be to endogenise the break point by allowing the data to determine the timing of the shifts in both the

Table 1: Summary Statistics of Δr

Country	Statistics	Full sample	Sub-samples	
		1985:12-2005:7	1985:12-1997:7	1997:8-2005:7
Hong Kong	Mean	-0.0029	0.0006	-0.0069
	Std. dev	1.0792	0.4434	1.6012
	Skewness	0.4866	-0.0686	0.3730
	Kurtosis	83.1963	8.8892	43.2543
	JB	294209	870.4388	28231
	Q(5)	254.5913*	15.6951*	125.1238*
	ARCH(5)	266.6446*	42.1368*	106.7443*
		1988:1-2005:7	1988:1-1997:7	1997:8-2005:7
Singapore	Mean	-0.0013	0.0030	-0.0058
	Std. dev	0.4709	0.2762	0.6289
	Skewness	4.4397	-0.5024	4.1489
	Kurtosis	97.2199	8.6938	68.4999
	JB	363352	692.2849	75920.97
	Q(5)	94.2419*	11.9740*	71.9819*
	ARCH(5)	115.3167*	36.6285*	51.5286*
		1992:1-2003:6	1992:1-1997:7	1997:8-2003:6
Thailand	Mean	-0.0091	0.0311	-0.0442
	Std. dev	1.6831	2.4373	0.8400
	Skewness	0.2078	0.1066	0.2030
	Kurtosis	7.7933	5.5814	19.7281
	JB	1791	80.7897	4876.606
	Q(5)	47.6248	27.7186	63.0231
	ARCH(5)	110.1106	30.5828	164.2881
		1987:11-2005:7	1987:11-1997:7	1997:8-2005:7
Indonesia	Mean	-0.0058	0.0033	-0.0202
	Std. dev	4.3796	1.4296	6.6242
	Skewness	1.2520	0.0210	0.8885
	Kurtosis	37.9519	12.3376	19.0093
	JB	61421	2187	4518
	Q(5)	121.9004*	58.3176*	54.9875*
	ARCH(5)	236.9098*	65.3536*	177.4096*
		1990:1-2005:7	1990:1-1997:7	1997:8-2005:7
Phillippines	Mean	0.0036	0.0534	-0.0601
	Std. dev	7.0823	9.1295	4.3818
	Skewness	0.8172	-0.1204	7.6115
	Kurtosis	29.4093	15.0915	160.8377
	JB	29317	2395	437932
	Q(5)	100.8066*	53.4997*	105.1064*
	ARCH(5)	87.0993*	82.4888*	33.5749*

Note: Q(5) is the LjungBox statistics with ten lags, ARCH(5) is the Lagrange multiplier heteroskedasticity test and JB is the Jarque-Bera normality test.

unconditional mean and variance of short rate change. This is the approach taken by the MSADF test discussed in the next section.[2]

[2]Unit root test procedures that endogenously allow for the possibility of a break in the data generating process such as Zivot and Andrews (1992) test is not considered here because the test does not admit changes in the variance of the process.

Table 2: Unit Root Tests Results for r

Country	Statistics	Full sample 1985:12-2005:7	Sub-samples 1985:12-1997:7	1997:8-2005:7
Hong Kong	ADF	-3.6379*	-2.3158	-1.8251
	PP	-4.9717*	-2.3463	-1.5065
		1988:1-2005:7	1988:1-1997:7	1997:8-2005:7
Singapore	ADF	-2.1007	-2.0412	-1.8708
	PP	-2.0364	-2.0987	-2.1250
		1992:1-2003:6	1992:1-1997:7	1997:8-2003:6
Thailand	ADF	-2.0483	-5.9736*	-2.1213
	PP	-2.2308	-5.6640*	-2.2994
		1987:11-2005:7	1987:11-1997:7	1997:8-2005:7
Indonesia	ADF	-2.5960***	-4.5054*	-2.0108
	PP	-5.1962*	-5.2768*	-2.2378
		1990:1-2005:7	1990:1-1997:7	1997:8-2005:7
Philippines	ADF	-4.9370*	-13.2982*	-1.8143
	PP	-20.4175*	-13.6745*	-1.9932

Note: The critical values for the ADF and Phillips-Perron tests with intercept at 1%, 5% and 10% significance levels are -3.43, -2.86, and -2.56. * and *** denote significance at 1% and 10% levels respectively.

In summary, the plot of the data, the descriptive statistics and the unit root tests results suggest that the empirical distributions and the underlying dynamics of the short rates have been influenced by the structural change which occurred in 1997.

4 Testing for Markov-Switching Dynamics and Unit Roots

Based on the statistical motivation outlined above, we incorporate switches in the data generating process of the short-term rate. We do this by modeling the evolution of interest rate as a Markov regime switching process. Using the discretized short-rate diffusion model of Vasicek (1977)

$$\Delta r_t = \mu + \lambda r_{t-1} + e_t \tag{2}$$

where μ is the intercept, λ is the mean-reverting coefficient and e_t is a sequence of independent and identically distributed random variables drawn from the distribution $(0, \sigma^2)$, we formulate the short rate process to have the features of mean reversion and different variance within each regime. The Markov-switching Vasicek model is

$$\begin{aligned}
\Delta r_t &= \mu_t + \lambda_t r_{t-1} + e_t \\
\mu_t &= \mu_0 s_t + \mu_1 (1 - s_t) \\
\lambda_t &= \lambda_0 s_t + \lambda_1 (1 - s_t) \\
e_t &\sim iid(0, \sigma_t^2) \\
\sigma_t^2 &= \sigma_0^2 s_t + \sigma_1^2 (1 - s_t)
\end{aligned} \tag{3}$$

where $s_t \in \{0, 1\}$ is the first-order Markov chain state variable which governs the time-varying nature of the parameters with the following transition matrix

$$\begin{aligned}
P &= \begin{bmatrix} p(s_t = 0 | s_{t-1} = 0) & p(s_t = 1 | s_{t-1} = 0) \\ p(s_t = 0 | s_{t-1} = 1) & p(s_t = 1 | s_{t-1} = 1) \end{bmatrix} \\
&= \begin{bmatrix} p_{00} & p_{01} \\ p_{10} & p_{11} \end{bmatrix}.
\end{aligned} \tag{4}$$

Here p_{ij} denotes the transition probability of $s_t = j$ given that $s_{t-1} = i$ for $i, j = 0, 1$. The transition probabilities satisfy the condition $p_{i0} + p_{i1} = 1$ for $i = 0, 1$. σ_i^2 denotes the interest rate process variance for the two regimes $i = 0, 1$.[3]

As highlighted in the descriptive statistics for the whole sample, all short rates display volatility clustering. A superior characterisation of the short rate requires that the model specifies such a feature using the generalised autoregressive conditional heteroskedasticity (GARCH) specification. Nevertheless, given that the objective of the current paper is to identify the stationarity properties of interest rates and not in modelling the conditional volatility dynamics, we do not estimate a Markov-switching GARCH model.

[3] In theory, one can allow for more regimes. For our purpose, we do not impose an *ad hoc* number of regimes. Instead, we perform a formal test to determine the number of regimes in the data.

4.1 Testing for Markov-switching Dynamics

Testing for Markov-switching dynamics in a short rate model is a difficult task because conventional testing approaches are not applicable due to the presence of unidentified nuisance parameters under the null of linearity. The presence of the nuisance parameters gives the likelihood surface sufficient freedom so that one is unable to reject the possibility that the apparently significant parameters could arise from sampling variation. The scores associated with the parameters of interest under the alternative may be identically zero under the null. To overcome the problem of testing in the presence of unidentified nuisance parameters, Davies (1977, 1987) derived an upper bound for the significance level of the likelihood ratio test statistic under the nuisance parameters. Hansen (1992, 1996) and Garcia (1998) proposed formal tests of the Markov-switching model against linear alternative employing standardized likelihood ratio test designed to deliver asymptotically valid inference. Nevertheless, these methods are computationally demanding.

Cheung and Erlandsson (2005) document the importance of formally testing for the presence of Markov switching dynamics. They also propose a simpler method of testing for it. They show that when fitting a Markov switching model to a (near) unit root process, the persistence in unit root can potentially give rise to spurious results and lead to misidentification of long swings as regime-switching behaviour. We avoid spurious identification of markov-switching dynamics by conducting their test that is set up with the following null and alternative hypotheses:

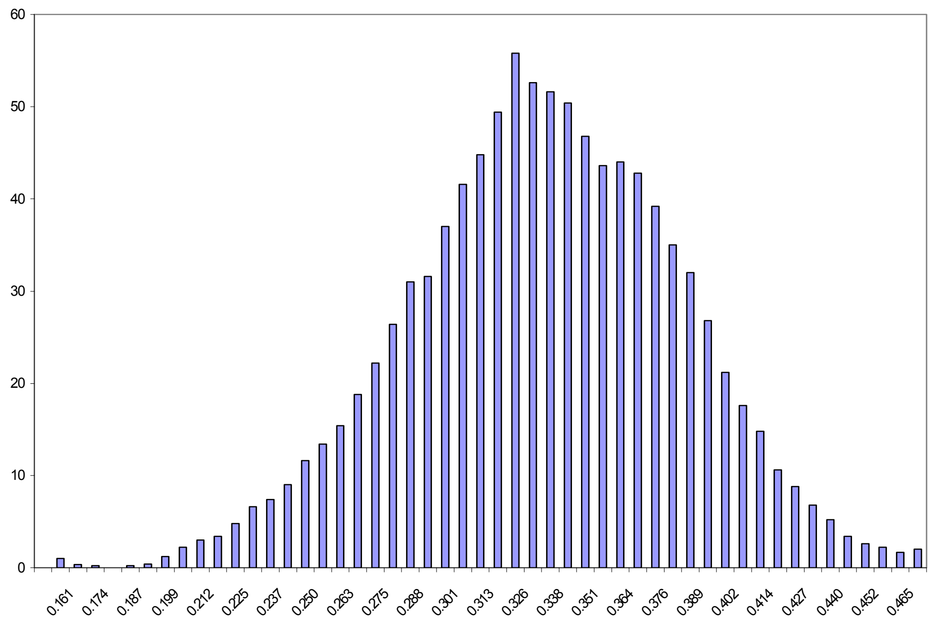
$$\begin{aligned} H_0 &: \text{ There are } I \text{ regimes in the data} \\ H_1 &: \text{ There are } I+1 \text{ regimes in the data.} \end{aligned}$$

We start with the case of $I = 1$ that is, the data are drawn from the null of a single regime defined by model (2) and from two different regimes under the alternative described by model (3). By denoting $\hat{\theta}_I$ and $\hat{\theta}_{I+1}$ as the maximum likelihood estimators (MLE) of the parameter vectors under the null and alternative hypotheses, the likelihood ratio statistic is

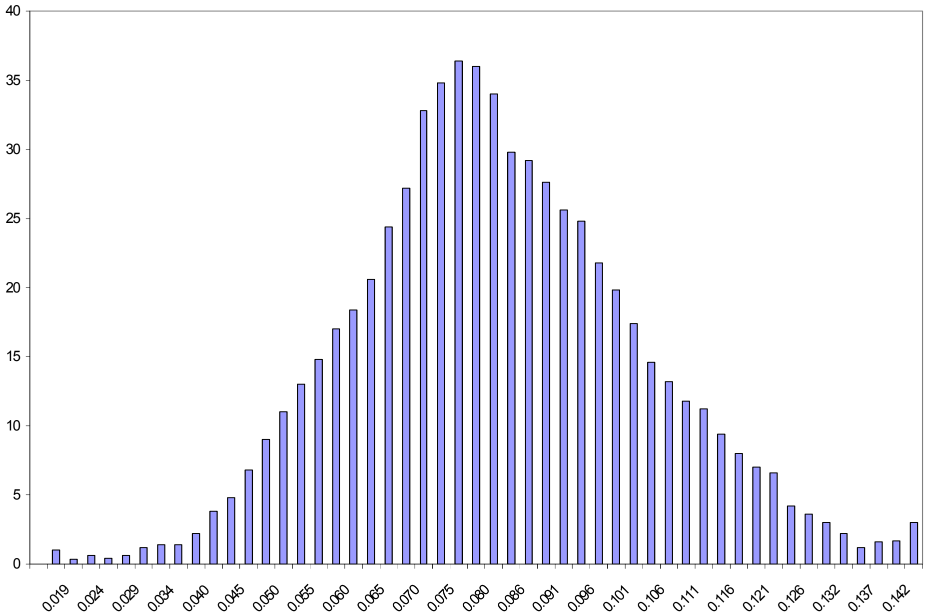
$$LR = 2[L(\hat{\theta}_{I+1}) - L(\hat{\theta}_I)] \tag{5}$$

where $L(\hat{\theta}_I)$ and $L(\hat{\theta}_{I+1})$ are the log-likelihood functions evaluated under the null and alternative hypotheses respectively. To determine the significance of the statistic computed from the actual data, we perform Monte Carlo simulations to derive the empirical distribution of the likelihood ratio statistic. The construction of the empirical distribution for the null of no markov-switching dynamics involves the following steps: (a) Use the maximum likelihood procedure to obtain $\hat{\theta}_I$; (b) generate a sample of artificial data using $\hat{\theta}_I$; (c) compute the likelihood ratio statistic (5) from the generated data; (d) repeat steps (b) and (c) M times and store the simulated likelihood ratio statistics, and (e) determine the number of simulated statistics, m, that are larger than the likelihood ratio statistic computed from the actual data series. The empirical p value of the test is computed as $(m+1)/(M+1)$. If the null of no Markov-switching dynamics fails to be rejected, we terminate the test at $I = 1$, and conclude that the short rate displays no evidence of regime dependence dynamics, otherwise we proceed to test for $I = 2$. That is, the data are drawn from two regimes under the null as opposed to three regimes under the alternative hypothesis. We set $M = 1000$ as opposed to 500 in the case of Cheung and Erlandsson (2005).

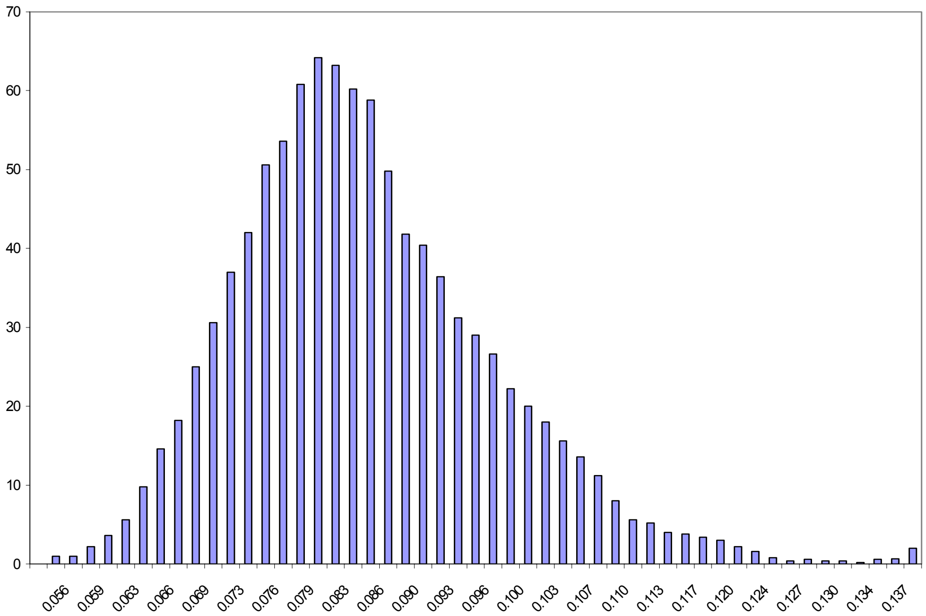
Figure 2: LR Test Statistic Empirical Distribution



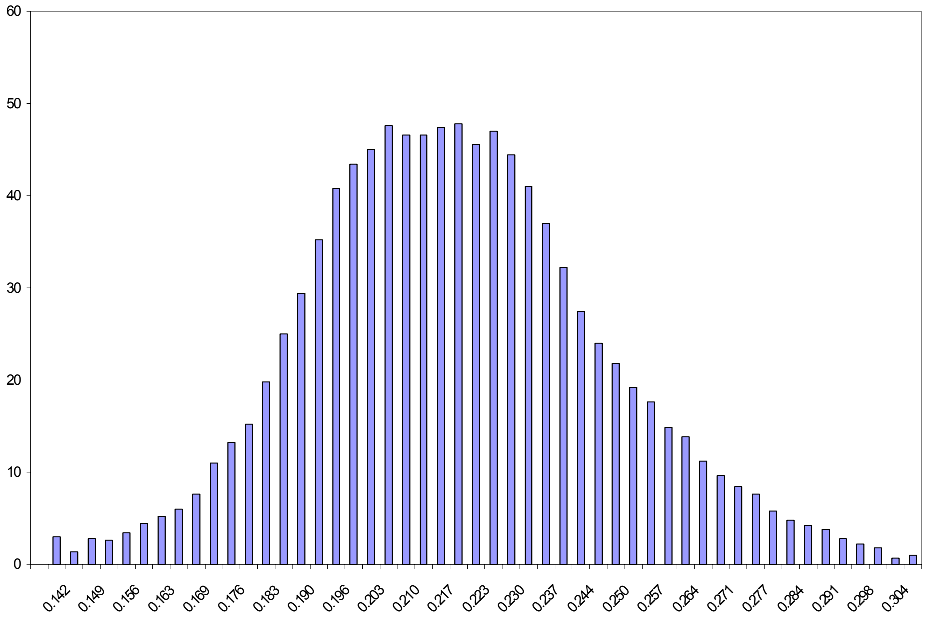
Hong Kong



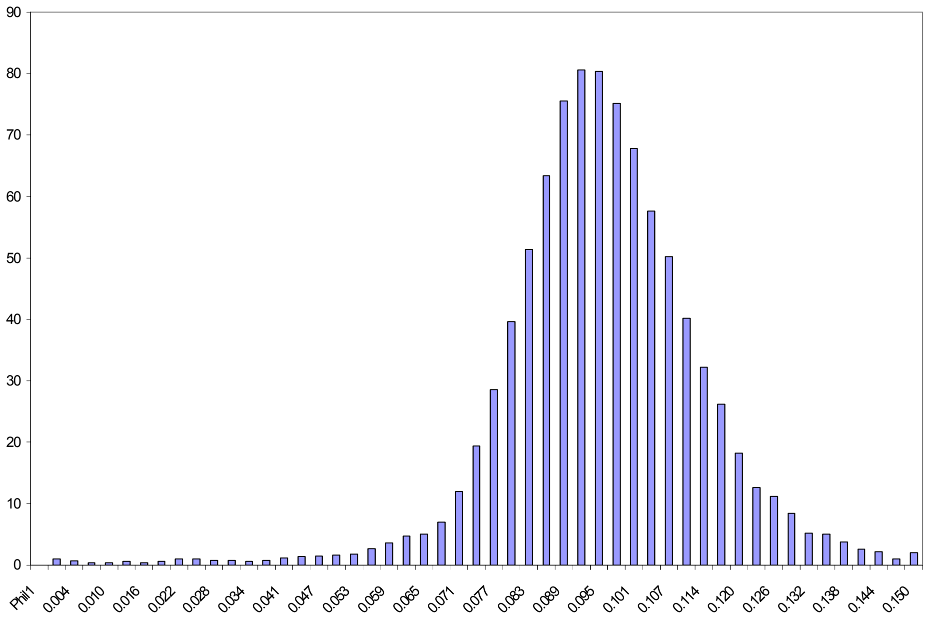
Singapore



Thailand



Indonesia



Philippines

Table 3 reports the results of the likelihood ratio test. According to the p-values the likelihood ratio test rejects the null hypothesis of single regime in favour of the two state Markov switching model for all short rates. In addition, as indicated by the p-values of the likelihood test under the null of two state Markov-switching dynamics, the test is not in favour of a Markov switching model with three states.

To conserve space, we only report the empirical distributions (see Figure 2) and the descriptive statistics of the LR test statistic under the null of linearity for all series (see Table 3). The descriptive statistics for the LR test statistics under the null of a two-state Markov-switching dynamic is available from the authors upon request. For all short rates, the empirical distributions have their means larger than the medians. The LR statistic distribution is positively skewed in the case of Singapore, Thailand and Indonesia. As for the rest, they are negatively skewed. It can be seen that the empirical distributions are series-specific.

Table 3: Testing for Markov-Switching Dynamics

	Sample LR-statistic	
	H_0: I=1 vs. H_1: I=2	H_0: I=2 vs. H_1: I=3
Hong Kong	1.5638 [0.0000]	0.3121 [0.6105]
Singapore	1.2060 [0.0000]	0.1014 [0.3150]
Thailand	1.1553 [0.0000]	0.0731 [0.6051]
Indonesia	2.5485 [0.0000]	0.2105 [0.5363]
Phillippines	3.6599 [0.0000]	0.1208 [0.2179]

	Empirical Distribution of LR statistic under $H_0 : I = 1$				
	Mean	Median	SE	Skew	Max
Hong Kong	0.3297	0.3293	0.0481	-0.1251	0.4682
Singapore	0.0810	0.0793	0.0205	0.2263	0.1430
Thailand	0.0840	0.0824	0.0121	0.7140	0.1380
Indonesia	0.2159	0.2146	0.0280	0.2398	0.3059
Phillippines	0.0953	0.0954	0.0158	-0.6321	0.1516

Note: Figures in [ ] are p-values.

4.2 Testing for a unit root in a Markov-switching Framework

Based on the evidence of a two state Markov-switching dynamics, we proceed to test for a unit root for each regime. Following the approach of Kanas and Genius (2005), we account for two distinct Markov-switching regimes by modifying the Augmented Dickey Fuller (ADF) regression as follows

$$\begin{aligned} \Delta r_t \;=\;& s_t(\mu_0 + \lambda_0 r_{t-1} + \sum_{j=1}^{p} \beta_{0j}\Delta r_{t-j}) + \\ & (1 - s_t)(\mu_1 + \lambda_1 r_{t-1} + \sum_{j=1}^{p} \beta_{1j}\Delta r_{t-j}) + \\ & [\sigma_0(1 - s_t) + \sigma_1 s_t]e_t. \end{aligned} \tag{6}$$

Here the parameters are defined in the same way as the Markov-switching interest rate model (3). Lagged of Δr are included in the unit root regression to ensure that the residual follows a white noise process. The optimal lag length is determined on the basis of Akaike Information Criterion. A lag length of one is chosen for Singapore and the Philippines short rates while zero lag length is chosen for others. Details of the estimation procedure for regression (6) using the quasi-maximum likelihood is laid out in the Appendix.

Here the unit root test for each regime corresponds to the t-tests of the null hypotheses $\lambda_0 = 0$ and $\lambda_1 = 0$ against the respective one-sided alternatives $\lambda_0 < 0$ and $\lambda_1 < 0$. However, under the null that $\lambda_i = 0$ for $i = 0, 1$ the transition probabilities, p_{ii}, are not identified. The presence of unidentified parameters invalidates the asymptotic theory for which the test statistic can be shown to follow a t-distribution. Hence, Monte Carlo simulations involving the following steps are performed to obtain the p-values of the t- statistics: (a) Estimate regression (6) under the null $\lambda_i = 0$ for $i = 0, 1$; (b) generate a sample of size equal to the data series that follows the estimated data generating process (DGP) in (a)[4]; (c) fit regression (6) to each realization of the sample and obtain the two t-statistics for

[4]For the purpose of generating the artificial data, the estimated transition probabilities in (a) are used.

the parameter λ, one for the volatile regime and the other for the tranquil regime; (d) repeat steps (b) and (c) 10,000 times and store the two series of t-statistics; and (e) compute the resulting p-values by expressing the percentage of the generated t-ratios that are below the t-values from the estimated model under the alternative hypothesis.

Table 4: Parameter Estimates of the MSADF Regression

		Hong Kong	Singapore	Thailand	Indonesia	Phillipines
Regime 1	μ_0	0.0383 (0.0206)	0.0121 (0.0120)	0.0183 (0.2503)	0.0192 (0.0333)	0.1054 (0.0295)
	λ_0	−0.0081 (0.0038)	−0.0048 (0.0042)	−0.0020 (0.0038)	−0.00001 (0.0026)	−0.0140 (0.0030)
	β_{01}	−	−0.0568 (0.0394)	−	−	0.0014 (0.0028)
	σ_0	0.2839 (0.0099)	0.1510 (0.0072)	0.2003 (0.0147)	0.2116 (0.0085)	0.1854 (0.0088)
	P_{00}	0.9789 (0.0065)	0.9802 (0.0068)	0.9331 (0.0218)	0.8965 (0.0142)	0.8874 (0.0157)
Regime 2	μ_1	2.9702 (0.6942)	0.3286 (0.1686)	1.0178 (0.0178)	1.4555 (0.5230)	10.5461 (1.4557)
	λ_1	−0.4134 (0.0846)	−0.0807 (0.0367)	−0.0993 (0.0229)	−0.0616 (0.0181)	−0.7114 (0.0944)
	β_{11}	−	−0.2781 (0.0769)	−	−	0.0134 (0.1153)
	σ_1	2.9517 (0.2305)	0.9920 (0.06859)	2.2641 (0.0864)	7.2743 (0.2799)	9.7813 (0.4281)
	P_{11}	0.8107 (0.0528)	0.9081 (0.0276)	0.9534 (0.2503)	0.8350 (0.0277)	0.7822 (0.0313)
		MSADF Test Results				
Regime 1	$H_0 : \lambda_0 = 0$	−2.1316 [0.1378]	−1.1429 [0.2043]	−0.5263 [0.3459]	−0.0038 [0.5683]	−4.6667* [0.0051]
Regime 2	$H_0 : \lambda_1 = 0$	−4.8953* [0.0030]	−2.1989 [0.1098]	−4.3362* [0.0164]	−3.4033* [0.0417]	−7.5360* [0.0015]

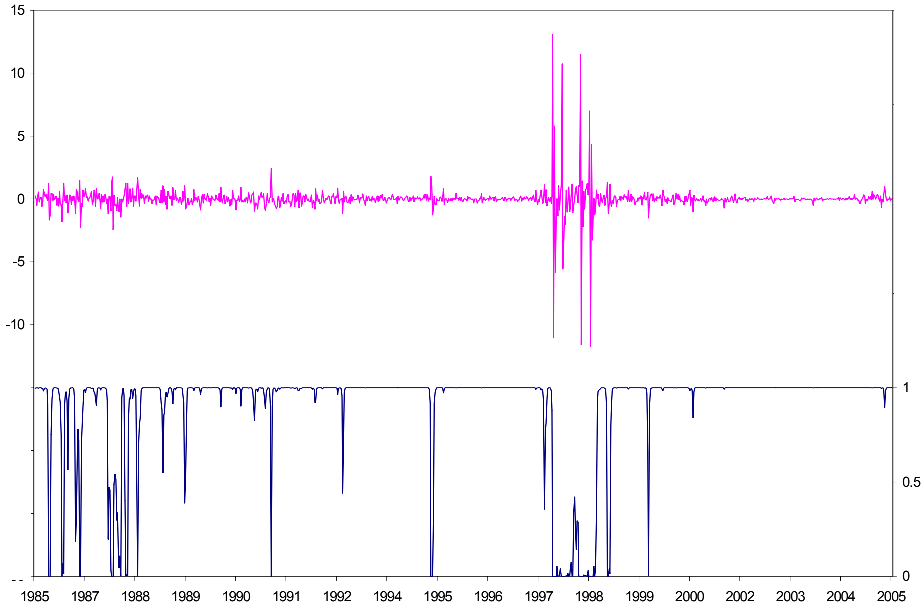
Note: Figures in ( ) and [ ] are standard errors and p-values respectively.

Table 4 shows the estimated MSADF regression results. The parameter estimates of the intercept ($\hat{\mu}$), mean-reversion coefficient ($\hat{\lambda}$) and unconditional variance ($\hat{\sigma}$) are vastly distinct in magnitude across the two regimes. In regime 2, the unconditional standard deviation $\hat{\sigma}_1$ is significantly larger than $\hat{\sigma}_o$ in regime 1 which suggests that the more volatile state is represented by regime 2. Likewise, the short rate level is higher in regime 2 than in regime 1 (i..e. $\hat{\mu}_1 > \hat{\mu}_0$). In addition, there is a large difference in the magnitude of the unconditional variance across the two regimes obtained from the MSADF regression relative to the descriptive statistics of the two subsamples. In particular, for countries like Indonesia and the Philippines that have historically high level of interest rates and large changes in the short rates, the difference in the unconditional variance of the two regimes is as large as 95.64 for the Philippines and 52.87 for Indonesia. The transition probabilities for both regimes are generally high with $p_{00}(p_{11})$ ranging from 0.9802 (0.9534) and 0.8874 (0.7822) thereby implying that both regimes are very persistent. Figure 3 presents the plot of the smoothed probabilities of the tranquil regime (regime 1) and the change in short rate in the lower and upper panels respectively. A common feature of the smoothed probabilities plots is that the impact of the 1997 crisis is clearly visible. The ergodic probability of short rates being in tranquil regime is almost always zero suggesting that short rates have a tendency to stay at high levels and are volatile. The regime 1 smoothed probabilities for Thailand and the Philippines also indicate that their short rates have remained rather volatile even prior to the crisis. All short rates exhibit persistence in the tranquil regime after 1998.

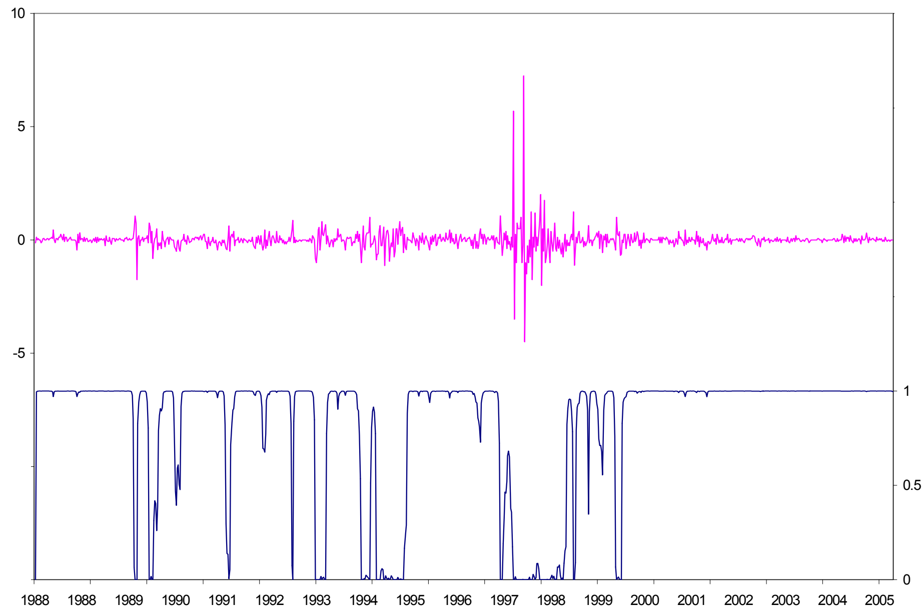
Based on the simulated p-value, the test statistic for the mean-reversion coefficient $\hat{\lambda}_1$ in tranquil regime fails to reject the null of nonstationarity for all short rates other than the Philippines'. As for $\hat{\lambda}_2$ the p-value of its test statistic rejects the null of nonstationarity in all cases except for Singapore. In summary, the MSADF test results show that short rates in Hong Kong, Thailand and Indonesia behave like a random walk when they are in tranquil regime but mean-revert when the volatility and short rate levels are high. In contrast, the Philippines short rate mean-reverts in both tranquil and volatile regimes, although the speed of mean reversion differs in the two regimes.[5] Noticeably, the Philippines short rate levels and volatility are high in both regimes implying that short rate tends to adjust to its regime long run mean. Singapore short rate stands out from the rest in that it follows a unit root process in both low and high volatility regimes. The nonstationarity of Singapore short rate can be explained by the low short rate levels and volatility in both regimes. Overall, the results predict that there is strong mean-reversion only during periods of high interest rates and high volatility. During periods of low interest rates and low volatility, interest rates seem to behave like an I(1) process.

[5] We test the null hypothesis that the mean-reverting coefficient of the Philippines short rate is the same in both regimes (i.e. $\lambda_0 = \lambda_1$). This is done by re-estimating regression (5) under the null. The LR test statistic that is 4.15 is distributed as $\chi^2(1)$ and it rejects the null hypothesis at the 5% significance level.

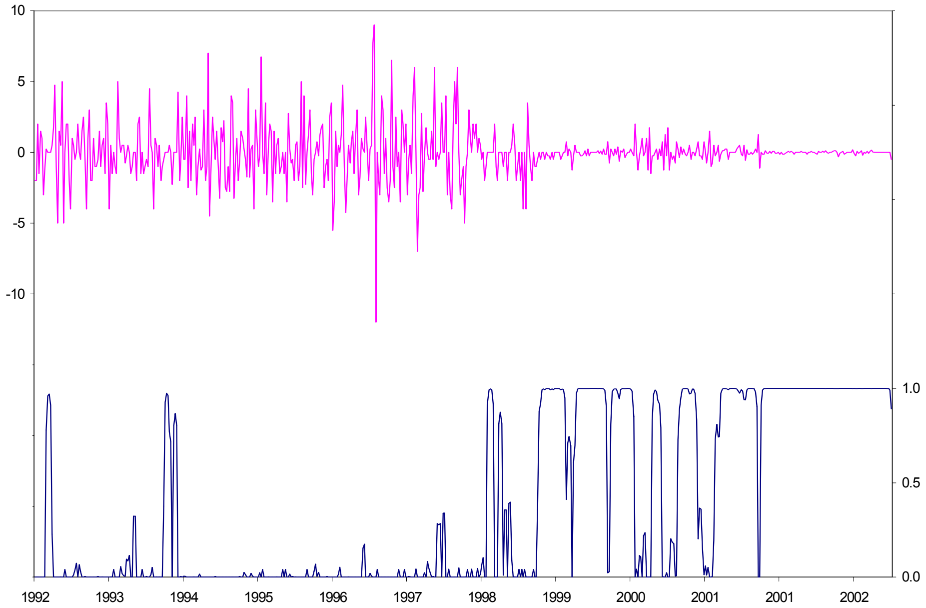
Figure 3. Smoothed Probabilities in Regime 1 (Lower Panel) and Change in Short Rate (Upper Panel)



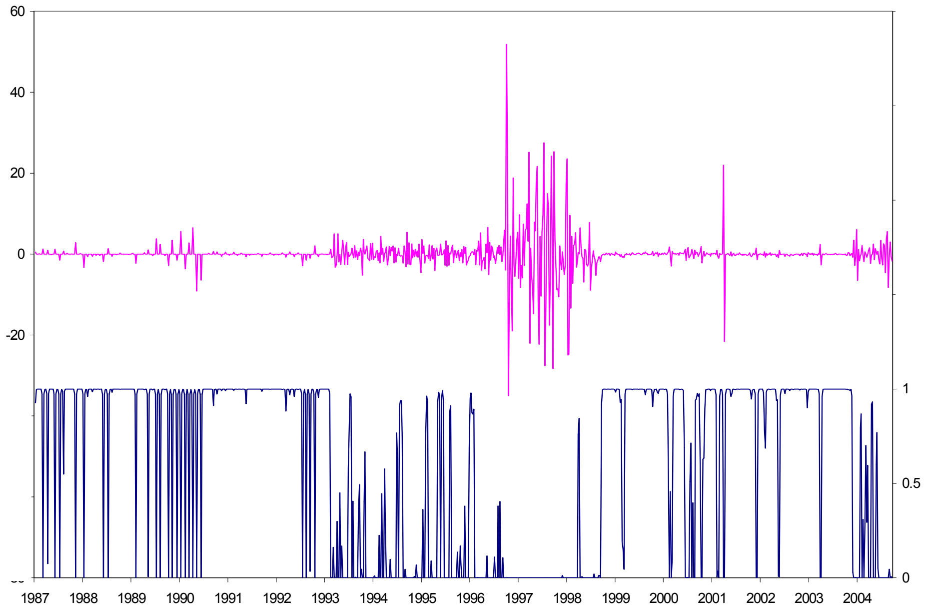
Hong Kong



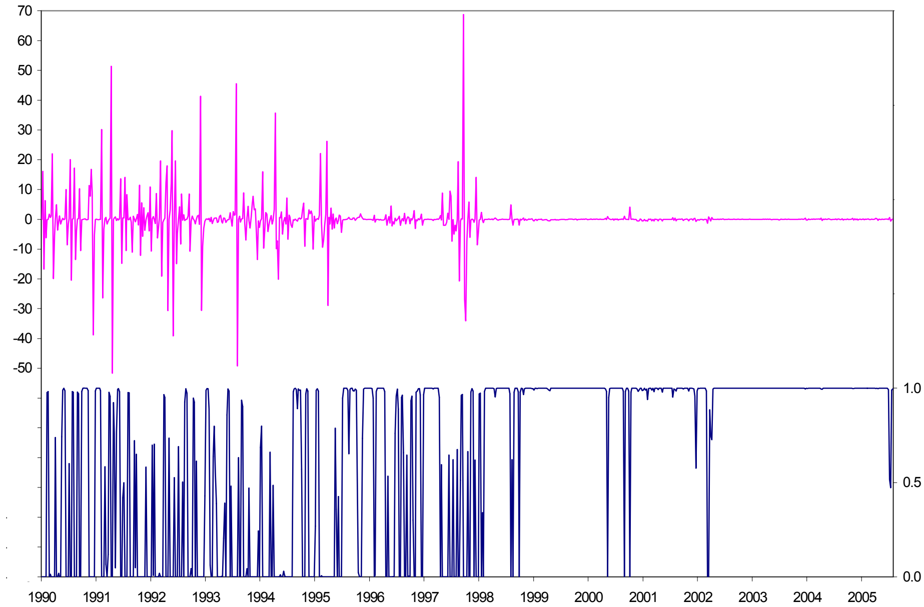
Singapore



Thailand



Indonesia



Philippines

Comparing the MSADF with the standard ADF tests results, it can be seen that the two tests arrive at different conclusions regarding the mean-reversion property of Asian short rates (with the exception of Singapore and the Philippines). For example, the ADF test detects stationarity for the whole sample of Indonesia short rate even though the MSADF test shows that it is stationary only when it is in the volatile regime. Nelson et al. (2001) and Kanas and Genius (2005) document the sensitivity of the ADF test power to the presence of occasional unit root process. To show that the MSADF can perform better in detecting the occasional unit root process from an integrated process of order one, we conduct a Monte Carlo experiment to compare the power of these two tests based on our sample. Given the occasional integratedness in the short rates for Hong Kong, Thailand and Indonesia, we generate data using the estimates in Table 4 and impose $\lambda_o = 0$ in regime 1. The sample size of the generated series is matched to the country-specific short rate sample. The ADF and MSADF tests are then applied to the generated short rate series and the percentage of replications for which each of the tests rejects the unit root hypothesis is computed. The experiment is repeated for 10,000 times. Table 5 reports the Monte Carlo experiment results.

Table 5: Power Comparison of ADF and MSADF Tests

	ADF Test	MSADF Test	
		Regime 1	Regime 2
Hong kong	55.9	-	93.1
Thailand	83.1	-	100
Indonesia	77.2	-	93.5
Phillippines	100	100	99.06

For the three scenarios, the ADF test distinguishes the occasional integrated process from an I(1) process at a rate that is lower than the MSADF test. The power of the MSADF test is impressive exceeding 90% at all times. The power of the ADF test, on the contrary, is much lower. In the case of Hong Kong, the power of the ADF test (which is 56%) is the weakest, this is because the coefficient of mean-reversion for the stationary regime is furthest from zero. That is, the low power could be attributed to the ADF test inability to differentiate the occasional integrated process from an I(1) process when the stationary regime has a high speed of adjustment to its long-run mean. To confirm our intuition, we simulate the data for Hong Kong using $\lambda_1 = -0.05$ while keeping the other parameter estimates unchanged. The result shows a significant improvement in the power of the ADF test to 81%. We also perform the Monte Carlo experiment using the Philippines short rate which shows evidence of stationarity in both regimes but with different speed of mean-reversion. Interestingly, we find that not only is the ADF test power robust in such a case, the MSADF test also exhibits as good a power.

5 Sources of (Non)Stationarity in Asian Short Rates

The finding that short-term interest rate models yield one (near) unit-root and one more mean-reverting regime is commonly reported in the empirical literature using U.S. short rate data (see Gray (1996) Ang and Bekaert (1998), Holst, Lindgren, Holst and Thuvesholmen (1994), and *inter alia*). Our results for Hong Kong, Thailand and Indonesia concur with the empirical literature. Nevertheless, the results for the Philippines and Singapore differ widely from the established U.S. short rate dynamics. We argue that these short rate dynamics are closely related to the Asian economies underlying macroeconomic fundamentals, exchange rate regime and their monetary policy objectives.

In the case of Singapore, the result is reflective of the economy exchange rate centred monetary policy. Due to Singapore's high reliance on imports, the domestic inflation is largely determined by changes in foreign prices and the exchange rate. Traditional monetary policy instruments, such as money supply and interest rates, are therefore less effective in controlling prices. The Monetary Authority of Singapore maintains a strong and credible currency and explicitly discourages the internationalisation of the Singapore dollar (SGD), that is the use of the SGD outside Singapore for activities unrelated to its real economy. These policies which contribute to the strength and stability of the SGD have instilled confidence and kept inflation low relative to her neighbouring countries. Consequently, Singapore short rate has remained low and follows a random walk process. While the Asian financial crisis may have impacted the economy's short rate, the effect on the short rate is mild and short lived (see the mild rise in short rate level in July 1997 relative to the other economies in Figure 1). Singapore's sound economic fundamentals and credible exchange rate policy have also rendered speculation on the SGD unattractive. It is therefore not surprising to find that the short rate fails to mean-revert even in the relatively volatile regime.[6]

Contrary to Singapore, between the early 1990s and 1998 the Philippines economy has witnessed volatile and high levels of short rate. The initial high levels of short rate is due to the central bank lack of commitment in targeting the monetary aggregate. As shown by Gochoco (1991), the central bank was more commited to targeting the exchange rate rather than the monetary aggregate. Apart from the large and discrete devaluations of the peso in 1983 and 1984, Gochoco (1991) argued that the exchange rate remained very stable while monetary growth and interest rates displayed large variability. The erratic nature of the monetary policy also resulted in wild swings in the inflation rates. In 1992, the foreign exchange market was liberalised and the capital account was virtually opened, however,

[6] Note that Singapore's level of short rate in the volatile regime is significantly lower than the other short rate estimates.

these events were not backed by adequate strengthening of bank supervision and regulation. There was also large increase in foreign borrowing caused by high domestic market interest rates relative to those in both developed countries and other ASEAN countries (with the exception of Indonesia). The large capital inflows interferes with the conduct of monetary policy as the simultaneous targeting of the exchange rate and the money supply is not feasible when capital is mobile. Sterilization was therefore adopted to gain some degree of control over monetary policy, even though such a policy could not be undertaken indefinitely. Consequently, the peso became prone to speculation. In 1997, there was a large depreciation of the peso and a tightening of liquidity ensued which caused the Philippines short rate to shoot up to an unprecedented level of 80%. One reason why interest rates mean revert when it reaches new heights is that "market participants may expect the Federal Reserve to credibly return the short-term interest rate to its middle range at some point, but are uncertain about the precise timing of the intervention" Ait-Sahalia (1996, pp. 407). The high levels of short rates relative to the others in tranquil and volatile regimes ($\hat{\mu}_0 = 0.1054$ and $\hat{\mu}_1 = 10.5461$) are also indicative of the mean-reverting behaviour in the Philippines short rate in both states.

The exchange rate regime of the Hong Kong, Indonesia and Thailand economies potentially explain the similarity in the dynamics observed in these economies short rates to that of the U.S.. Thailand adopted a pegged exchange rate regime between the second world war until June 1997. The low GDP growth rate and the oil crises of the 1970s led to the devaluation of the Thai Baht and the pegging of the Baht to the US$ and later to a basket of currencies. Nidhiprabha (1993) noted that Thailand monetary policy was designed to be in line with the pegged exchange rate regime and had little to do with price stabilisation. The case of Indonesia is similar. The Indonesian Rupiah was also pegged to the US$ prior to the crisis and had undergone numerous major devaluations. Bank Indonesia's role of managing the exchange rate between the Rupiah and foreign currencies sometimes conflicted with the objective of controlling the amount of bank credit which indirectly influences the level of interest rates. Finally, monetary policy in Hong Kong is tied to maintaining the nominal exchange rate linked to the U.S. dollar. While the Hong Kong Monetary Authority respond to market pressures by occasionally adjusting liquidity through interest rate changes and intervention in the foreign exchange and money markets, the adjustments to its interest rate cannot cause it to deviate significantly from the U.S. rate.

6 Conclusion

In this paper, we examine the stationarity properties of five East Asian economies short-term interest rates within a Markov-switching framework. The likelihood ratio test statistic developed by Cheung and Erlandsson (2005) is employed to detect the presence of Markov-switching dynamics in these short rates. In addition, a regime-switching ADF test of Kanas and Genius (2005) is used to capture the periodic changes in the stationarity of the short rate introduced by these regime changes.

The empirical results suggest that nonlinearity in the form of Markov-switching dynamics is present in all Asian short rates. The regime-switching ADF test shows that short rates tend to revert to some long-run mean when the interest rate levels are high and volatile. However, during periods of low interest rates and low volatility, interest rate behaves more like a random walk process. These findings are consistent with the empirical literature that is largely based on the U.S. short rates. Hong Kong, Indonesia and Thailand short rates mean-revert in periods when their interest rate levels are high and volatile, but fail to do so when they are in tranquil regime. Their short rate behaviour mimics that of the U.S. short rate because their monetary policy, to a large extent, is tied to maintaining the nominal exchange rate to the U.S. dollar. Singapore short rate, on the other hand, exhibits nonstationarity in both regimes. These results are attributed to her credible exchange rate centred monetary policy and sound macroeconomic fundamentals that have kept interest rates at low levels with little volatility, even during the 1997 Asian financial crisis. On the contrary, the Philippines economy has suffered from long periods of high and volatile interest rates which are the results of erratic monetary policy that conflicts with the objective of targetting the exchange rate. The high inflation rates also contributed to the unusually high and volatile short rates. As a result, the Philippines short rates tend to mean-revert to its long-run mean in both regimes.

Finally, we show that employing a Markov-switching unit root test on the short-term interest rate uncovers the occasional stationarity properties of the process. This would not have been possible when using the standard ADF test because of its low power in the presence of occasional unit roots. These findings bear important implications for both interest-rate modeling and forecasting. For modeling and forecasting purposes, a regime-switching framework incorporating different mean-reversion speeds in each regime provides a superior characterisation of the short rate and is more likely to improve the predictive power of the model. An extension of this research agenda is to utilise the MSADF test considered in this paper to examine the mean-reverting properties of interest rates across various maturities and frequencies of actual data.

Appendix

The MSADF regression (6) has the vector of parameters $\theta = (\lambda_0, \lambda_1, \beta_{01}, ..., \beta_{0p}, \beta_{11}, ..., \beta_{1p}, \sigma_0, \sigma_1, p_{00}, p_{11})'$. Let $r_T = \{r_t, r_{t-1}, ...r_1\}$ denote the sample of all the observed short rates up to time t. Under normality assumption, the density of r_t conditional on r_{t-1} and $s_t = i$ (for $i = 0, 1$) is

$$f(r_t|s_t = i, r_{t-1}; \theta) = \frac{1}{\sqrt{2\pi[\sigma_0(1-i) + \sigma_1 i]^2}} \times$$

$$\exp\left(\frac{-\left(\Delta r_t - \lambda_0(1-i)r_{t-1} - \lambda_1(i)r_{t-1} - \sum_{j=1}^{p}[\beta_{0j}(1-i) + \beta_{1j}i] \cdot \Delta r_{t-j} \right)^2}{2[\sigma_0(1-i) + \sigma_1 i]^2} \right). \tag{A1}$$

Given the prediction probability $P(s_t = i|r_{t-1}; \theta)$, the density of r_t conditional on r_{t-1} can be obtained from equation (A1) as

$$\begin{aligned} f(r_t|r_{t-1}; \theta) &= P(s_t = 0|r_{t-1}; \theta) f(r_t|s_t = 0, r_{t-1}; \theta) \\ + P(s_t &= 1|r_{t-1}; \theta) f(r_t|s_t = 1, r_{t-1}; \theta). \end{aligned} \tag{A2}$$

The filtered probabilties of s_t for $i = 0, 1$ are

$$P(s_t = i|r_t;\theta) = \frac{P(s_t = i|r_{t-1};\theta)f(r_t|s_t = i, r_{t-1};\theta)}{f(r_t|r_{t-1};\theta)}. \tag{A3}$$

By invoking Bayes theorem, the relationship between the filtered and prediction probabilities is

$$P(s_t = i|r_{t-1};\theta) = p_{0i} \cdot P(s_{t-1} = 0|r_{t-1};\theta) + p_{1i} \cdot P(s_{t-1} = 1|r_{t-1};\theta), \tag{A4}$$

where $p_{0i} = P(s_t = i|s_{t-1} = 0)$ and $p_{1i} = P(s_t = i|s_{t-1} = 1)$ are the transition probabilities. Note that equations (A1)-(A4) form a recursive system for $t = 1, ..., T$.

With the initial values $P(s_m = i|r_{m-1};\theta)$, we can iterate equations (A1)-(A4) to obtain the filtered probabilities $P(s_t = i|r_t;\theta)$ and the conditional densities $f(r_t|r_{t-1};\theta)$ for $t = m, ...T$.[7] This gives the quasi-log-likelihood function

$$L(\theta) = \frac{1}{T}\sum_{t=1}^{T} \ln f(r_t|r_{t-1};\theta). \tag{A5}$$

The vector of parameter estimates $\hat{\theta}$ is obtained by maximising the log-likelihood function (A5) using the BFGS algorithm.

The estimated filtered and prediction probabilities can be obtained by substituting $\hat{\theta}$ into equations (A3) and (A4) respectively. The smoothed probabilities $P(s_t = i|r_T;\theta)$ are obtained by substituting $\hat{\theta}$ into

$$P(s_t = i|r_T;\theta) = P(s_t = i|r_t;\theta)\left(\frac{p_{i0}P(s_{t+1} = 0|r_T;\theta)}{P(s_{t+1} = 0|r_t;\theta)} + \frac{p_{i1}P(s_{t+1} = 1|r_T;\theta)}{P(s_{t+1} = 1|r_t;\theta)}\right). \tag{A6}$$

See Kim (1994) for a discussion on the derivation of the smoothed probabilities.

[7] Hamilton (1994) suggests setting the initial value $P(s_0 = i|r_0;\theta)$ in equation $(A4)$ to its limiting unconditional counterpart $\frac{1-p_{11}}{2-p_{00}-p_{11}}$ and $\frac{1-p_{00}}{2-p_{11}-p_{00}}$ for $i = 0$ and 1 respectively.

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