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Are Inflation Rates Mean-reverting Processes? Evidence from Six Asian Countries

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This paper tests for the (non)stationarity of inflation in six Asian countries, that is, Indonesia, Malaysia, the Philippines, Thailand, Singapore and India, by considering the possibility of structural breaks and nonlinearity. The results are mixed when a battery of conventional linear unit root tests is used. However, the inflation rate is shown to be a stationary process after considering the structural breaks and nonlinear properties of the threshold and exponential smooth transition. The mean reversion in inflation favors the hypothesis of the natural rate of inflation and the sticky-price model, indicating that any shock has a transitory effect on inflation. The implication of stationary inflation is that any shock has a transitory effect on inflation.

Keywords: inflation, stationarity, structural break, nonlinearity

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1 Introduction

Being aware of the persistence in inflation has long been the focus of attention of academia, government and the populace alike because it is helpful not only in distinguishing between different economic hypotheses, but is also important to the monetary authorities in implementing the policies of disinflation.[1] There are two kinds of theoretical discussion on inflation persistence in the literature. First, the accelerationist hypothesis and the Fisher hypothesis suggest that inflation is a nonstationary process. The accelerationist view implies an ever-increasing level of inflation in order to keep unemployment below its natural rate. Therefore, in the case where the unemployment rate is kept below the natural rate, the inflation would be characterized as a unit root (Romero-Avila and Usabiaga, 2009). Moreover, the Fisher hypothesis claims that if the nominal interest rate contains a unit root, in order for the real interest rate to be stationary, it is necessary for the inflation rate to follow a unit root process and to be cointegrated with the nominal interest rate.

On the contrary, a number of theoretical models, for example, the rational expectations version of Cagan (1956), the sticky-price model (Dornbusch, 1976; Taylor, 1979) and the 'higher-order' Phillips curve model (Calvo, 1983; Ball, 1993), assert that the inflation rate is a mean-reverting process. Traditionally, the New Keynesian Phillips curve has been estimated under the assumption that inflation is a stationary process (Gali and Gertler, 1999; Gali *et al*., 2001; Nymoen *et al*., 2012). Moreover, the hypothesis of a "natural rate of inflation" assumes that inflation is a stationary process. As outlined in Russell (2011), inflation could also be a stationary process around shifting means. The rational expectations hypothesis proposes that the stable growth of money supply implies stationary inflation (Yellen and Akerlof, 2006). Arize *et al*. (2005) and Arize and Malindretos (2012) also emphasize that the stationarity of the inflation rate is an important consideration in the estimation of money demand relations. Cecchetti and Debelle (2006) stress that stationary inflation will incur a lower cost for the monetary authorities in conducting monetary policy.

[1]Culver and Papell (1997) point out that the construction and evaluation of monetary policy rules is affected by the (non)stationarity of inflation.

A wealth of research reflects many efforts devoted to the issue of inflation persistence since the studies conducted by Nelson and Plosser (1982) and Rose (1988). Interested readers are referred to, for example, to name a few studies, Choi (1994), Baillie *et al*. (1996), Crowder and Hoffman (1996), Culver and Papell (1997), Lee and Wu (2001), Henry and Shields (2004), Rapach and Weber (2004), Romero-Avila and Usabiaga (2009), Beechey and Osterholm (2009), Cuestas and Harrison (2010), Mourelle *et al*. (2011), Tsong and Lee (2011), Narayan and Popp (2011), Arize *et al*. (2005), Arize and Malindretos (2012), Chang *et al*. (2013), Zhang (2013) and Martins and Rodrigues (2014) for the details.

Several features characterize previous studies in the literature. First, an important feature of previous studies is that distinct results based on previous research are due to differences in methodology, approaches and samples and are subject to diverse interpretations, thus making it difficult to reach a corroborative position on the stationarity property of inflation. Second, early research employs the traditional method in testing the null hypothesis of a unit root in inflation. It is well known that the power of the traditional unit root test is significantly reduced if the true data-generating process of a series exhibits structural breaks (Perron, 1989). Therefore, recent studies such as Culver and Papell (1997), Lee and Wu (2001), Costantini and Lupi (2007), and Basher and Westerlund (2008) have started to adopt newly-developed (panel) unit root tests with or without a break (Im *et al*., 2005; Westerlund, 2005) to investigate the stationarity property of inflation rates. Third, a number of studies (Cuestas and Harrison, 2010; Omay and Hasanov, 2010; Mourelle *et al*., 2011; Chang *et al*., 2013), however, find that inflation tends to be better characterized by a nonlinear data-generating process than a linear model.

Although a large number of studies have been based on the persistence of inflation for the US and European countries (for example, Lee and Wu, 2001; Gregoriou and Kontonikas, 2006; Baillie and Morana, 2012; Giannellis, 2013; Chen, 2015),[2] little is known about inflation persistence for Asian countries (Gerlach and Tillmann, 2012). This paper attempts to fill this gap. In this study we examine the inflation persistence for six Asian countries, including India and the five largest economies in the Association of Southeast Asian Nations (ASEAN5), i.e., Indonesia,

[2]Readers are referred to the appendix of Murray *et al*. (2009) for a summary of studies on US inflation persistence.

Malaysia, the Philippines, Singapore and Thailand. Studies on these Asian countries are attractive because the rapid economic growth and development in these countries enable them to play more important roles in the world economy. Therefore, the empirical results from this study are useful not only in classifying which theoretical hypothesis is suitable for explaining the time-series properties of the inflation series in these six Asian countries, but also allow us to understand the attitude of the monetary authorities in conducting monetary policy to deal with the inflation in these countries.

In order to obtain robust inferences on inflation persistence, it is not necessary to assume a *one-size-fits-all* approach. In particular, Culver and Papell (1997), Malliaropulos (2000) and Costantini and Lupi (2007) have shed light on the importance of recognizing the possibility of a structural shift in testing for the null hypothesis of a unit root in inflation rates. Neglecting structural changes could lead to the erroneous acceptance of nonstationarity. If the true adjustment process is asymmetric, then the restrictive symmetric adjustment in standard unit root tests that is implicitly assumed is indicative of model misspecification (Pippenger and Goering, 1993). In order to take account of the possibility of the structural break and asymmetric adjustment in the inflation rate, we consider the following unit root tests in this study.

First, it is well-known that in the presence of a structural break, the power to reject a unit root decreases if the stationary alternative is true and a structural break is ignored (Perron, 1989). To address this problem, in addition to the Zivot and Andrews (1992) (hereafter ZA) sequential one trend break model and the Lumsdaine and Papell (1997) (hereafter LP) two trend breaks model, we adopt two newly-developed structural break unit root tests proposed by Popp (2008) and Narayan and Popp (2010) in this study.[3]

Second, to address the influence of nonlinearity on the unit root test, we employ two different types of nonlinear unit root test in this study. (i) In line with the literature (e.g., Giannellis, 2013), we adopt the threshold autoregressive (hereafter TAR) and the momentum threshold autoregressive (hereafter MTAR) unit

[3]We discuss the advantages of the Popp (2008) and Narayan and Popp (2010) structural break unit root tests, compared to the Zivot and Andrews (1992) and Lumsdaine and Papell (1997) tests, in the empirical results.

root tests, proposed by Enders and Granger (1998).[4] (ii) Based upon the contribution of Sargent and Wallace (1973), inflation may behave as a nonlinear process with multiple equilibria. This type of nonlinearity is well-characterized by the exponential smooth transition autoregressive (ESTAR) model. Therefore, we apply three different ESTAR-type nonlinear unit root tests, proposed by Kapetanios *et al*. (2003), Rothe and Sibbertsen (2006) and Kruse (2011), in this study.

As compared to the literature, the contributions of this paper are twofold. First, by considering different nonlinearities and adopting the corresponding nonlinear unit root tests in this study, we are allowed to know that either one or all of them are important in determining the dynamics of inflation. Second, rather than attempting to model any structural change in trend as an instantaneous trend break, we permit the possibility of smooth transitions between two different trend paths over time. In the context of economic time series this has considerable intuitive appeal.

The major findings of this study are briefly foreshadowed as follows. First, by using standard unit root tests, we cannot reach a unanimous conclusion on the stationarity of inflation rates. Second, the results from the Popp (2008) and Narayan and Popp (2010) tests show that inflation rates are stationary processes after taking account of structural breaks. Third, the results from the TAR, MTAR and ESTAR-type nonlinear unit root tests provide strong evidence that the inflation rates are mean-reverting processes. This fact implies that nonlinearity (threshold or exponential smooth transition) is another important property in testing for inflation persistence because it is inclined to accept the null hypothesis of a unit root if nonlinearity is ignored. Fourth, the results of the Fourier unit root test, which accounts for smooth breaks, are consistent with the findings of the TAR, MTAR and ESTAR-type nonlinear unit root tests. Overall, the mean reversion in inflation favors the hypotheses of, for example, the natural rate of inflation and the sticky-price model, which implies that any shock has a transitory effect on inflation.

The remainder of this paper is organized as follows. Section 2 reviews some of

[4]Mourelle *et al*. (2011, p. 69) indicate that “the further the inflation rate deviates from the equilibrium or inflation target, the higher the efforts of the government will be to control it and, therefore, the speed of reversion of the variable (Gregoriou and Kontonikas, 2006). This implies the existence of a threshold for the inflation rate within which the monetary authority may not apply any particular policy, not only when inflation targets are set in terms of a threshold of values, but also when the costs of applying monetary policy offset the benefits of its application (see Orphanides and Wieland’s (2000) model).”

the recent empirical studies on the stationarity of the inflation rate. Section 3 outlines the statistical methods used for testing for nonlinearity and unit roots. Section 4 discusses the data used and the empirical results and compares our results with those of the extant literature. Finally, Section 5 concludes.

2 Review of the Recent Literature

Motivated by the statistical power of the advances in panel unit root tests, an increasing number of authors have applied these new tools to test for inflation persistence since 2000. For example, the results of Costantini and Lupi (2007) favor the ‘I(0) plus break’ hypothesis for a panel of 19 OECD countries. Ho (2009) applies Chang’s (2002) nonlinear panel unit root test to the inflation rates of 19 OECD countries and finds that they are nonstationary processes, concluding that inflation rates may accelerate after all. On the contrary, the findings of Romero-Avila and Usabiaga (2009) strongly reject the presence of a unit root in inflation for 13 OECD countries over the period 1957–2005, thus lending firm support to the existence of regime-wise stationarity. Basher and Westerlund (2008) test the robustness of Culver and Papell’s (1997) results and find that inflation follows an I(0) series by using a group of panel unit root tests that allow for general forms of cross-sectional dependence and multiple structural breaks in each cross-section.

Alternatively, Pippenger and Goering (1993), Gonzalez and Gonzalo (1997), and Enders and Granger (1998) demonstrate the poor performance of the ADF statistic in distinguishing between a unit root and a stationary root with a strong threshold effect. Therefore many studies employ recent newly-developed nonlinear unit root tests to examine the nonstationarity of the inflation rates for different countries or areas. For example, Henry and Shields (2004) adopt Caner and Hansen’s (2001) approach and find that inflation in the UK and Japan is well described as a two-regime threshold unit root process. Shocks to inflation are highly persistent in one regime, but have finite lives in the other regime. Giannellis (2013) tests the existence of persistent inflation rate differentials in the euro area by adopting Caner and Hansen’s (2001) threshold unit root test. The results argue that nonstationarity has been observed in the cases of Austria, Finland, France, Germany, Malta, the Netherlands, Ireland, Italy, Portugal, and Slovenia. Cuestas and Harrison (2010) test for the order of integration of monthly inflation rates for twelve Central and Eastern European countries by applying linear (panel) and nonlinear unit root tests. They find that the inflation rate turns out to be stationary for a total of 7 out of 12 countries after taking into account the possibility of nonlinearities in the inflation dynamics. Using data for 11 OECD countries, Chen (2010) employs a Markov switching unit root regression to investigate the issue of the nonstationarity and nonlinearity in inflation. The results convincingly support the view that the inflation rates in the OECD countries are characterized by a two-regime Markov switching unit root process.

Mourelle *et al*. (2011) test for the inflation persistence hypothesis and model the long-run behavior of the inflation rates in a pool of African countries using nonlinear unit root tests and fractional integration. Their results show that the hypothesis of inflation persistence does not hold empirically for most of the countries. Arize and Malindretos (2012) employ the ESTAR-type unit toot tests based on symmetric adjustment (Kapetanios *et al*., 2003) and asymmetric adjustment (Sollis, 2009 and Pascalau, 2007) for thirty-four African countries. Their results provide robust evidence in favor of the nonstationarity of CPI inflation for 34 countries.

Zhou (2013) examines the stationarity of the inflation rates for 12 European countries by using the Kapetanios *et al*. (2003) approach and finds that the majority of the inflation rates of these countries are stationary processes during the floating exchange rate periods. Chang *et al*. (2013) apply a flexible Fourier stationarity test, proposed by Becker, Enders and Lee (2006) to investigate the mean reversion of inflation in 22 OECD countries over the period from 1961 to 2011. Their empirical results show that the inflation exhibits a mean-reversion process in all 22 OECD countries based on the flexible Fourier stationarity test.

Tsong and Lee (2011) apply the regression quantile approach developed by Koenker and Xiao (2004) to investigate the dynamic behavior of inflation in 12 OECD countries. They find that, in general, the inflation rates are not only mean-reverting but also exhibit asymmetries in their dynamic adjustments, in which large negative shocks tend to induce strong mean reversion, while on the contrary, large positive shocks do not. Cicek and Akar (2013) also use Koenker and Xiao’s

(2004) method to examine the dynamic behavior of the overall inflation rate, its subgroups, and the inflation rates of traded and non-traded goods in Turkey from 1994 to 2012.

Some authors employ the fractionally integrated mode to examine inflation persistence. For example, Kumar and Okimoto (2006) investigate the dynamics of inflation persistence using fractionally integrated processes and find that there has been a clear decline in inflation persistence in the United States over the past two decades. Baillie and Morana (2012) propose an Adaptive ARFIMA model, which uses a flexible Fourier form to allow for a time-varying intercept, to examine the G7 inflation persistence. The results indicate the presence of both long memory and structural change not only in the conditional mean dynamics, but also in the conditional variance dynamics of G7 inflation rates. Martins and Rodrigues (2014) propose a new approach to detect persistence change in fractionally integrated models based on recursive forward and backward estimation of regression-based Lagrange Multiplier tests and apply it to several world inflation rates. They find that there is persistence change in most of the series.

Beechey and Osterholm (2009) investigate how inflation persistence in the Euro area has evolved between 1991 and 2006. Employing an ARMA(1,11) model with a time-varying autoregressive parameter, they find that inflation persistence has fallen markedly since the third stage of the EMU began in January 1999 and inflation no longer exhibits unit-root behavior. Narayan and Popp (2011) apply the modified seasonal unit root test with seasonal level shifts at unknown times proposed by Popp (2007) to the G7 inflation rates. They find that there is a non-seasonal unit root in Canada’s inflation rate, a semi-annual unit root in Germany’s inflation rate, and no seasonal unit root at the annual frequency for any of the G7 countries.

3 Methodology

3.1 Narayan and Popp’s (2010) Test for a Unit Root

Narayan and Popp (2010) consider an unobserved components model to represent the data generating process (DGP). The DGP of a time series y_t has two components, a deterministic component (d_t) and a stochastic component (ε_t), as follows:

$$y_t = d_t + u_t \,, \tag{1}$$
$$u_t = \rho u_{t-1} + \varepsilon_t \,, \tag{2}$$
$$\varepsilon_t = \Psi^*(L)e_t = A^*(L)^{-1}B(L)e_t \,, \tag{3}$$

with $e_t \sim iid(0,\sigma_e^2)$. It is assumed that the roots of the lag polynomials $A^*(L)$ and $B(L)$, which are of order p and q, respectively, lie outside the unit circle. Narayan and Popp (2010) consider two different specifications which are both for trending data: one allows for two breaks in levels (denoted as model 1 or $M1$) and the other allows for two breaks in levels as well as slope (denoted as model 2 or $M2$). Both model specifications differ in terms of how the deterministic component, d_t , is defined:

$$d_t^{M1} = \alpha + \beta t + \Psi^*(L)(\theta_1 DU'_{1,t} + \theta_2 DU'_{2,t}) \,, \tag{4}$$
$$d_t^{M2} = \alpha + \beta t + \Psi^*(L)(\theta_1 DU'_{1,t} + \theta_2 DU'_{2,t} + \gamma_1 DT'_{1,t} + \gamma_2 DT'_{2,t}) \,, \tag{5}$$

where $DU'_{i,t} = 1(t > T'_{B,i})$, $DT'_{i,t} = 1(t > T'_{B,i})(t - T'_{B,i})$, $i = 1,2$, $T'_{B,i}, i = 1,2$.

Here, $T'_{B,i}, i = 1,2$ denotes the true break dates. The parameters δ and κ denote the magnitude of the level and slope breaks, respectively. Narayan and Popp (2010) show that the inclusion of $\psi^*(L)$ allows breaks to occur slowly over time. Hence, the proposed model is an innovative outlier class of models, and is based on the idea that a series responds to shocks to the trend function in the same way as it reacts to shocks to the innovation process, ε_t.

The test regressions are then simply the reduced form of the corresponding structural model. They take the following forms:

$$\begin{aligned} y_t^{M1} &= \rho y_{t-1} + \alpha_1 + \beta^* t + \theta_1 D(T'_B)_{1,t} + \theta_2 D(T'_B)_{2,t} \\ &\quad + \delta_1 DU'_{1,t-1} + \delta_2 DU'_{2,t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_t \,, \end{aligned} \tag{6}$$

$$\begin{aligned} y_t^{M2} &= \rho y_{t-1} + \alpha^* + \beta^* t + \kappa_1 D(T'_B)_{1,t} + \kappa_2 D(T'_B)_{2,t} + \delta_1^* DU'_{1,t-1} \\ &\quad + \delta_2^* DU'_{2,t-1} + \gamma_1^* DT'_{1,t-1} + \gamma_2^* DT'_{2,t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_t . \end{aligned} \tag{7}$$

The null hypothesis of a unit root is tested as $\rho = 1$ against the alternative hypothesis of $\rho < 1$, based on a *t*-statistic of $\hat{\rho}$ in Eqs. (6) and (7). The break dates are selected using the sequential procedure. Readers are referred to Narayan and Popp (2010) for the details.

3.2 The TAR and MTAR Unit Root Tests

The well-known Dickey-Fuller (1981) test and its extensions assume a unit root as the null hypothesis and a symmetric adjustment process under the alternative. These tests are misspecified if the adjustment dynamics is nonlinear or asymmetric. A formal way to quantify an asymmetric adjustment process as a generalization of the Dickey-Fuller test is given by the TAR and MTAR models proposed by Enders and Granger (1998) and Enders and Siklos (2001). Let y_t denote the inflation rate and consider the following regression:

$$\Delta y_t = I_t \rho_1 \left(y_{t-1} - \tau\right) + \left(1 - I_t\right) \rho_2 \left(y_{t-1} - \tau\right) + \varepsilon_t, \tag{8}$$

where the indicator variable is defined as:

$$I_t = \begin{cases} 1, & \text{if} \quad y_{t-1} \geq \tau \\ 0, & \text{if} \quad y_{t-1} < \tau \end{cases}, \tag{9}$$

$$I_t = \begin{cases} 1, & \text{if} \quad \Delta y_{t-1} \geq \tau \\ 0, & \text{if} \quad \Delta y_{t-1} < \tau \end{cases}, \tag{10}$$

and τ denotes the value of the threshold which is derived by minimizing the residual sum of squares. Eqs. (8)–(9) represent the TAR model and Eqs. (8)–(10) the MTAR model, respectively. The MTAR model allows the speed and direction of adjustment, represented by ρ_1 and ρ_2, to depend upon the previous period's change in y_{t-1}, which is especially valuable when the adjustment is believed to exhibit more momentum in one direction than the other.

If the system is convergent, $\Delta y_t = \tau$ is the long-run equilibrium value. In the case where Δy_t is above its long-run equilibrium value, the adjustment is $\rho_1 y_{t-1}$, and if it is below its equilibrium value, the adjustment is $\rho_2 y_{t-1}$. The Dickey-Fuller test is a special case of the MTAR model (8) and (10) in the case of asymmetry in the error correction process where $\rho_1 = \rho_2$.

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

The threshold autoregressive model (TAR) allows the degree of autoregressive decay to depend on the state of the inflation rate, measuring the "deep" cycles. For instance, if the autoregressive decay is rapid when inflation is above the trend and slow when inflation is below the trend, troughs will be more persistent than peaks. Likewise, if the autoregressive decay is slow when inflation is above the trend and rapid when inflation is below the trend, peaks will be more persistent than troughs. On the other hand, the momentum threshold autoregressive model (MTAR) allows the inflation to display differing amounts of autoregressive decay depending on whether the inflation rate is increasing or decreasing, measuring the "sharpness" of cycles (Payne and Mohammadi, 2006).

3.3 The ESTAR Unit Root Test

This type of nonlinearity is related to the possibility of an *asymmetric speed of adjustment towards equilibrium*, i.e., the further the inflation rate deviates from its fundamental equilibrium, the faster will be the speed of mean reversion. This implies that the inflation rate may be a unit root process for a given threshold of values (inner regime), but a stationary process when the inflation rate reaches the outer regime.[5] In order to take account of the possibility of an asymmetric speed of adjustment towards equilibrium when testing for the unit roots, we apply the exponential smooth transition autoregressive (ESTAR) unit root tests proposed by

[5]As explained in Mourelle *et al*. (2011, p71), "Controlling for this source of nonlinearity is interesting when dealing with the inflation rate, since policy makers may decide not to react when the inflation is within range of certain values, given that the costs of any policy decision may overwhelm the benefits. However, when the inflation rate is outside a given threshold, the monetary authority might intervene in the markets in order to return the inflation rate to a more sensible value."

Kapetanios *et al*. (2003) (KSS hereafter), Rothe and Sibbertsen (2006) and Kruse (2011).

Kapetanios *et al*. (2003) have developed a new technique for the null hypothesis of a unit root against the alternative of a nonlinear but globally stationary smooth transition autoregressive process. In particular, Kapetanios *et al*. (2003) test for the null hypothesis of $\gamma = 0$ in the following model:

$$\Delta y_t = \beta y_{t-1}\left[1 - \exp\left(-\gamma y_{t-1}^2\right)\right] + \varepsilon_t \,. \tag{11}$$

The test is carried out by a *t*-test of the coefficient of y_{t-1}^3 being zero in the auxiliary regression:

$$\Delta y_t = \delta y_{t-1}^3 + \sum_{j=1}^{p} \rho_j \Delta y_{t-j} + \eta_t \,, \tag{12}$$

with the p augmentations in order to correct for serially correlated errors. The null hypothesis to be tested with Eq. (12) is $H_0 : \delta = 0$ (a unit root in the outer regime) against the alternative of $H_1 : \delta < 0$ (stationarity in the outer regime). Kapetanios *et al*. (2003) show that the *t*-statistic for $\delta = 0$ against $\delta < 0$ does not have an asymptotic standard normal distribution and they tabulate the asymptotic critical values of the *t* statistics via stochastic simulations. In the presence of constants and trends, the data are first demeaned or detrended. We refer to this test as the *K*SS nonlinear augmented Dickey-Fuller test and label it as $K\mathrm{SS}\left(\mathrm{t}_{\mathrm{NL}}\right)$.

Rothe and Sibbertsen (2006) propose a Phillips-Perron type, semi-parametric testing procedure to distinguish a unit root process from a mean-reverting exponential smooth transition autoregressive one. The test statistic is as follows:

$$Z_{NL}(t) = \frac{\hat{\sigma}}{\hat{\lambda}} t_{\hat{\beta}} - \frac{3}{2}\sum_{t=1}^{T} y_{t-1}^2 (\hat{\lambda}^2 - \hat{\sigma}^2)(\hat{\lambda}^2 \sum_{t=1}^{T} y_{t-1}^6)^{-1/2} \,, \tag{13}$$

where $\hat{\lambda}^2$ is the consistent estimator of the long-run variance $\hat{\lambda}^2$ and $\hat{\sigma}^2$ is the consistent estimator of the variance σ^2. Their simulation results show that the power of $\mathrm{Z}_{\mathrm{NL}}\left(\mathrm{t}\right)$ dominates that of $K\mathrm{SS}\left(\mathrm{t}_{\mathrm{NL}}\right)$ in the case where γ is small and where the error sequence is an MA(1).

Kruse (2011) proposes an extension of the *K*SS unit root test, which relaxes the assumption of a zero location parameter *c*, i.e., Kruse (2011) considers the following modified ADF regression:

$$\Delta y_t = \beta y_{t-1}\left[1-\exp\left(-\gamma y_{t-1}^2 - \mathrm{C}\right)\right]+\varepsilon_t , \tag{11}$$

$$\Delta y_t = \delta_1 y_{t-1}^3 + \delta_2 y_{t-1}^2 + \sum_{j=1}^{p}\rho_j \Delta y_{t-j} + \mu_t . \tag{12}$$

Eq. (15) incorporates lags of the dependent variable in order to eliminate serial correlation in the error terms. In order to test the null hypothesis of a unit root, $H_0: \delta_1 = \delta_2 = 0$, against a globally stationary ESTAR process, $H_1: \delta_1 < 0, \delta_2 = 0$. Kruse (2011) proposes a τ test, which is a version of the Abadir and Distaso (2007) Wald test.

4 Data and Results

4.1 Data Description and Basic Statistics

We use the quarterly inflation rates of six Asian countries in our empirical study, namely, India and five founding members of the Association of Southeast Asian Nations (ASEAN), i.e., Indonesia, Malaysia, the Philippines, Singapore and Thailand. We collect the data set from the International Financial Statistics (IFS) database of the International Monetary Fund. The sample period is determined primarily based on the availability of the data. For India, Indonesia, Malaysia, the Philippines, and Thailand, the data cover the period from 1970Q1 to 2014Q2, and provide a total of 178 observations. For Singapore, the sample period is from 1970Q1 to 2014Q1.

We calculate the basic summary statistics, including the sample mean, standard error, skewness and excess kurtosis, for the inflation rates of these Asian countries. The results are displayed in Table 1. Several interesting facts are observed from Table 1. First, the coefficients of skewness of all of the series are positive, implying that the inflation rates are flatter to the right compared to the normal distribution. Second, the coefficients of excess kurtosis for all of the inflation rates are much higher than zero, indicating that the empirical distributions of these samples have fat tails. The coefficients of skewness and excess kurtosis reveal non-normality in the data. This is confirmed by the Jarque-Bera normality test as shown in Table 1. Third, the Ljung-Box Q statistics, LB(24), denote significant autocorrelations for all of the series. We also report a standard ARCH test for these inflation rates. The test results indicate that a significant ARCH effect exists for all series.

Table 1: Summary Statistics

	India	Indonesia	Malaysia	Philippines	Singapore	Thailand
Mean	8.095	11.785	3.611	10.113	3.101	5.080
S.D.	5.541	11.259	3.190	9.675	4.928	5.063
SK	0.703	3.416	2.835	2.575	3.677	2.157
EK	4.481	14.429	11.027	8.511	16.765	5.576
JB	163.573**	1890.239**	1140.282**	734.002**	2471.692**	368.67**
LB(24)	359.165**	266.053**	372.966**	422.654**	391.648**	509.897**
ARCH(4)	333.528**	168.78**	388.209**	218.914**	339.405**	369.178**

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

4.2 Results from the Unit Root Tests with and without Structural Breaks

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

The results of applying these tests are reported in Table 2. We find that, with the exceptions of Indonesia and Thailand, the null hypothesis of a unit root cannot be rejected at the 5% significance level for the ADF statistics. Based on the well-known low power problem of the ADF test, we turn our attention to other statistics. The results from the DF-GLS (see Elliott *et al*., 1996) suggest that, with the exceptions of India and Indonesia, the inflation rates of these countries are nonstationary processes. However, the results of the SP test (see Schmidt and Phillips, 1992), with parametric correction, can reject the unit root hypothesis with linear trend and quadratic trend at the five percent significance level, respectively, for all of the inflation rates.[6] That is, the results from the SP tests indicate that the inflation rates for these countries are stationary processes. Based on the linear unit root test results, we cannot reach a general agreement on the stationarity of these Asian inflation rates.

Table 2: Results of the Linear Unit Root Tests

	Linear trend			Quadratic trend and breaks tests		
Country	ADF	SP(1)	DF-GLS	SP(2)	ZA	LP
India	-2.846	-4.364**	-2.913**	-4.304**	-7.282**	-6.901**
Indonesia	-3.499**	-4.512**	-3.296**	-4.500**	-7.655**	-8.051**
Malaysia	-2.841	-3.643**	-2.634	-3.995**	-6.146**	-6.175
Philippines	-2.190	-4.400**	-2.865	-4.716**	-8.849**	-7.990**
Singapore	-2.066	-3.898**	-2.449	-3.770**	-6.865**	-7.725**
Thailand	-3.664**	-3.400**	-2.303	-3.573**	-6.568**	-6.878**

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

A possibility for getting mixed results from the linear unit root tests is that we ignore the feature of structural breaks.[7] As Perron (1989) pointed out, in the presence of a structural break, the power to reject a unit root decreases if the stationary alternative is true and the structural break is ignored. To address this, we

[6]The terms SP(1) and SP(2) denote the Schmidt-Phillips *t* tests with the linear and quadratic trend, respectively.

[7]For example, Malliaropulos (2000) finds that the US inflation rate is better characterized by trend stationarity with a structural break.

use Zivot and Andrews' (1992) (hereafter ZA) sequential one trend break model and Lumsdaine and Papell's (1997) (hereafter LP) two trend breaks model to investigate the order of the empirical variables. We use the 't-sig' approach proposed by Hall (1994) to select the lag length (*k*). We set $k_{\max} = 12$ and use the approximate 10% asymptotic critical value of 1.60 to determine the significance of the t-statistic on the last lag. We use the 'trimming region' [0.10T, 0.90T] and select the break point endogenously by choosing the value of the break that maximizes the ADF t-statistic. We report the results in Table 2. The results from the ZA and LP tests suggest that, with the exception of Malaysia for the LP test, the null hypothesis of a unit root can be rejected at the 5% significance or better level for all of the inflation rates, implying that we could wrongly conclude that the inflation rate is a unit root process but in fact it is a stationary process with a structural break.

Lee and Strazicich (2003) shows that the ZA and LP unit root tests, which do not allow for a break under the null hypothesis, suffer from severe spurious rejections in finite samples when a break is present under the null hypothesis. The root of the problem of spurious rejections is that parameters of the test regression have different interpretations under the null and alternative hypotheses, which is crucial since the parameters have implications for the selection of the structural break date. Following Schmidt and Phillips (1992), Popp (2008) and Narayan and Popp (NP, 2010) propose a new endogenous structural break unit root test by formulating the data-generating process (DGP) as an unobserved components model to avoid the above problem. Narayan and Popp (2013) show that the NP test has better size properties and identifies the breaks more accurately than its main two-break unit root rivals; namely, the Lumsdaine and Papell (1997) and Lee and Strazicich (2003) tests. In this study, we also test for inflation persistence by employing Popp's (2008) one-break and Narayan and Popp's (2010) two-break unit root tests. We consider Model 1 (two breaks in the intercept) and Model 2 (two breaks in the intercept and trend) and extract appropriate critical values from Popp (2008) and Narayan and Popp (2010), respectively. We report the Popp (2008) and NP test results for the inflation rates in Tables 3 and 4, respectively.

Beginning with the Popp (2008) one-break test (Table 3), we find that we are able to reject the unit root null hypothesis for the inflation rates of all countries at the 5% significance level or better. In the case of the NP test (Table 4), again, the null hypothesis of a unit root is rejected at the 5% level for Model 1 and Model 2. The results of the Popp (2008) and the NP tests echo the results of the ZA and LP tests, suggesting that any shock has a transitory effect on inflation. The mean reversion in Asian inflation rates is in favor of the hypothesis of, for example, the natural rate of inflation and the sticky-price model.

Table 3: Results of the Popp One-Break Unit Root Test

	Model 1			Model 2		
Country	Test statistic	TB1	*k*	Test statistic	TB1	*k*
India	-4.998**	1988.4	5	-5.107**	1988.4	5
Indonesia	-5.810**	1998.2	5	-5.420**	1998.2	5
Malaysia	-4.696**	2008.3	5	-4.686**	2008.3	5
Philippines	-6.520**	1984.2	5	-6.656**	1984.2	5
Singapore	-6.886**	1974.3	5	-6.381**	1974.4	5
Thailand	-4.409**	1974.4	5	-5.204**	1974.2	5

Notes: (1)*, **, *** denote significance at the 10%, 5% and 1% levels, respectively.

Table 4: Results of the Narayan and Popp Two-Break Unit Root Test

	Model 1				Model 2			
Country	Test statistic	TB1	TB2	*k*	Test statistic	TB1	TB2	*k*
India	-5.042**	1988.4	1998.4	5	-5.263**	1988.4	1998.4	5
Indonesia	-8.367**	1997.4	1998.2	5	-7.593**	1997.4	1998.2	5
Malaysia	-5.616**	1980.4	1984.1	5	-5.630**	1980.4	1990.3	5
Philippines	-7.709**	1983.4	1984.2	5	-7.535**	1983.4	1984.2	5
Singapore	-5.697**	1981.4	1982.1	5	-6.673**	1979.1	1981.1	5
Thailand	-5.408**	1979.4	1980.2	5	-5.401**	1979.2	1980.2	5

Notes: (1) *, **, *** denote significance at the 10%, 5% and 1% levels, respectively.

4.3 Results Based on the Nonlinear Unit Root Tests

Another possibility for the mixed results from the linear unit root test is that we ignore the nonlinearity of the inflation rate. Parallel to Perron (1989), the simulation results from Pippenger and Goering (1993) suggest that the test is inclined to accept the null hypothesis of a unit root if the nonlinear property is ignored. That is, the probability of a type II error will increase. In order to validate the nonlinear unit root used in this paper, we conduct several nonlinearity tests for these inflation rates. Psaradakis *et al*. (2001) examine the relative performance of some popular nonlinearity tests. The nonlinearity tests considered include the RESET-type tests, the Keenan test, the Tsay test, the McLeod-Li test, the BDS test, the White dynamic information matrix test, and the neural network test. We adopt these statistics to examine whether any nonlinearity exists in the inflation rate. The results are reported in Table 5, which shows that, with the exception of the Keenan (1985) test, almost all of the *p*-values of these nonlinear tests are below the 5% significance level or better for all inflation rates considered in this paper, indicating that the inflation rates of these Asian countries are characterized by nonlinearity.

Table 5: *p*-values for a Battery of Nonlinear Tests

	India	Indonesia	Malaysia	Philippines	Singapore	Thailand
RESET1	0.0493	0.002	0.041	0.005	0.000	0.0204
RESET2	0.003	0.001	0.050	0.001	<0.001	0.001
KEENAN	0.394	0.083	0.351	0.23	0.817	0.804
TSAY	0.002	<0.001	0.066	<0.001	<0.001	0.013
MCLEOD	0.392	<0.001	0.002	<0.001	<0.001	<0.001
BDS	0.007	<0.001	<0.001	<0.001	<0.001	0.000
WHITE1	0.052	0.140	0.017	0.006	0.004	0.011
WHITE2	0.000	<0.001	<0.001	<0.001	<0.001	0.000
NEURAL1	0.026	0.019	0.108	0.344	0.004	0.027
NEURAL2	0.123	0.058	0.038	0.475	0.136	0.038

Notes: (1) RESET1: Ramsey and Schmidt (1976). (2) RESET 2: Thursby and Schmidt (1977). (3) KEENAN: Keenan (1985). TSAY: Tsay (1986). (4) MCLEOD: McLeod and Li (1983). (5) BDS: Brock *et al*. (1996). (6) WHITE1 and WHITE2 are White's (1987) information matrix tests. (7) NEURAL1 and NEURAL2 are the neural networks proposed by White (1989a,b). (8) < 0.001 indicates that the number is less than 0.001.

Next we test for inflation persistence by considering two different types of nonlinear unit root tests. First, following the literature,[8] we adopt threshold unit root

[8]For example, Henry and Shields (2004) find the two-regime threshold unit root process in Japan and

tests, proposed by Enders and Granger (1998), in examining the stationarity of Asian inflation rates. We report the results for the demeaned, as well as demeaned and detrended data for inflation rates (y_t) based on the following reason: if there is a time trend in the data and the regression equation does not contain a trend term, then the test has low power. On the other hand, if the regression equation contains a trend term but a trend does not exist in the data, then the null hypothesis is rejected too often. The inflation rate is demeaned by regressing y_t on a constant, *C*, and, alternatively, demeaned and detrended, *C*, *T*, by regressing y_t on a constant, as well as a linear trend prior to estimation in the TAR and MTAR regression equations. Hence, we allow for a constant term and a linear trend as attractors. We perform the tests with a linear time trend included due to its possible impact on the properties of the tests.

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

The results of applying the TAR tests for the demeaned as well as the demeaned and detrended data of the inflation rates of six Asian countries are reported in Table 6. For the demeaned data (columns (1)–(6) in Table 6, the F_C statistics are significant at the 5% or better level, indicating that the null hypothesis of a unit root ($H_0 : \rho_1 = \rho_2 = 0$) in the inflation rates of all countries must be rejected. Because the null hypothesis of a unit root is rejected using the TAR specification, the null hypothesis of symmetry where $\rho_1 = \rho_2$ is tested. The F_A statistics show that this hypothesis is rejected at the 5% significance level for Indonesia, Malaysia and the Philippines, therefore indicating the presence of an asymmetric adjustment phenomenon.

For the demeaned and detrended data in the bottom panel of Table 6, the null hypothesis of a unit root ($H_0 : \rho_1 = \rho_2 = 0$) in the inflation rates of six Asian countries must be rejected at the 5% significance level. Thus, for those countries in

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which the null hypothesis of a unit root is rejected using the TAR specification, the null hypothesis of symmetry where $\rho_1 = \rho_2$ is tested. The F_A statistics show that this hypothesis is rejected at the 5% significance level for Indonesia, Malaysia, the Philippines, Singapore and Thailand in Table 6, therefore indicating the presence of an asymmetric adjustment phenomenon. Taking Singapore as an example, the point estimates suggest that the inflation tends to decay at the rate of $|\rho_1| = 0.2.375$ for y_{t-1} above the attractor 30.5162 − 0.0566*t* and at the rate of $|\rho_2| = 0.005$ for y_{t-1} below the attractor. In regard to the speeds of adjustment, it appears that $|\rho_1|$ is greater than $|\rho_2|$. This is evidenced by the null hypothesis of symmetry ($\rho_1 = \rho_2$) which is rejected. Thus, the evidence suggests that the 'deep' cycles (adjustments) around the threshold value of the inflation are asymmetric.[9]

Table 6: Results of the TAR Unit Root Test

	Demeaned data					
	India	Indonesia	Malaysia	Philippines	Singapore	Thailand
F_C	21.783**	21.807**	18.870**	25.028**	17.585**	12.437**
F_A	0.838	16.833**	21.980**	26.114**	2.388	0.224
	[0.361]	[0.000]	[0.000]	[0.000]	[0.124]	[0.637]
ρ_1	-0.172**	-0.27**	-0.0784**	-0.261**	-0.132**	-0.101**
ρ_2	-0.222**	-0.064**	-1.0730**	-0.048	-0.070**	-0.126**
	Demeaned and detrended data					
	India	Indonesia	Malaysia	Philippines	Singapore	Thailand
F_C	21.887**	30.518**	22.037**	32.863**	28.166**	20.429**
F_A	0.666	54.017**	28.412**	0.314	54.416**	17.022**
	[0.416]	[0.001]	[0.000]	[0.576]	[0.000]	[0.000]
ρ_1	-0.175**	-0.516**	-0.0697**	-0.238**	-2.375**	-0.252**
ρ_2	-0.218**	-0.0221	-1.6172**	-0.208**	-0.005	-0.055**

Notes: (1) *, **, *** denote significance at the 10%, 5% and 1% levels, respectively. (2) F_C and F_A denote the *F*-statistics for the null hypothesis of a unit root $H_0: \rho_1 = \rho_2 = 0$ and symmetry $H_0: \rho_1 = \rho_2$, respectively. (3) The 10%, 5% and 1% critical values for the F_C statistic for the demeaned data are 3.74, 4.56 and 6.47, respectively. (4) The 10%, 5% and 1% critical values for the F_C statistic for the demeaned and detrended data are 5.18, 6.12 and 8.23, respectively. (5) The numbers in parentheses are standard errors. (6) The numbers in square parentheses are *p*-values.

Next, the results of the MTAR unit root test for these Asian inflation rates are

[9]The estimates of the attractors are available from the authors upon request.

reported in Table 7. It is shown that, for all countries, the null hypothesis of a unit root can be rejected at the 5 percent significance level or better for the demeaned data. Moreover, the asymmetric adjustment hypothesis is supported for India, Malaysia and the Philippines, since the F_A statistic is rejected at the 5% significance level. For the demeaned and detrended data, the null hypothesis of nonstationarity is rejected at the 5% significance level for all countries, implying that the inflation rates of the six Asian countries are stationary processes. In addition, the asymmetric adjustment hypothesis is supported for India, Malaysia and the Philippines since the F_A statistic is rejected at the conventional level. Thus, it appears that the 'sharpness' cycles (adjustments) around the threshold value of the inflation rates for India, Malaysia and the Philippines are asymmetric.

Table 7: Results of the MTAR Unit Root Test

	Demeaned data					
	India	Indonesia	Malaysia	Philippines	Singapore	Thailand
F_C	25.09**	20.524**	16.963**	27.13**	18.516**	14.201**
F_A	5.835**	1.428	8.597**	10.407**	1.663	3183
	[0.017]	[0.234]	[0.004]	[0.002]	[0.199]	[0.076]
ρ_1	-0.137**	-0.229**	-0.039	-0.072**	-0.095**	-0.070**
ρ_2	-0.261**	-0.145**	-0.196**	-0.24**	-0.157**	-0.150**
	Demeaned and detrended data					
	India	Indonesia	Malaysia	Philippines	Singapore	Thailand
F_C	25.188**	22.100**	22.793**	36.894**	20.593**	18.507**
F_A	6.673**	0.763	11.106**	6.641**	0.760	3.462
	[0.011]	[0.384]	[0.001]	[0.011]	[0.385]	[0.065]
ρ_1	-0.131**	-0.236**	-0.077	-0.155**	-0.114**	-0.092**
ρ_2	-0.264**	-0.174**	-0.262**	-0.292**	-0.155**	-0.179**

Notes: (1)*, **, *** denote significance at the 10%, 5% and 1% levels, respectively. (2) F_C and F_A denote the F-statistics for the null hypothesis of a unit root $H_0: \rho_1 = \rho_2 = 0$ and symmetry $H_0: \rho_1 = \rho_2$, respectively. (3) The 10%, 5% and 1% critical values for the F_C statistic for the demeaned data are 4.05, 4.95 and 6.91, respectively. (4) The 10%, 5% and 1% critical values for the F_C statistic for the demeaned and detrended data are 5.64, 6.65 and 8.85, respectively. (5) The numbers in parentheses are standard errors. (6) The numbers in square parentheses are p-values.

Table 8: Results of the ESTAR-Type Unit Root Tests

		$KSS(t_{NL})$	
Country	raw data	demeaned	detrended
India	-3.494**	-3.943**	-4.055**
Indonesia	-7.282**	-7.340**	-7.088**
Malaysia	-4.081**	-4.211**	-4.502**
Philippines	-6.816**	-8.500**	-6.662**
Singapore	-6.430**	-6.564**	-6.728**
Thailand	-4.817**	-5.136**	-5.506**
		$Z_{NL}(t)$	
Country	raw data	demeaned	detrended
India	-2.234**	-2.186	-2.234**
Indonesia	-2.922**	-2.867	-2.782**
Malaysia	-1.965	-1.931	-2.073**
Philippines	-2.406**	-2.247	-2.297**
Singapore	-2.022	-2.050	-2.129**
Thailand	-2.210	-2.325	-2.562**
		τ	
Country	raw data	demeaned	detrended
India	12.410**	17.049**	17.463**
Indonesia	53.375**	53.953**	50.073**
Malaysia	16.568**	25.169**	25.423**
Philippines	46.193**	46.276**	50.115**
Singapore	42.983**	50.878**	52.166**
Thailand	24.156**	34.630**	31.982**

Notes: (1) *, **, *** denote significance at the 10%, 5% and 1% levels, respectively. (2) $KSS(t_{NL})$: Kapetanios *et al*. (2003). $Z_{NL}(t)$: Rothe and Sibbertsen (2006). *t*: Kruse (2011). (3) The critical values for the three statistics are obtained from Kapetanios *et al*. (2003) and Kruse (2011).

Finally, we examine the size nonlinearity which is related to the possibility of an asymmetric speed of adjustment towards equilibrium. That is, the further the inflation rate deviates from its natural rate or fundamental equilibrium, the faster will be the speed of mean reversion.[10] We apply the $KSS(t_{NL})$ (Kapetanios *et al*., 2003), $Z_{NL}(t)$ (Rothe and Sibbertsen, 2006) and τ (Kruse, 2011) statistics to the raw data, as well as the demeaned and detrended data of the inflation rates. The results are reported in Table 8, and with the exceptions of $Z_{NL}(t)$, we see that the results of the $KSS(t_{NL})$ and *t* statistics point to the rejection of the null hypothesis of a unit root against the alternative of a globally stationary ESTAR process in the inflation rates of six Asian countries. This implies that the size nonlinearity is a vital feature of the inflation rates for India, Indonesia, Malaysia, the Philippines, Singapore and Thailand. If we overlook this feature, then we will be inclined to reach a spurious conclusion that the inflation is a nonstationary process. In fact, it is a nonlinear mean-reverting process that favors the natural rate hypothesis and the sticky-price model.

4.4 Test Allows for Multiple and Smooth Breaks

Previous tests assume that there are at most two structural breaks and a sharp break. Since the number of structural breaks could be more than two and switches are smooth rather than abrupt, in order to account for these features, we have decided to employ the Fourier augmented Dickey-Fuller unit root test, proposed by Christopoulos and Leon-Ledesma (2011) and Enders and Lee (2012), in this study.[11] Several seminal studies, such as Becker *et al*. (2004) and Enders and Lee (2012), illustrate that a Fourier series approximation can capture the behavior of an unknown function, even if the function itself is not periodic. Moreover, the flexible Fourier testing framework requires only the specification of the proper frequency in the estimating equations. Chang *et al*. (2013) have applied the Fourier unit root test to detect the stationarity properties of 22 OECD countries and have found that mean-reverting inflation holds in all 22 OECD countries.

Here we briefly illustrate the Fourier ADF test. Readers are referred to Christopoulos and Leon-Ledesma (2011) and Enders and Lee (2012) for the details.

[10] As explained in Mourelle *et al*. (2011, p71), "Controlling for this source of nonlinearity is interesting when dealing with the inflation rate, since policy makers may decide not to react when the inflation is within range of certain values, given that the costs of any policy decision may overwhelm the benefits. However, when the inflation rate is outside a given threshold, the monetary authority might intervene in the markets in order to return the inflation rate to a more sensible value."

[11] We owe this point to an anonymous referee.

Following Enders and Lee (2012), we consider the following regression:

$$y_t = \delta_0 + \delta_1 t + \alpha_k \sin(2\pi kt / T) + \beta_k \cos(2\pi kt / T) + \rho y_{t-1} + \varepsilon_t , \tag{13}$$

where k is the Fourier frequency. We are interested in testing the null hypothesis of a unit root ($H_0 : \rho = 1$) against the alternative hypothesis of stationarity ($H_0 : \rho < 1$). Since the frequency (k) that provides the best fit to the nonlinear trend is unknown, Enders and Lee (2012) suggest a grid search — specifically, selecting a value of k over $1 \le k \le 5$ that minimizes the sum of squared residuals from Eq. (13). In our application we use an interval for $k = [0.1, 0.2, ..., 4.9, 5]$. The resulting test is denoted by $FDF(\hat{k})$. Moreover, Enders and Lee (2012) suggest pretesting for nonlinearity via the usual *F*-statistic ($F(\hat{k})$) for the null hypothesis of linearity (i.e., $\alpha_k = \beta_k = 0$) in (13) to determine whether $FDF(\hat{k})$ should be used.

The results of the Fourier ADF test are summarized in Table 9. The significant $F(\hat{k})$ statistic as shown in the third column of Table 9 indicates that both sine and cosine terms should be included in the estimated model. For Malaysia, the Philippines, Singapore and Thailand, the non-integer value of $\hat{k}$ generates permanent breaks in the series. For India and Indonesia, the integer value of $\hat{k}$ denotes only temporary breaks in the series. The results from the Fourier ADF statistics, i.e., $FDF(\hat{k})$, suggest that allowing for nonlinearity and structural breaks results in the rejection of the unit root null hypothesis for all six Asian countries. The empirical findings echo the findings of the TAR, MTAR and ESTAR unit root tests.

The corresponding time series plots of the inflation rates are graphed in Figure 1. Visual inspection of Figure 1 clearly highlights the explosion in inflation experienced during the 1970s and early 1980s caused by the first and second oil crises. Accordingly, it appears sensible to allow for structural breaks in testing for a unit root. The estimated time paths of the time-varying intercepts are also shown in Fig. 1. Further examination of these graphs indicates that all of the Fourier approximations seem to be reasonable and support the notion of long swings in inflation rates.

Table 9: Results of the Fourier ADF Unit Root Tests

Country	$\hat{k}$	$F(\hat{k})$	Lags	$FDF(\hat{k})$
India	5	9.661**	6	-5.056**
Indonesia	2	19.859**	6	-5.356**
Malaysia	0.7	16.299**	6	-5.356**
Philippines	0.7	29.059**	6	-6.180**
Singapore	0.8	15.850**	6	-5.756**
Thailand	0.8	19.610**	6	-5.069**

Notes: (1) *, **, *** denote significance at the 10%, 5% and 1% levels, respectively. (2) $\hat{k}$ denotes optimal frequency. (3) $F(\hat{k})$ indicates that the test is distributed as a F statistic under the null hypothesis with two degrees of freedom. The critical values are taken from Table 1 of Becker *et al*. (2006). (4) The optimal lag was selected optimally using the BIC.

5 Concluding Remarks

The purpose of this paper is to study the property of persistence of six Asian inflation rates. A variety of unit root tests ranging from univariate estimators to nonlinear testing principles are employed in an effort to obtain inferences that are robust to problems associated with nonstationarity. The empirical results show that we cannot reach a unanimous conclusion on the stationarity of the inflation rates by using a battery of linear unit root tests. We are of the opinion that the mixed results are due to ignorance regarding structural breaks and nonlinearity in testing for the unit root. Hence, on the one hand, we adopt newly-developed structural break unit root tests (Popp, 2008; Narayan and Popp, 2010) to re-examine the inflation persistence. On the other hand, we consider two types of nonlinearities, i.e., threshold and exponential smooth transition, in this study. For this purpose, we adopt the TAR, the MTAR, the ESTAR-type unit root tests which help detect the nonlinear inflation relationship without specifying the threshold in advance.

For the benefit of readers, we summarize our empirical results in Table 9. We reach the following key conclusions. First, the results from the Popp (2008) and Narayan and Popp (2010) tests show that the inflation rates are stationary after taking account of structural breaks, indicating that we are inclined to accept the null hypothesis of the unit root if we overlook the structural break. Second, the results from the TAR and MTAR unit root regressions suggest that the inflation rates are stationary series in the long-run with asymmetric adjustment in the short-run. Third, empirical evidence from the ESTAR-type unit root tests favors a globally stationary ESTAR process in the inflation rates of six Asian countries, implying that the further the inflation rate deviates from its natural rate, the faster will be the speed of mean reversion.

To sum up, the results of the structural break unit root tests or nonlinear unit root tests provide strong evidence that the inflation rates are mean reversion processes. These facts imply that structural breaks and nonlinearity (threshold or exponential smooth transition) are important properties in testing for inflation persistence because such a test is inclined to accept the null hypothesis of a unit root if structural breaks or nonlinearity are ignored. Therefore, an adequate assessment of the persistence of shocks to inflation should account for the inherent structural breaks or nonlinearity in the data.

Table 10: Summary of a Variety of Nonlinear Unit Root Tests

Country	SB	TAR(asymmetry)	MTAR(asymmetry)	ESTAR	FADF
India	yes	yes(no)	yes(yes)	yes	yes
Indonesia	yes	yes(yes)	yes(no)	yes	yes
Malaysia	yes	yes(yes)	yes(yes)	yes	yes
Philippines	yes	yes(yes)	yes(yes)	yes	yes
Singapore	yes	yes(yes)	yes(no)	yes	yes
Thailand	yes	yes (yes)	yes(no)	yes	yes

Notes: (1) SB denotes the unit root test with a structural break proposed by Popp (2008) and Narayan and Popp (2010). (2) The term “yes (no)” indicates that the null of a unit root is rejected and in favor of a stationary process with symmetric adjustment. (3) The term “yes (yes)” indicates that the null of a unit root is rejected and in favor of a stationary process with asymmetric adjustment. (4) TAR and MTAR denote the nonlinear unit root tests proposed by Enders and Granger (1998). (5) FADF denotes the Fourier augmented Dickey-Fuller unit root tests.

In line with the discussions in Chang *et al*. (2013), a major policy implication of the present study is that a stabilization policy may not have a permanent effect on the inflation rates of these six Asian countries (India and ASEAN5) under study. Our empirical findings are essentially in favor of the hypothesis of, for example, the natural rate of inflation and the sticky-price model, indicating that any shock has a transitory effect on inflation. Therefore, it would be less costly to exercise the policies of disinflation for the monetary authorities of the sample countries than for those of the countries with nonstationary inflation.

It is worth noting that some recently related works on the nonlinear unit root tests, for example Chen *et al*. (2013), Chen and Lee (2015), and Chen *et al*. (2015) have been proposed in the literature. These three papers handle three types of nonlinear unit root tests for high-frequency time series based on the Bayesian approach. Chen *et al*. (2013) propose a Bayesian test to detect for the unit root in multi-regime threshold autoregression with heteroskedasticity. Chen and Lee (2015) propose a Bayesian unit root to detect the presence of a local unit root vs. mean-reverting nonlinear smooth transition heteroskedastic alternative hypothesis. Chen *et al*. (2015) propose a Bayesian hypothesis test to detect the presence of a local unit root in the mean equation using Markov switching GARCH models. It is still important to apply these new tests to test for the nonstationary properties of macroeconomic or financial variables. This is beyond the scope of the present work. We leave this as an avenue for future research.

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Appendix

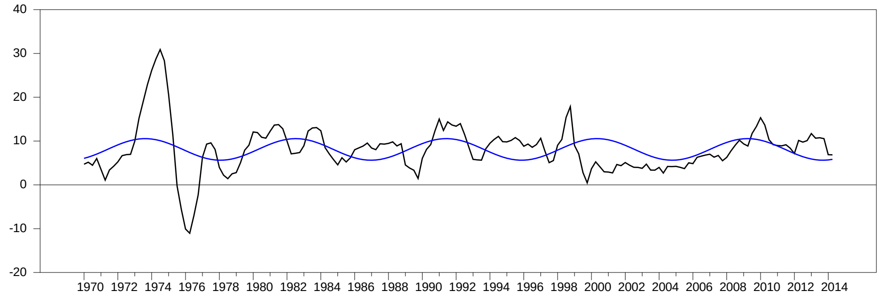

Figure 1: Time Series Plot of India's Inflation Rate (Black) and Fitted Nonlinear Trend (Blue)

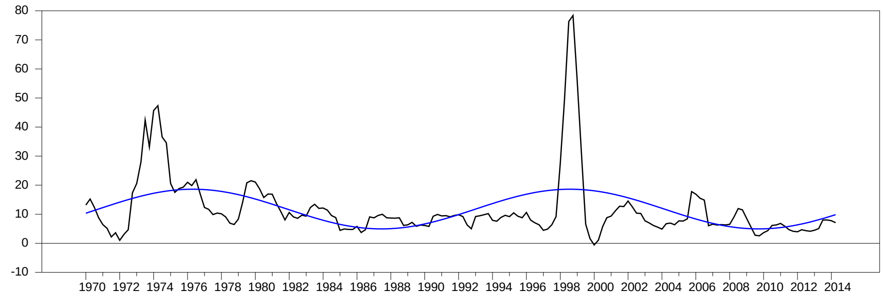

Figure 2: Time Series Plot of Indonesian Inflation Rate (Black) and Fitted Nonlinear Trend (Blue)

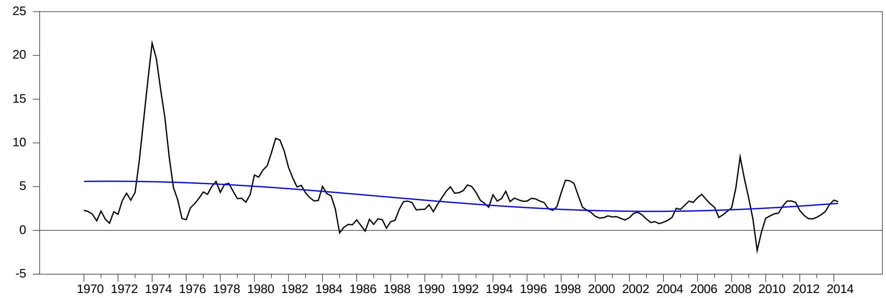

Figure 3: Time Series Plot of Malaysian Inflation Rate (Black) and Fitted Nonlinear Trend (Blue)

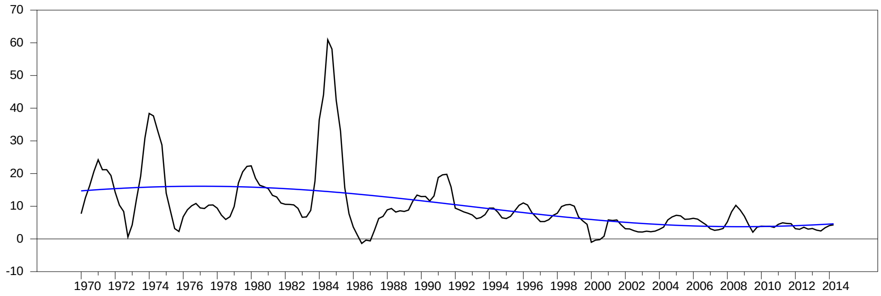

Figure 4: Time Series Plot of the Philippines' Inflation Rate (Black) and Fitted Nonlinear Trend (Blue)

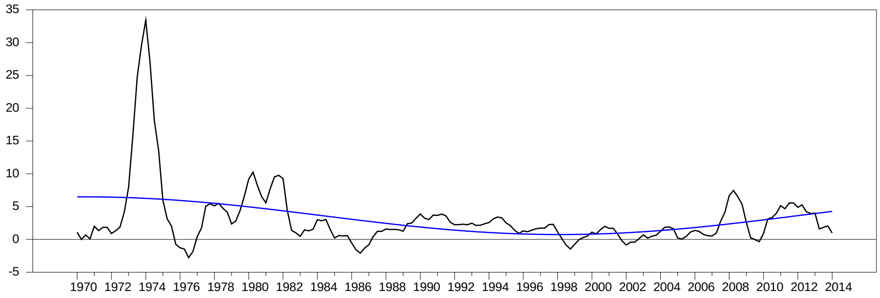

Figure 5: Time Series Plot of Singapore's Inflation Rate (Black) and Fitted Nonlinear Trend (Blue)

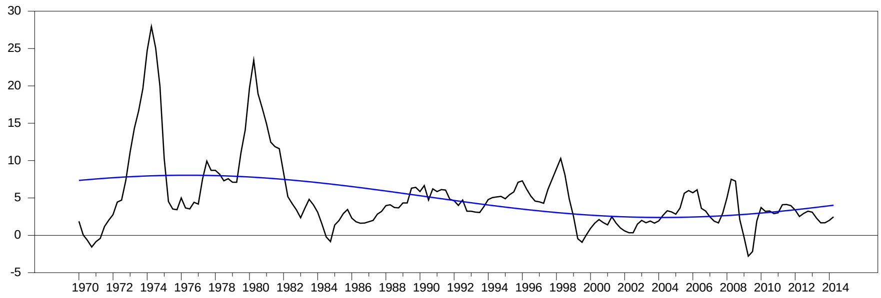

Figure 6: Time Series Plot of Thailand's Inflation Rate (Black) and Fitted Nonlinear Trend (Blue)

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