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Is the Hybrid New Keynesian Phillips Curve Stable? Evidence from Some Emerging Economies

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ABSTRACT: *One of the central issues in macroeconomics is the relationship between inflation and unemployment / output gap, called the 'Phillips curve'. Of the several version of this relationship, the latest one is known as 'hybrid new Keynesian Phillips curve' (HNKPC). The primary focus of this paper is to check whether the HNKPC holds or not for four important emerging economics viz., Brazil, Russia, India and South Africa. This has been empirically examined after testing for the structural stability of this relationship so that it can be studied accordingly with due consideration to this important issue. Some of the econometric issues like the unit root tests and estimation of output gap have also been done appropriately. Our findings suggest that the HNKPC is not stable for all the four countries. Finally, the analysis based on the two sub-periods thus formed clearly shows mixed evidence in respect of holding of this relationship.*

KEY WORDS: Inflation; generalised method of moments; Phillips curve; structural break

JEL CLASSIFICATION: C32; E31

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1. Introduction

One of the central issues in macroeconomics is to deal with the relationship between inflation and unemployment or, more generally, between inflation and a measure of real economic activity such as output gap. This relationship, known, more widely, as 'Phillips curve', was first empirically tested by Phillips (1958) using the U.K. data. Subsequently, the original version of the Phillips curve was sharply criticized by Friedman (1968) and Phelps (1968) because of the fact that there is no consideration to inflation expectation term in this formulation. According to them, once this expectation is allowed in the inflation dynamics, permanent trade-off between inflation and output gap no longer exists. However, a short-run relation continues to exist due to the presence of lags in the adjustment of expectations. This is typically known as the 'traditional Phillips curve' which relates current inflation to lagged inflation and output gap.

It is documented that the dynamics of inflation over most of the post-second World War period till around mid-seventies of the last century, is captured quite well by the traditional Phillips curve for a number of industrialised countries. Subsequently, however, two issues have emerged in the context of studying the Phillips curve. First, the traditional Phillips curve is subject to the Lucas critique (1976) in that its coefficients may not be invariant across policy regimes. Second, according to Sargent (1971), the treatment of expectation in traditional Phillips curve is not sufficient to capture the forward looking rational behaviour of an individual. Due to these shortcomings, the so called 'New Keynesian' economists challenged the concept of the traditional Phillips curve and came up with a new version of it, known as the 'new Keynesian Phillips curve' (NKPC). The NKPC uses micro foundation to derive relationship between inflation, expectation of future inflation and the current value of the cyclical indicator. The important characteristic of NKPC is that it can be derived from the optimal price setting behaviour of firms. According to the model proposed by Calvo (1983), firms set their prices optimally subject to a constraint on the frequency of price adjustments. Hence, the parameters can be viewed as the structural parameters, thus providing some immunity to the Lucas critique.

One problematic feature of the NKPC, from consideration of economic theory, is that a fully credible central bank can manage a disinflation of any size without bearing a cost in terms of output losses. But the empirical applications of the historical data support the view that disinflation requires a recession. Also, Clarida and Gertler (1997) have shown that countries with highly credible central banks have experienced very costly disinflation. Due to the empirical limitations stated above, a number of studies following Gali and Gertler (1999) have considered a hybrid model that incorporates expected future inflation and lagged inflation in the inflation dynamics. The inclusion of the lagged terms is for the purpose of capturing the past dynamics of inflation. The hybrid version of NKPC, called the hybrid new Keynesian Phillips curve (HNKPC), incorporates both the traditional Phillips curve and the pure forward-looking NKPC as the two extreme cases. Consequently, empirical verification of the HNKPC is a crucial issue.

It may, however, be noted that most of the literature on Phillips curve are concerned with studying this relationship for the developed economics only. It is only very recently that similar studies on developing and emerging countries have gained momentum. In this study, our main purpose is to verify if HNKPC holds for four important emerging economics *viz*., Brazil, Russia, India and South Africa. To state it differently, this paper examines the relative importance of past and expected future inflation along with output gap, to explain the dynamics of inflation for the aforesaid four countries. Although, these four countries are now being regarded as important emerging economies, there have been very few serious studies on the HNKPC involving these countries. It is, therefore, our understanding that it would be of interest to verify if this relationship holds for these countries or not. It is also to be mentioned that some econometric issues like testing of stationarity in consideration to structural breaks, proper estimation of output gap in presence of structural breaks and the stability of HNKPC relationship have been duly considered in this study using recent advances made on these issues, which are not quite so in the existing studies.

To be more specific on the latter points, our study takes a closer look at the issue of stationarity or nonstationarity of the time series on inflation and output of the countries concerned. This issue has an econometric relevance since the previous studies focusing on the estimation of the NKPC considers that the inflation series follows a stationary process (see, for instance, Rudd and Whelan (2007), and Nason and Smith (2008)), whereas some recent works (see, for example, Lee and Nelson (2007), and Goodfriend and King (2009)) have modeled the Phillips curve by assuming a stochastic process for inflation and provides mixed evidence regarding the unit root condition of the series. It may be noted that the time span of the data sets used in this study being 17 years (from May 1994 to May 2011), it is quite probable that structural breaks have occurred in the time series, and hence this should be tested applying the recent advances made in the literature on unit root tests. To that end, we have performed the structural break test due to Perron and Yabu (2009) to detect the presence of a structural break, if any, in the trend function of a time series. In case this test suggests that there is one break in the deterministic trend, we have applied the test developed by Kim and Perron (2009) where the null hypothesis is that of a unit root with a break against the alternative of stationarity with a break in the deterministic trend function. We have applied these two tests to inflation and output series to find whether any shock really persists or it is transitory in nature around a break in the deterministic trend of the relevant series. By applying these tests to our data sets, we have found that except India inflation series of each of the other three countries *viz*., Brazil, Russia and South Africa, follows a stationary process having a break in their respective deterministic trend. For India, no break in deterministic trend has been found for inflation. However, for the output series, the above tests suggest that the series are nonstationary with a deterministic trend break for all the four countries.

Taking the findings on output series into consideration, we have estimated the output gap, where the estimation procedure endogenously incorporates the structural break in the trend function. It is known that the Hodrick-Prescott (HP) filter is the favourite alternative technique among researchers for studying the output gap. However, different shortcomings and drawbacks of the HP filter have been pointed out in the literature. According to Cogley and Nason (1995), the HP technique extracts a smooth estimate of trend from actual output series and it can thus lead to an excessive smoothing over structural break. In the entire literature on structural breaks, the change in the dynamics of a variable is often modelled as being sudden. In the above situation when there exists a structural break in the trend function (intercept and/or slope coefficient), the HP technique may provide some distorted results. One would, therefore, like to have a detrending procedure that is able to account for such features in an endogenous fashion. In our analysis, we have estimated the potential output using a modified Hamilton Markov regime switching model due to Lam (1990). This model has an advantage that it endogenously accounts the trend break in the estimation procedure of output gap.

It is well documented that because of its derivation from a model with optimizing agents, the HNKPC has a strong theoretical advantage over the backward looking Phillips curve. On the theoretical ground, the parameters of the HNKPC can be considered to be structural ones or these are invariant to different policy regimes and thus immune to the Lucas critique. But a statistical rechecking is needed to test the stability of HNKPC relationship. To the best of our knowledge, only a limited number of studies deal with this issue of stability of HNKPC. Among these, the study by Jondeau and Bihan (2005) have used the approach developed by Andrews (1993), and Andrews and Ploberger (1994) to test for the parameter stability of the hybrid model. Based on a regime switching model, Kim and Kim (2008) have also studied the structural break issue in the HNKPC analysis. However, in this paper we have used the structural break test suggested by Andrews (1993) to detect whether the HNKPC relationship is structurally stable or not for each of these four countries.

The organisation of the paper is as follows. Section 2 briefly describes the HNKPC. The methodology along with the econometric issues involved are discussed in Section 3. Empirical findings are discussed in Section 4. The paper ends with some concluding remarks in Section 5.

2. The Hybrid New Keynesian Phillips Curve (HNKPC)

The theory of Phillips curve began with the publication of the seminal work of A.W.Phillips in 1958. Phillips empirically established a negative relationship between wage inflation and unemployment rate in the U.K. However, studies covering subsequent time periods have shown that this relationship either does not exist or it provides weak explanations to inflation dynamics. This led researchers to look into it more carefully resulting in some concerns being found with regard to this relationship. Subsequently, the original 'Phillips Curve' was modified a number of times and the latest version, called the Hybrid New Keynesian Phillips Curve (HNKPC), is now used to study inflation dynamics. One very influential paper in the literature on HNKPC is due to Gali and Gertler (1999), where the specification depends on optimal price setting behaviour of firms. Gali and Gertler (1999) rationalized the hybrid model by assuming that there exist two types of firms. A fraction of firms set their price optimally subject to constraint on the frequency of price adjustments, as in Calvo's (1983) model, while the remaining firms use a rule of thumb or behave in backward-looking fashion in price setting. In Christiano et al. (2001), all firms adjust their prices at each period, but some of them are not able to re-optimize their prices so that their prices are indexed to last period's respective inflation rates. Thus, in the general framework of HNKPC, inflation can be expressed as a function of expected future inflation, lag inflation and real marginal cost.

Rotemberg and Woodford (1997) showed that under certain assumptions, the output gap is linearly related to the real marginal cost, thus providing a rationale for introducing the output gap in the Phillips curve equation. Again, Fuhrer and Moore (1995), Fuhrer (1997), and Neiss and Nelson (2002) have also used the output gap as a proxy for marginal cost. Thus, by using output gap as a proxy of real marginal cost, the hybrid specification of Phillips curve is written as follows:

$$\pi_t = \gamma_f E_t(\pi_{t+1}) + \gamma_b \pi_{t-1} + \lambda x_t + u_t \qquad (2.1)$$

where π_t denotes inflation at time t, x_t is the output gap, $E_t(\pi_{t+1})$ is the expectation at time t about the inflation at the future time point $t+1$ and π_{t-1} the lag inflation.

3. Methodology

In this section, the method of estimation of output gap, HNKPC and the stability of HNKPC are discussed.

3.1 Estimation of output gap by modified Hamilton Markov switching model

The output gap i.e., the gap between the actual output and the potential output is the key variable determining the evaluation of price behaviour. In terms of macroeconomics concepts, this is essentially the business cycle. It is argued that output gap is a summary indicator of the relative demand and supply components of the economic activity where supply side effects are captured by potential output and demand side effects by output gap. The prevailing economic thinking during 1960s was that supply side economy was deterministic, and accordingly all movements in real output about the time trend were interpreted as results of demand shock.

It is only after the influential work by Nelson and Plosser (1982) that trend in economic time series started being modelled as stochastic trend with or without a deterministic trend, as the case may be. This implies that movement in output can occur as a result of shocks to aggregate demand as well as to potential output. During the last two decades, a number of statistical techniques have been developed to extract stochastic trend from real output. Among these the well-known Hodrick-Prescott filter (HP) is used extensively. However, as pointed out by Cogley and Nason (1995), one serious drawback of the HP filter is that it can lead to an excessive smoothing over structural break(s). According to an empirical study by Perron and Wada (2009), a single break in trend can affect the outcome of various detrending procedures and such a change can drastically alter the resulting cyclical component. Hence, a detrending procedure would be more powerful if it is able to incorporate the statistical feature of structural breaks in an endogenous fashion. On the other hand, the modified Hamilton Markov switching model, due to Lam (1990), has an important advantage that it estimates the output gap by endogenizing the structural break feature in the estimation procedure.

According to Hamilton (1989), a seasonally adjusted time series is perceived as a sum of two independent unobserved components, trend and cycle, the latter being the measure of output gap. Further, it is assumed that one of the components follows a random walk process with a drift which evolves according to a two-state Markov process and the other follows an autoregressive process with a unit root. Lam (1990) generalized this model by relaxing the unit root assumption so that the autoregressive process may be stationary or it may contain a unit root.

The model proposed by Lam (1990) involves the following equations:

$Y_t = Y_t^* + c_t,$

$Y_t^* = \alpha_0 + \alpha_1 s_t + Y_{t-1}^*,$

and $\phi(B)c_t = \varepsilon_t$

where s_t, Y_t^*, and c_t are the state variable, trend and cyclical component, respectively and $\varepsilon_t \sim N(0,\ \sigma^2)$. Here, the trend component Y_t^* is assumed to follow a random walk with a drift, which evolves according to a two-state Markov process. The binary state variable s_t represents either a high or a low growth state of the economy at time *t*. The transition probability i.e., the probability that state 'i' is followed by state 'j' is denoted as

$prob(s_t = i | s_{t-1} = j) = p_{ij}, i, j = 0, 1$ where $\sum_j p_{ij} = 1.$

If one of the roots of $\phi(B) = 0$ is unity, then Y_t^* and c_t are not identified, and the model collapses to the original Hamilton model. In most of such studies (see, Kim (1994) and Bautista (2003), for instance) where such decomposition is considered, the cyclical component is assumed to be driven by a second order autoregressive process so as to cover a wide array of possible dynamics.

Kim (1994) revised Lam's (1990) specification by using a state-space technique, where the first equation is the measurement equation and the second and third are the transition equations. Kim has shown that his smoothing algorithm yields approximate maximum likelihood estimates that are very close to Lam's exact maximum likelihood estimates.

In this study, we have used Kim's algorithm for the modified Hamilton Markov switching model because it is easy to use and it reduces computational time substantially.

3.2 Estimation and stability of the HNKPC

The hybrid Phillips curve, as specified in equation (2.1), tells us that current inflation depends on expected future inflation, lagged inflation and output gap. Obviously, the expected future inflation is an unobserved variable. Replacing the expected future inflation term in the above equation by realized inflation and adjusting it accordingly, we get,

$$\pi_t = \gamma_f \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda x_t + \varepsilon_t, \tag{3.1}$$

where $\varepsilon_t = u_t - (\pi_{t+1} - E_t(\pi_{t+1}))$.

This leads to an estimation problem arising out of correlation between the regressors and the error. Consequently, as a solution to this estimation problem, the generalised method of moments (GMM) is used.

Now, the orthogonality condition required for the GMM method of estimation involves the error term ε_t and a $(k \times 1)$ vector of instruments, say Z_t, and this condition is $E(\varepsilon_t Z_t) = E\big(g_t(\theta)\big) = 0$ where $g_t(\theta) = \big(\pi_t - \gamma_f \pi_{t+1} - \gamma_b \pi_{t-1} - \lambda x_t\big) Z_{t-1}$ and $\theta = (\gamma_f \quad \gamma_b \quad \lambda)'$ denotes the vector of unknown parameters. An efficient GMM estimator of θ is obtained by minimizing the following with respect to θ:

$$\bar{g}_T(\theta)' \left(S_T\big(\tilde{\theta}_T\big)\right)^{-1} \bar{g}_T(\theta)$$

where $\bar{g}_T(\theta) = (1/T) \sum_{t=1}^{T} g_t(\theta)$, $S_T\big(\tilde{\theta}_T\big)$, the weighting matrix, is a consistent estimator of the covariance matrix $\sqrt{T}\bar{g}_T(\theta)$ at $\theta = \tilde{\theta}_T$ and $\tilde{\theta}_T$ is a consistent estimate of θ.

Estimation of parameters in equation (3.1) depends on k orthogonality conditions. Among the several methods available for the GMM estimation we have used the two-step GMM estimator proposed by Hansen (1982). An asymptotic estimator of the weighting matrix $S_T\big(\tilde{\theta}_T\big)$ is found by choosing a consistent estimator of V where $V = E[\bar{g}_T(\theta)\bar{g}_T(\theta)']$ is the long-run covariance matrix of $g_t(\theta)$. If it is assumed that the error term is heteroscedastic and autocorrelated, then the covariance matrix is consistently estimated by the Newey and West (1987) procedure.

When the number of orthogonality conditions i.e., the number of instruments, exceeds the number of parameters to be estimated, the model is said to be over-identified and the GMM estimation will provide biased estimates. To avoid the bias originated by the use of an excessive number of instruments, we have applied the Newey-West (1987) test for over-identification. This test is based on the minimized value of the objective function, the *J*-statistic. Under the null hypothesis that the over-identifying restrictions are satisfied, the value of the *J*-statistic times the number of observations asymptotically follows a chi-square distribution with degrees of freedom equal to the number of over-identifying restrictions.

We now briefly describe the stability test that has been used for studying the stability of HNKPC. To that end, we have applied the Wald test statistic, as developed by Andrews (1993), for testing the structural stability of the HNKPC model with an unknown change point. The advantage of this test is that it can be applied to a wide class of parametric models that are suitable for GMM estimation. The procedure assumes that the break can occur at any time point in the period $[\pi_0 T,\ (1-\pi_0)T]$, where π_0 is a chosen fraction of the total observations *T*. For each of these probable break points, we estimate the HNKPC for the periods before and after the break. The 'Sup Wald' test statistic of Andrews (1993) is defined as

$$sup\ W_T = sup_{\pi\epsilon\,|\pi_0,(1-\pi_0)|}\, W_T(\pi) \tag{3.2}$$

where

$$W_T(\pi) = T(\hat{\theta}_1(\pi) - \hat{\theta}_2(\pi))' \left(\frac{\hat{V}_1(\pi)}{\pi} + \frac{\hat{V}_2(\pi)}{1-\pi}\right)^{-1} (\hat{\theta}_1(\pi) - \hat{\theta}_2(\pi)), \tag{3.3}$$

and $\hat{\theta}_1(\pi)$ and $\hat{\theta}_2(\pi)$ are the two estimates of the parameter vector based on observations for the two periods, before and after a candidate break point, and $\hat{V}_i(\pi)$, i =1,2, denote the estimated variance-covariance matrix of the estimated parameters for the two sub-periods. The asymptotic distribution of the Wald test statistic is nonstandard, and its asymptotic critical values have been provided by Andrews (1993, 2003). In this test, the test statistic is obtained as a function of all possible break dates. This sequence of test statistic values are plotted against the candidate break points, and then checked if the maximum of these values has exceeded the appropriate critical value. In case it so happens, then it is concluded that the relationship has a structural break.

4. Empirical Results

4.1 Data

This empirical study is based on monthly time series data for the period from May 1994 to May 2011 for four most important emerging economies *viz*., Brazil, Russia, India and South Africa. All the data sets for Brazil, Russia and South Africa have been downloaded from the official website of Federal Reserve Bank of St. Louis (*http://research.stlouisfed.org/fred2/categories/32263*) while for India the source is the official website of Reserve Bank of India (RBI) (*www.rbi.org.in*). Except India, consumer price index (CPI) is used to obtain rate of inflation, and it is constructed as the monthly difference of the logarithm values of CPI. In case of India, wholesale price index (WPI) is used instead of CPI.

As argued by Agenor et al. (2000), output in the industrial sector corresponds roughly to output in the traded goods sector, and hence it is closely related to business cycle shocks of a country. Thus, the index of industrial production series can be considered to be a good proxy for measuring aggregate cycle behaviour of any country. Accordingly, we have used the monthly series of index of industrial production as a measure of real economic activity. Other than inflation and output gap, the instrument set for estimation of the HNKPC includes monetary growth and exchange rate growth. We have used broad money supply (M3) and real effective exchange rate to compute monetary growth and exchange rate growth, respectively. Time series data for all these variables have been downloaded from the two websites mentioned earlier.

4.2 Results of the tests on presence on structural break and unit roots

To test for stationarity or otherwise of a time series, most often the augmented Dicky-Fuller (ADF) (1979) test is used, where the null hypothesis of nonstationarity in the sense of being integrated or having unit roots tested against the alternative of stationarity. However, due to the influential work by Perron (1989), the commonly used ADF test has been criticised because of its bias towards non-rejection of the null hypothesis of a unit root against the alternative of trend stationarity in the presence of a structural break in the deterministic trend, and also for its low power for near integrated process. Subsequently, Perron (1989, 1990) proposed an alternative unit root test which allows the possibility of a structural change in the trend function under both the null and alternative hypotheses. This means that if a change/break is present, it is allowed under both the null and alternative hypotheses, and thus improves the power of the test as against the usual ADF test. However, one serious limitation to this is that it is based on the assumption of a known break date. Subsequently, Zivot and Andrews (1992), Perron (1997) and Vogelsang and Perron (1998), among others, have treated the break date to be unknown, which can be endogenously determined from the model. However, Kim and Perron (2009) have pointed out that in all such studies, a trend break is not allowed under the null hypothesis. Those tests consider a break in the time series under the alternative only. Hence, tests in this line of work are inferior in terms of size and power. To overcome this limitation, Kim and Perron (2009) have developed a new test procedure on the line of Perron's (1989) original formulation of trend break being allowed under both the null and alternative hypotheses, but the break date is now assumed to be unknown. Since the existence of a structural break in the trend function is a problem of long-horizon data – we propose to apply this unit root test of Kim and Perron because of its distinct advantage over the ADF type unit root tests.

Now, prior to applying this unit root test, it is most important to have the information on whether or not a structural break is present in a given time series especially when the break date is assumed to be unknown as in the case with the Kim-Perron (2009) test. Further, it may be noted in this context that in the absence of a trend break, the well-known ADF test has the highest power than any other alternative test and is most appropriate for testing the presence of unit roots in a time series. It is all the more important that the knowledge on presence or absence of a break in the trend function is available before deciding on the appropriate unit root tests. However, as pointed out by Perron and Yabu (2009), testing for structural break in the trend function depend on whether the noise component is stationary or nonstationary having unit roots. Thus there is some sort of a circular problem in testing for the twin issues. To deal with this, recently Perron and Yabu (2009) have proposed a test for structural change in the trend function of a univariate time series, which can be performed without any prior knowledge on whether the noise component is stationary or nonstationary containing unit roots[1]. Since this test is very general in its approach insofar as the assumption on noise is concerned, we have performed this test to detect the presence of structural break,

[1] Perron and Yabu (2009) have considered a quasi-feasible generalized least squares technique that uses a super-efficient estimate of the sum of autoregressive parameters, which governs the stationary or integrated behaviour of a time series.

if any, in the trend function of a time series. Thus, in case the Perron-Yabu test suggests that there is no break in the deterministic trend, we apply the ADF test; otherwise, we use the Kim-Perron test to test the null hypothesis of unit roots against the alternative of stationarity with a break in the deterministic trend function under both the hypotheses[2].

It may be noted that three models - *Model I*, *Model II* and *Model III* - representing a single change in intercept only, a one-time change in the slope of linear trend without a change in level, and a simultaneous change in the intercept and slope coefficients, respectively, are considered for the Perron-Yabu test. For this test, we have chosen the trimming parameter to be 0.15. Since the relevant test statistic for *Model III* has the highest power against the alternatives, we have considered this model only. Test statistic values are reported in Table 1.

Table 1: Results of the Perron-Yabu test on structural break and the Kim-Perron test on unit roots on inflation and output

Country	*Perron-Yabu structural break test*	*Kim-Perron unit root test*	
	Statistic value for Model III	*Statistic value for Model A3*	*Statistic value for Model A3(trimmed)*
		Inflation	
Brazil	18.93*	-6.81*[0.26]	-5.57*[0.26]
Russia	117.29*	-5.61*[0.15]	-5.53*[0.15]
India	1.55	--	--
South Africa	5.68*	-6.01*[0.51]	-5.86*[0.51]
		Output	
Brazil	16.47*	-4.01[0.57]	-3.44[0.57]
Russia	21.62*	-2.89[0.33]	-2.06[0.33]
India	9.64*	-2.74[0.73]	-3.34[0.73]
South Africa	4.86*	-2.85[0.85]	-2.48[0.85]

[indicates significance at 5% level of significance. The fractions in parentheses denote the estimated break fractions.]*

For inflation, the Perron-Yabu test statistic values clearly reject the null hypothesis of $H_0: \mu_1 = \beta_1 = 0$ for Brazil, Russia and South Africa. Hence, it is concluded that there is a

[2] A GAUSS program code for Perron and Yabu (2009) structural break test and a MATLAB code for Kim and Perron (2009) unit root test in presence of a structural break have been taken from the official website of Pierre Perron.

structural break in inflation in case of these three countries. For India, however, the test statistic value is 1.55 which is not significant even at 10% level of significance. On the other hand, the Perron-Yabu test statistic values for output series are found to be significant for all the four countries, and thus we can conclude that there exists a structural break in the deterministic trend in the output series of all the four countries.

As mentioned earlier, the ADF unit root test is the most appropriate test if there is no structural break, and accordingly for India's inflation series only. For the three remaining inflation series and all output series, Kim and Perron (2009) unit root test has been carried out.

In the ADF test, the estimating equation includes both the intercept and a deterministic linear trend, and the value of the optimum lag has been obtained by applying the Schwarz's information criterion (SIC). The computed value of -9.749 of the ADF test statistic for India's inflation is highly significant at 1% level of significance. Further, the coefficient associated with the linear trend term has also been found to be insignificant. Hence, we conclude that the time series on India's inflation is stationary.

After having these empirical finding that there is a break in the trend function, we have applied the third variant of the additive outlier model, *Model* A3[3], as described in Kim and Perron (2009), to inflation (for Brazil, Russia and South Africa) and output (for all the four countries) series for testing the null hypothesis of unit root with a deterministic trend break against the alternative of stationarity with a break in the deterministic trend. The values of this statistic are reported in Table 1. The computed values of *t*-statistic are compared with the appropriate critical values available in Perron (1989) and Perron and Vogelsang (1993). The findings clearly suggest that inflation of each of Brazil, Russia and South Africa has no unit root, whereas each of the output series follows a nonsationary process with a break in their respective deterministic trend part. Since the inflation of Brazil, Russia and South Africa are trend stationary processes (TSP). Hence, these series have been transformed to stationary series by using appropriate (deterministic) trend removal procedure.

[3] Kim and Perron (2009) have considered three types of additive outlier models *viz*., A1, A2 and A3, representing the occurrence of a structural break only in the intercept, only in the slope, and both in the intercept and slope coefficients of a time series, respectively.

4.3 Estimation of output gap

In this section, we report results of estimation of output gap for each of the four countries following the method discussed in Section 3.1. It may be stated that in the modified Hamilton model, a regime switching behaviour has been introduced, and it takes care of the feature of sudden changes in the trend function and then yields the estimate of potential output accordingly.

A constrained maximum likelihood method has been used to obtain the maximized value of the log likelihood function subject to p_{00} and p_{11} being constrained to lie between 0 and 1, and σ^2 and α_1 being positive while the roots of the autoregressive components are left unconstrained. The estimated covariance matrix has been computed numerically from the negative inverse of the Hessian of the log-likelihood evaluated at these estimates. A GAUSS program code which is available on the website of Kim and Nelson has been used for these computations. Table 2 below presents the results of estimation of output gap by the modified Hamilton Markov regime switching model.

Table 2: Parameter estimates of the modified Hamilton Markov switching model

Country *Parameter*	*Brazil*	*Russia*	*India*	*South Africa*
p_{00}	0.937*	0.970*	0.972*	0.963*
	(0.022)	(0.021)	(0.024)	(0.018)
p_{11}	0.252***	0.825*	0.985*	0.663*
	(0.136)	(0.107)	(0.012)	(0.154)
ϕ_1	0.570*	0.542*	0.493*	0.318*
	(0.103)	(0.093)	(0.069)	(0.079)
ϕ_2	-0.081*	0.018	0.345*	-0.025***
	(0.030)	(0.082)	(0.069)	(0.013)
σ	1.925*	1.931*	1.367*	1.863*
	(0.117)	(0.113)	(0.068)	(0.099)
α_0	-4.465*	-1.594*	0.420*	-1.996*
	(0.522)	(0.330)	(0.028)	(0.429)
α_1	5.030*	2.130*	0.560*	2.359*
	(0.527)	(0.336)	(0.086)	(0.429)

[, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively. Figures within parentheses indicate standard errors of the estimates.]*

We observe from this table that all the parameters are statistically significant. The estimates of p_{00} and p_{11} are very close to 1 for Russia, India and South Africa indicating that the processes are very persistent in nature. But the estimate of p_{11} for Brazil is 0.252, and hence it suggests that if the output series falls in the high growth state at any time point, the probability of remaining in that state in the next time period is very low. Further, the estimates of all the autoregressive coefficients are found to be much less than unity. The sum of ϕ_1 and ϕ_2 is also less than unity for each of these countries, and hence we can conclude that there is no explosive cycle visible in any of the output series.

Figure 1 gives the plots of the estimated smoothed probability of a low growth state. From the diagram it can be concluded that the model is well characterized by recurrent shifts between high and low growth periods.

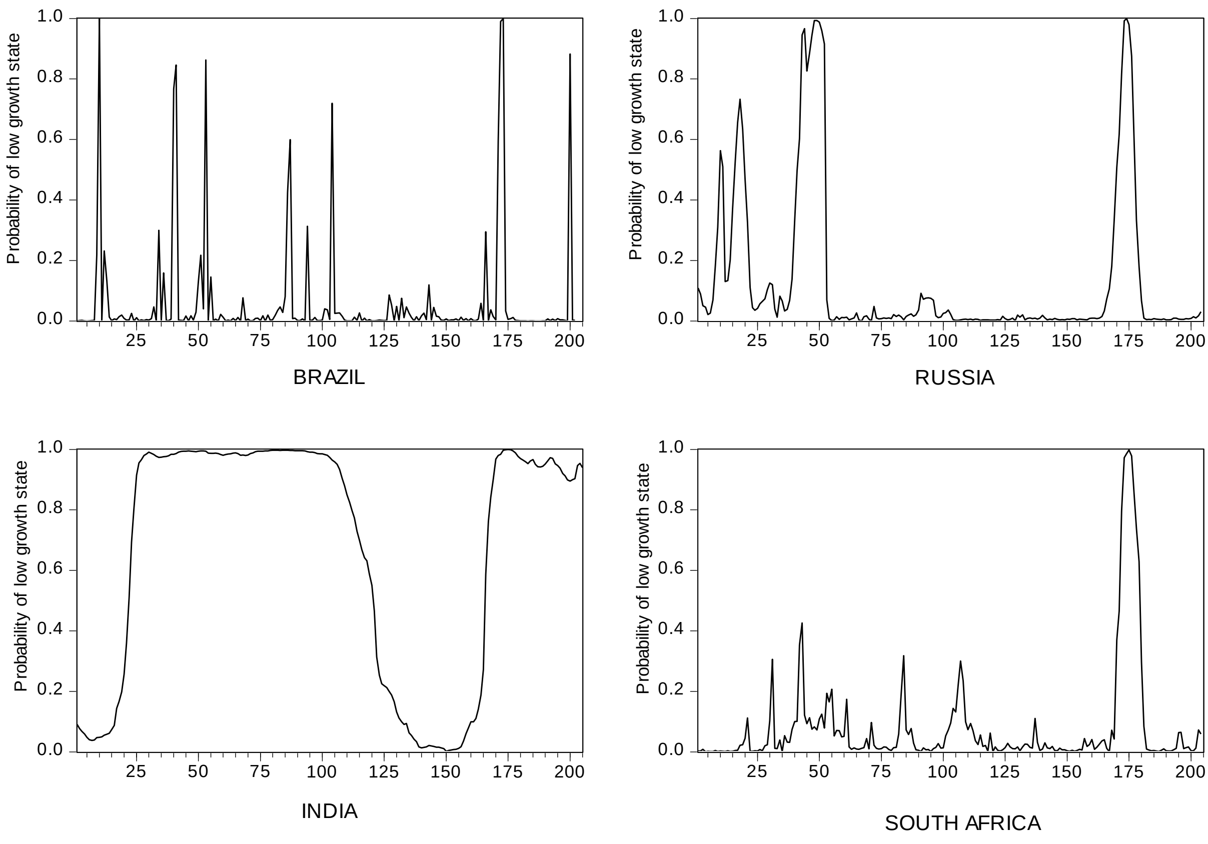

Figure 1: Smoothed probability of a low growth state

4.4 Stability of the HNKPC relationship

We first report the results of GMM estimation of the parameters of the HNKPC relation. The GMM estimation results are reported in Table 3.

Table 3: Full sample parameter estimates of the HNKPC model

Country / *Parameter*	*Brazil*	*Russia*	*India*	*South Africa*
γ_f	0.608* (0.079)	0.766* (0.071)	0.685* (0.046)	0.379* (0.109)
γ_b	0.387* (0.042)	0.218* (0.052)	0.290* (0.040)	0.339* (0.062)
λ	-0.009 (0.007)	-0.023 (0.020)	0.001 (0.011)	0.010 (0.019)
J-statistic	10.17	7.61	8.54	13.83
p-value of J-statistic	0.75	0.91	0.86	0.46

[, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively. Figures in parentheses indicate standard errors of the estimates.]*

It is noted from the entries of this table that both the parameters of HNKPC *viz.*, γ_f and γ_b are statistically significant for all the four countries. Thus, both forward as well as backward looking behaviours are found to be important in studying the inflation dynamics for the four countries. However, the coefficient of output gap, i.e., λ is found to be insignificant for all the countries.

Now, as pointed out by Jondeau and Bihan (2005), the stability of HNKPC may suffer due to regime changes in policy, and hence a test is required to check for the structural stability of the estimated relationship. Keeping this in mind, we have carried out Andrews' sup-Wald test to detect the structural change, if any, in this relationship. It may be noted that the maximum values of the four sequences of Wald test statistics for Brazil, Russia, India and South Africa have been obtained as 102.22, 103.60, 31.62 and 26.10, respectively. Since the critical value is 16.65 at 1% level of significance, we can conclude that the HNKPC model is structurally unstable for each of these countries. The estimated break dates have been found to be May 1998, December 1998, March 2008 and June 2008 for Brazil, Russia, India and South Africa, respectively.

The estimates of the break dates for these countries are found to be quite in order in terms of major policy initiatives undertaken by these governments during the period under study. For instance, in case of Brazil, the occurrence of this structural break quite correctly captures the change in its monetary policy regime in early 1999. As is officially recorded, the Central bank of Brazil adopted a pegged exchange rate regime in 1999 instead of a fixed exchange rate system in order to target the inflation rate for attaining a better macroeconomic stability, and this led to a more stable inflation regime during 2000s as compared to the period of 1990s. It is well-known that the Asian Financial Crisis of 1997 caused a negative shock throughout emerging market economies, and Russia was no exception. As pointed out by Esanov et al. (2005), this external cause decreased investment confidence in Russia, resulting in a large capital outflow. In the presence of large capital outflows, the Central bank of Russia attempted to preserve the financial system by injecting liquidity into the market. As a result, this excess liquidity led to an inflationary pressure during late 1990s. In the post-1998, the central bank of Russia tightened monetary policy to reduce the hike in inflation. Furthermore, due to the improvements in fiscal measurements and revenue collection, the fiscal performance of Russia improved significantly. As a result, during the early part of 2000s Russian inflation rate fell sharply and got stabilized. The timings of structural changes in case of India and South Africa are found to have coincided with the beginning of global financial crisis of 2008. Thus the possible explanations of the breaks in these two countries may be attributed to the consequences of this global crisis on the important macro variables including inflation.

Based on these estimated break dates, we have divided the entire time period of each country into two sub-periods. For instance, in case of Brazil the first sub-period spans from May 1994 to May 1998 and the second one from June 1998 to May 2011. Now for each sub-period thus obtained for all the four countries, we have estimated the HNKPC relationship, and the results are reported in Table 4.

Table 4: Estimates of parameters of the HNKPC model in the two sub-periods

Country *Parameter*	*Brazil*	*Russia*	*India*	*South Africa*
		First sub-period		
γ_f	0.381* (0.061)	0.530* (0.080)	0.711* (0.056)	0.285** (0.138)
γ_b	0.281* (0.041)	0.196* (0.062)	0.280* (0.049)	0.334* (0.061)
λ	-0.074* (0.011)	0.049 (0.058)	-0.007 (0.015)	0.036 (0.031)
J-statistic	8.99	8.99	13.61	10.43
p-value of J-statistic	0.83	0.83	0.48	0.73
		Second sub-period		
γ_f	0.524* (0.086)	0.497* (0.069)	0.462* (0.050)	0.588* (0.058)
γ_b	0.441* (0.043)	0.445* (0.038)	0.399* (0.036)	0.435* (0.035)
λ	-0.003 (0.005)	0.001 (0.011)	0.013 (0.011)	0.042* (0.009)
J-statistic	10.56	6.15	7.41	7.55
p-value of J-statistic	0.72	0.96	0.92	0.91

[, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively. Figures in parentheses indicate standard errors of the estimates.]*

It is evident from the above tables that the coefficients γ_f and γ_b are statistically significant in both the sub-periods for all the countries. It is also noted that in the first sub-period, the estimate of the forward-looking component i.e., γ_f, is larger than that of the backward-looking component, γ_b, for Russia and India. For instance, the estimates of γ_f and γ_b for Russia are 0.530 and 0.196, respectively. However, for Brazil and South Africa these estimates do not differ much. The *p*-values of *J*-statistic suggest that over-identifying restrictions are satisfied for all the countries and hence the models are correctly specified. It is also worthwhile to note that, for all countries, the coefficient of backward looking component increases uniformly in the second period than in the first. The estimate of γ_f increases from the first sub-period to the second sub-period only for Brazil and South Africa whereas it decreases in the second period in case of Russia and India. In the second sub-period, the estimated values of γ_f and γ_b do not differ much from each other for all the countries. This suggests that the backward-looking component is equally important to the forward-looking component in understanding inflation dynamics for all of them. Finally, insofar the coefficient of output gap is concerned, the parameter i.e., λ is found to be statistically insignificant in all but one sub-period for each of Brazil and South Africa. In case of Brazil, it is significant and negative in the first sub-period. But this parameter is significant and positive for South Africa in the second sub-period. This is in contrast to the finding of non-existence of any such relationship between inflation and output gap for the estimation results on the basis of entire time series.

5. Concluding Remarks

The aim of this paper was to verify whether the HNKPC holds for four important emerging economics *viz*., Brazil, Russia, India and South Africa. However, in so doing, we have given due importance to the issue of stability of HNKPC and accordingly applied the Andrews (1993) test so that the empirical verification of HNKPC can be done in the light of stability of this relationship. To start with, we first took a closer look at the stationarity condition of the time series on inflation and output for all the four countries. The test for stationarity has been carried out by applying the newly developed tests by Perron and Yabu (2009) and Kim and Perron (2009). The results show that, except India, inflation in the other three countries are trend stationary processes with a break in their respective trend functions. In case of India, it is a trend stationary process without a break. Regarding the output series, the above tests suggest that the output series for all the four countries have unit roots with a deterministic trend break. We have applied the modified Hamilton Markov switching model to calculate the output gap for each country. Using the output gap thus obtained, we have estimated the HNKPC taking in to account the stability aspect of this relationship.

Regarding the stability of HNKPC, it has been found that there exists a structural break in the above relationship for all the four countries. Accordingly, the time period was divided into two sub-periods and HNKPC was fitted separately to each sub-period. The findings show that both the lead and lag coefficients are highly significant for the two sub-periods, indicating that both the forward and backward looking behaviour are important to understand the dynamics of inflation of these four countries. However, the coefficient associated with output gap was found to be insignificant in most of the cases except for Brazil and South Africa. In case of Brazil the output gap is significant only in the first sub period with a negative sign, and it thus invalidates the typical Phillips curve theory that the output gap will positively affect inflation. And, the coefficient associated with output gap was found to be significant and positive for South Africa, which indicates that the short run relation between inflation and output gap exists only for the recent period in South Africa.

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