Core inflation for India

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Abstract

In this study, we construct a variety of core inflation measures using exclusion method, limited influentual methods and a common trends model. The estimated core inflation measures are put to empirical evaluation of how well they satisfy certain desirable properties of core inflation. The evidence reveal that the core inflation based on exclusion and limited influentual methods do not seem to be consistently satisfying certain important criteria under consideration. Nevertheless, the core inflation estimated from a common trends model is found to be unbiased to headline inflation, less volatile, highly correlated with growth rate of M3 money, cointegrated with headline inflation, and a powerful attractor of headline inflation. The evidence suggests that overemphasizing exclusion based core measures for its simplicity does not rule out the potential use of model based and economically interpretable measures of core inflation for policy purposes.

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1. Introduction

In recent years, the literature on measuring low frequency component of headline inflation or constructing a measure of core inflation gained importance due to growing reliance of central banks on an explicit inflation targeting policy to ensure price stability. It is widely recognized that there is a mismatch between what is conceptually known as inflation – a sustained rise in the general price level – and the measured headline inflation, which is noisy. Moreover, the noise elements in observed inflation should not affect policy decisions as they are purely transitory phenomenon reflecting mostly shocks in relative prices. In this context, empirical estimates of core inflation, which is free of transitory price fluctuations, turned out to be crucial in policy designs that aim at price stability.

Broadly speaking, the empirical studies largely arrive at a measure of core inflation using either exclusion method that identifies and eliminates unwanted noise elements of headline inflation or smoothing methods. Under exclusion method, core inflation is constructed by ignoring those commodities whose prices are highly volatile. The rationale behind this method is to retain only those goods in the price index that are not affected by any exogenous shocks. Although such measures are widely used for policy purposes, they suffer from severe criticism on the ground that excluding prices in *ad hoc* manner might result in loss of potential information about the underlying inflation.

Other methods involve smoothing actual inflation using moving average method, Kalman filter, Hodrick-Prescott filter etc. Bryan and Cecchetti (1993) proposed a method in which they trim the tails of the cross sectional price variations in the price index to overcome the disturbances in the distribution due to relative price shocks. However, Quah and Vahey (1995) criticized all these methods arguing that they devoid of theoretical base and deserve little economic interpretation.

Instead, Quah and Vahey (1995) formulated a bivariate structural vector autoregression with output and inflation and imposed an identifying restriction consistent with vertical

Phillips curve that the permanent inflation shocks does not have any medium to long run impact on real output. Blix (1995) extended this approach in a common trends framework that accommodates cointegrated variables. Recently Bagliano and Morana (2003a,b) used a multivariate common trends method to construct core inflation for UK and US. Their argument towards a multivariate model is that the structural shocks can be precisely identified in a multivariate model than in a bivariate model. They interpreted core inflation as the long run forecast of the inflation conditional on the information and the long run cointegration properties contained in the data. Wynne (1999) evaluated various methods and bestow his consent on Quah and Vahey approach on two aspects among others. It has a theoretical basis and forward-looking characteristics.

Although the movement of annual point-to-point percentage change in wholesale price index (WPI) serves as an official measure of inflation, the importance of identifying an appropriate measure of core inflation has been often emphasized (Reddy, 1999) in India. Of late, the Reserve Bank of India began to observe the movements of alternative core inflation measures (RBI Annual Report, 2000). However, the empirical studies on measuring core inflation are scanty in India. There are few studies that attempted to examine the relevance of core measures using conventional methods. The study by Samanta (1999) computed core inflation following exclusion method. The study arrived at a measure of core inflation by excluding primary food and non-food articles and few commodities whose prices are administered with a combined weight of 46 percent in WPI. Although, the core measure displayed lesser volatility during 1993-94 to 1998-99 its relevance as an indicator of inflation in a developing country was questioned. For instance, Mohanty, Rath and Ramaiah (2000) argued that a large array of commodities show relative price volatility over time; hence, it is not appropriate to remove them all from core measure. The basket of commodities whose prices are volatile is not time invariant; therefore, excluding certain commodities permanently will result in loss of information about underlying inflation. Moreover, primary commodities are sizable in the consumption basket of consumers and their influence on underlying inflation is substantial. Therefore, core measure that excludes primary commodities cannot be a good indicator of inflation.

Mohanty, Rath and Ramaiah (2000) used trimmed mean method for monthly WPI series and found that 20 percent trimmed mean of WPI is a valid core inflation measure for India. Joshi and Rajpathak (2004) used new series on WPI and constructed core inflation following exclusion and trimmed mean methods. They conclude that WPI excluding food and 20 percent trimmed mean satisfy most of desirable properties of the core measure.

However, there is no empirical study that exploits the recent developments in the empirical literature to construct an appropriate measure of core inflation for India. Moreover, none of the studies could bring out a convincing measure of core inflation that satisfies all the desirable properties. In this regard, this study makes a comprehensive attempt to examine the relative merits of certain conventional measures vis-à-vis economically interpretable measure of core inflation for India. The rest of the paper is organized as follows: Section 2 deals with various methods for constructing core inflation. Section 3 presents estimates of various core measures. Section 4 evaluates their relative merits in terms of certain desirable properties of core inflation. Section 5 provides concluding remarks.

2. The Methodology

In this study, we adopt three different methods to construct core inflation: (i) exclusion method; (ii) limited influence method; and (iii) common trends method. Under exclusion method, prices of certain commodities are assigned zero weights in the construction of core inflation as they are identified as highly volatile and mostly driven by supply shocks. The use of exclusion based core inflation for policymaking is very popular since it is very simple to construct and easy to understand. In general, food and energy commodities are considered for exclusion as their prices are largely supply driven and subject to shocks due to frequent changes in administrative price mechanism.

However, Ball and Mankiw (1992) point out that relative price shocks render the distribution of price changes skewed. If the price setters want to revise their prices in response to a one time cross sectional price change, they have to face a menu cost associated with price revision. Hence, price setters tend to revise the price only when the price rise is too large. If the distribution of the shock is skewed then mean price level will change temporarily. In that case the weighted average of cross-sectional price changes losses its robustness and also *ad hoc* exclusion of commodities will not circumvent the problem of skewness.

Rather than excluding commodities that are prone to this kind of shocks, a better way of resolving this problem is to trim the tails of the skewed distribution or consider a weighted median instead of weighted mean. As noted by Bryan and Cecchetti (1993), the advantage of this trimming method is that it derives a robust measure of change in the mean price level and eliminates the measurement error induced bias in price changes. Also, it prevents the loss of information brought in by the incoming data.

Trimmed mean method

A trimmed mean based measure proposed by Bryan and Cecchetti (1993) and Bryan, Cecchetti and Wiggins (1997) involves the following steps in calculating core inflation. First, arrange the sample with associated weights in order. Second, construct the cumulative sum of weight 1 to i: $W_i = \Sigma w_i$. Third, decide the set of observation to be averaged for calculation as:

$$I_\alpha = \frac{\alpha}{100} < W_i < \left(1 - \frac{\alpha}{100}\right)$$

Fourth, the trimmed mean is calculated as:

$$\overline{\pi}_\alpha = \frac{1}{1 - 2(\frac{\alpha}{100})} \sum_{i \in I_\alpha} w_i \pi_i$$

where π_i is the annual inflation rate for i^{th} commodity.

Weighted Median Method

Bryan and Pike (1991) proposed an alternative measure of core inflation by calculating the median price changes instead of the weighted mean. They argue that all prices reflect relative price changes to some extent. The exclusion method is somewhat subjective as it assumes only those commodity prices are affected by some special factors. The extraction of core measure is based on identifying the persistent relative price changes of all the commodities and not that of certain commodities alone. Two methods have been used to construct weighted median core inflation. First, following Smith (2004), we define weighted median measure as:

$$\overline{\pi}_{\alpha} = \frac{1}{N-2m}\sum_{i=m}^{N-m}\pi^{i}_{t-k,t}$$

where $\pi^{i}_{t-k,t} = \ln\left(\frac{p^{i}_{t}}{P^{i}_{t-k}}\right)$, i indicates the commodity and for annualized monthly inflation rate $k = 12$. N is the number of commodities in the basket. If N is an odd number then $m = N\alpha$ *where* m is the largest integer less than or equal to $N\alpha$ for even number $m = N\alpha - 1$. This study has N =76, with $\alpha = 0.50$ the $m = 37$. By averaging the inflation rate at the median we get weighted median core inflation. Second, as an extension to it, as suggested by Bryan and Cecchetti (1993) we can construct the 'frequency the good is at median' by simply counting the number of times a good is at the median for the entire sample period. Instead of the weights, we can use this frequency as weights of the commodities to construct the weighted median measure[1]. The measures π^{W1} and π^{W2} have been computed from the respective methods.

The common trends approach

Under this approach, we try to capture the permanent component of headline inflation and call it as core inflation using the framework of Blanchard and Quah (1989) and Quah and Vahey (1995). Blanchard and Quah have decomposed permanent and transitory components non-cointegrated I (1) variables by imposing restrictions that variables do not respond in the long run to certain shocks. Using such empirical framework, Quah and Vahey measured core inflation as the permanent component of headline inflation. However, they assumed there is no cointegration among the variables. Nevertheless, Bagliano and Morana (2003) used the long run cointegrating properties of the variable to

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[1] With 76 commodities we get two frequency series. We constructed the weighted median measure using both the measure and selected the less variance measure for the evaluation purposes.

identify the permanent component of inflation[2]. In the present paper, we identify important determinants of inflation to construct core inflation, which can be defined as conditional forecast of inflation.

The methodology

Let $\boldsymbol{x_t}$ be a vector of $\boldsymbol{I\ (1)}$ variables. If there is $\boldsymbol{0 < r < n}$ cointegrating relationships then the cointegrated VAR system can be written as:

$$\Delta x_t = \Pi(L)\Delta x_{t-1} + \alpha\beta' x_{t-1} + \varepsilon_t \tag{1}$$

Where $\boldsymbol{L}$ is lag operator; $\boldsymbol{\Delta = 1 - L}$; $\boldsymbol{\Pi(L) = \Pi_1 + \Pi_2 L + \ldots + \Pi_p L^{p-1}}$; β is $\boldsymbol{n \times r}$ matrix and cointegration among variables indicates that $\boldsymbol{\beta' x_t}$ are stationary. α is corresponding $\boldsymbol{n \times r}$ matrix of factor loadings and $\boldsymbol{\varepsilon_t}$ is serially uncorrelated reduced form disturbances. The Wold representation expression in (1) can be written as:

$$\Delta x_t = C(L)\varepsilon_t \tag{2}$$

Further, by recursive substitution the expression in (2) can be define for levels of variables:

$$x_t = x_0 + C(1)\sum_{j=0}^{t-1}\varepsilon_{t-j} + C^*(L)\varepsilon_t \tag{3}$$

where $\boldsymbol{C\ (1)}$ captures the long run impact of $\boldsymbol{\varepsilon_t}$ on $\boldsymbol{x}$ and $\boldsymbol{x_0}$ is initial observation in the series. However, $\boldsymbol{\varepsilon_t}$ are reduced form disturbances; hence, we need to transform them into structural innovations. The structural shocks for $\boldsymbol{x_t}$ is:

$$\Delta x_t = \Gamma(L)\varphi_t \tag{4}$$

where $\boldsymbol{\varphi_t = (\psi_t,\ \nu_t)}$ in which $\boldsymbol{\psi_t}$ and $\boldsymbol{\nu_t}$ are subvectors having permanent and transitory impact on x_t respectively and $\boldsymbol{\Gamma(L) = \Gamma_0 + \Gamma_1 L + \ldots}$ Note that first term of $\boldsymbol{C\ (L)}$ in (2) is **I** and that of $\boldsymbol{\Gamma(L)}$ in (4) is $\boldsymbol{\Gamma_0}$. By equating these two we can define the reduced form errors in terms of structural innovations:

$$\varepsilon_t = \Gamma_0\varphi_t \tag{5}$$

The expression in (2) and (4) shows that

$$C(L)\Gamma_0 = \Gamma(L)$$

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[2] To understand the development this literature, see Stock and Watson (1988), Gonzalo and Granger (1995), King et al (1991), Mellander et al (1992) and Warne (1993).

Hence, ***C (1)*** $\Gamma_0 = \Gamma_1$ being the long run impact matrix. To identify that some shocks have permanent and some have transitory impact, we must impose the following restriction:

$$\Gamma(1) = (\Gamma_g \, 0) \tag{6}$$

From (4) the structural form for levels of *x* can be written as:

$$x_t = x_0 + \Gamma(1)\sum_{j=0}^{t-1}\varphi_{t-j} + \Gamma^*(L)\varphi = x_0 + \Gamma_g \sum_{j=0}^{t-1}\psi_{t-j} + \Gamma^*(L)\varphi \tag{7}$$

The permanent component in (7) can be expressed as a random walk process:

$$\tau_t = \tau_{t-1} + \psi_t \tag{8}$$

Using expressions in (8) and (7) the common trend representation for $\boldsymbol{x_t}$ can be defined as:

$$x_t = x_0 + \Gamma_g \tau_t + \Gamma^*(L)\varphi_t \tag{9}$$

Further, the long run forecast of x_t conditional upon the permanent shocks can be expressed as:

$$\lim_{h\to\infty} E_t x_{t+h} = x_0 + \Gamma_g \tau_t \tag{10}$$

where Γ_g is the long run impact matrix. The identification requires further restriction on the Γ_g that the impact of some shocks dies out in the long run.

3. Estimates of core inflation

We construct core inflation using monthly data for the period April 1994 to March 2005. The choice of the sample period is dictated by the availability of consistent time series on prices index with the latest base period 1993-94. The new measure of wholesale price index (WPI) comprises 435 commodities under three different categories: Primary articles; Fuel, Power, Light and Lubricants; and Manufactured Products with weights 22.025, 14.226 and 63.749 percent respectively. For computing core inflation, 435 commodities are classified into 76 items.

Following different approaches, seven alternative measures of core inflation are constructed. We obtain two core inflation measures from exclusion method; two from trimmed mean method; one from weighted median (π^{W1}) method of Smith (2004); one from weighted median method (π^{W2}) of Bryan and Cecchetti (1993); and one from common trends methods. As the discussion on exclusion based measures are already available in the literature, we consider only two measures of core inflation by excluding food alone (π^F) and food and energy items (π^{FE}). The commodities under these categories are considered for exclusion since their prices are largely supply driven and relatively more volatile[3]. Among these two groups of commodities, energy prices remained consistently on higher side during the sample period. The food articles have 15.402 and energy items have 14.226 weights in the WPI basket.

For estimating common trend core inflation, we construct a four variable VAR system including seasonally adjusted log of industrial production (y), annual point-to-point percentage change in WPI (π), annual point-to-point percentage change in M3 (m), and log of oil prices (o)[4]. The oil price is measured as US $ per barrel of crude and used as an endogenous variable so that we can capture the response of all the variables in the system to supply shocks.

The standard unit-root tests (not reported) showed that all the variables are integrated of order one. The cointegration test is conducted using the Johansen's (1991) maximum likelihood method. We have used two lags in short-run specification of the model as suggested by Akaike information, Schwartz, Hannen-Quin criteria and likelihood ratio test. The results of cointegration tests are presented in Table 1. The trace and max-eigen value statistics suggest existence of one cointegrating vector at 1 % level of significance. The cointegrating vector normalized with respect to m show that the long run cointegrating coefficients with respect to output and oil price are statistically insignificant. The likelihood ratio test cannot reject the restriction that the cointegrating vector {0,0,1, -1}. This implies m-π is a constant in the long run.

[3] Some studies have constructed core measures by excluding commodities whose prices are administered. We do not consider such measures since administrative price mechanism has been dismantled to a larger extent during the sample period. Moreover, administrative mechanism covers larger set of commodities having around 50 percent weight in WPI. Hence, excluding them might provide a core measure, which could be an inadequate representative of underlying inflation.

[4] Arguably oil price in the foreign market may not be a proxy for supply shock since it is administered. Nevertheless, it is not appropriate to assume that impact of oil price fluctuations on domestic prices is nil in the short to medium run, particularly when there is a persistent rise in oil prices in the international market during the sample period.

Table 1: Cointegration results

Hypothesis	Eigenvalue	λ_{TRACE}	95 % Critical Value	λ_{MAX}	95 % Critical Value
$r = 0$	0.266	75.63*	53.12	39.98*	28.14
$r \leq 1$	0.151	35.65	34.91	21.13	22.00
$r \leq 2$	0.075	14.51	19.96	10.14	15.67
$r \leq 3$	0.033	4.37	9.24	4.37	9.24

The cointegrating coefficients

y	o	m	π
-0.0017	0.0728	1	-0.865
(0.076)	(0.037)		(0.321)

The restricted cointegrating coefficients

y	o	m	π
0	0	1	-1
χ^2 (p value)		5.835 (0.119)	

Standard errors are in the parentheses; r is valid cointegration vectors; and * denote significance at 1% level.

With one cointegrating relationship among four variables there could be three distinct shocks having permanent effects on some of the variables under consideration. We define the impact of autonomous shocks on the variables as follows: (i) a foreign real shock (ψ_f); (ii) a domestic real shock (ψ_r); and (iii) a nominal shock (ψ_n). The nominal shock is assumed to have no long run impact on output whereas it has equal effect on money growth and inflation. The domestic real shock and nominal shock do not have long-run effect on oil prices.

The common trends representation for four variables can be defined as

$$\begin{pmatrix} o \\ y \\ m \\ \pi \end{pmatrix}_t = \begin{pmatrix} o \\ y \\ m \\ \pi \end{pmatrix}_0 + \begin{pmatrix} \kappa_{11} & 0 & 0 \\ \kappa_{21} & \kappa_{22} & 0 \\ \kappa_{31} & \kappa_{32} & \kappa_{33} \\ \kappa_{41} & \kappa_{42} & \kappa_{43} \end{pmatrix} \begin{pmatrix} \tau_f \\ \tau_r \\ \tau_n \end{pmatrix}_t + \Phi(L) \begin{pmatrix} \psi_f \\ \psi_r \\ \psi_n \\ \upsilon_1 \end{pmatrix}_t \tag{11}$$

The common trend estimate of core inflation is:

$$\pi_t^{CT} = \pi_0 + \hat{\kappa}_{41}\hat{\tau}_{f\,t} + \hat{\kappa}_{42}\hat{\tau}_{rt} + \hat{\kappa}_{43}\hat{\tau}_{n\,t}.$$

The estimated Γ_g matrix along with the long run forecast error variance decomposition is reported in Table 2.

Table 2: Estimates of common trends model

Variables	Shocks		
	ψ_f	ψ_r	ψ_n
	Long-run impact matrix Γ_g		
o	0.0866 (0.010) *	0	0
y	0.0007 (0.002)	0.0098 (0.002) *	0
m	0.0003 (0.001)	-0.0001 (0.002)	0.0069 (0.001)*
π	0.0003 (0.001)	-0.0001(0.002)	0.0069 (0.001)*
	Long-run forecast error variance decomposition		
o	1.00	0	0
y	0.005 (0.039)	0.995 (0.028)	0
m	0.001 (0.012)	0.000 (0.004)	0.999 (0.013)
π	0.001 (0.012)	0.000 (0.004)	0.999 (0.013)

Asymptotic standard errors are in the parentheses. * denote significance at 1% level.

The estimates show that the foreign real disturbances have significant impact on oil prices. The domestic real shock has significant positive impact on output and no significant impact on money stock and inflation. The nominal shock has a significant positive impact on both money stock and inflation. Further, we can infer the long run association among variables from the forecast error variance decomposition. The results in Table 2 show that 99.9 % of variation in both *m* and π can be attributed to domestic nominal shocks. The variation in industrial output due to foreign real shock seems to be negligible.

4. Evaluation of core inflation

There are several criteria proposed in the literature to examine the relative merits of core inflation. First, we examine the unbiasedness, volatility and how well core measures are associated with policy variables such as money stock. Accordingly, we present mean, standard deviation and the correlation between core inflation and annual point-to-point percentage change in M3 money in Table 3. One of the basic criteria that the core inflation must satisfy is that it is unbiased to headline inflation. This implies that in the long run the difference between average of headline and core inflation must be zero. In this sense, the mean of core inflation measures obtained from common trend (π^{CT}) approach and by excluding food articles (π^{F}) are much closer to mean of headline inflation. The mean of trimmed mean and weighted median based core inflation measures (π^{W1}) lie far off from the mean of actual inflation. Another important criterion is that core inflation should be noise free; hence, it is less volatile. By this criterion, we find trimmed mean measures and core inflation from common trend approach qualify as they have relatively less volatile.

Table 3: Descriptive statistics of inflation

Inflation	Mean	σ	ρ	Inflation	Mean	σ	ρ
π	5.90	2.98	0.35	π^{T20}	5.10	2.60	0.35
π^{F}	5.92	3.15	0.22	π^{W1}	4.26	2.62	0.28
π^{FE}	4.93	3.74	0.28	π^{W2}	5.37	3.00	0.37
π^{T15}	5.30	2.53	0.37	π^{CT}	5.88	2.17	0.73

σ is standard deviation and ρ is correlation between inflation and point-to-point annual percentage change in M3 money.

The correlation coefficient indicates that the exclusion and weighted median based core measures (π^{W1}) have relatively weak correlation with nominal money growth while core inflation estimated through common trends model has strong correlation. In sum, the statistics in Table 3 exhibit that core inflation from common trends is relatively a better measure as it consistently satisfies all the basic criteria under consideration.

Further, we examine whether the estimated core measures satisfy certain additional empirical criteria. Freeman (1998) argued that there must be cointegration between core and headline inflation. Marques, Neves and Sarmento (2000) proposed certain additional criteria to identify core inflation as an indicator of permanent trend component in headline inflation: (i) the targeted and core inflation must be cointegrated with cointegrating parameter of unity; (ii) core inflation must be an attractor of targeted inflation; and (iii) targeted inflation must not be an attractor of core inflation.

However, Ribba (2002) argued that any measure of core inflation should satisfy two conditions: (i) core (π^C) and headline inflation (π) are cointegrated with a cointegrating vector (1, -1); and (ii) there exists an error correction representation:

$$\left.\begin{aligned} \Delta\pi_t^c &= b_{11}(L)\Delta\pi_{t-1}^c + b_{12}(L)\Delta\pi_{t-1} + \varepsilon_{\pi ct} \\ \Delta\pi_t &= b_{11}(L)\Delta\pi_{t-1}^c + b_{12}(L)\Delta\pi_{t-1} - \alpha(\pi_{t-1}^c - \pi_{t-1}) + \varepsilon_{\pi t} \end{aligned}\right\} \quad (12)$$

where L is lag operator; $\Delta = 1 - L$; and $\varepsilon_t = (\varepsilon_{\pi ct}, \varepsilon_{\pi t})'$ such that $E(\varepsilon_t) = 0$ and $E(\varepsilon_t \varepsilon_t') = \Sigma_\varepsilon$. Hence, condition (ii) implies that shocks in core inflation can influence the long run forecast of headline inflation and not *vice versa*. In other words, there is one-way causality from core to headline inflation at zero frequency[5]. If so,

$$\lim_{h\to\infty} \partial E(\pi_{t+h}) / \partial\varepsilon_{\pi ct} \neq 0; \quad \text{and} \quad \lim_{h\to\infty} \partial E(\pi_{t+h}) / \partial\varepsilon_{\pi t} = 0 \quad (13)$$

The above conditions imply that the conditional forecast of headline inflation h period ahead depends only on core inflation. We investigate the validity of these conditions using Johansen's (1991) maximum likelihood approach. The results of ADF and PP tests presented in Table 4 indicate that headline and estimated core measures have unit root. The Akaike information, Schwarz, and Hannen-Quinn criteria suggested two lags in the short run specification of the VAR model.

[5] Marques et al. (2002) argued that π^C is an appropriate measure of core inflation if π^C and π are cointegrated and π does not Granger cause π^C at all frequencies. Nevertheless, Ribba (2002) has shown that this not a necessary condition to identify π^C as an indicator of the permanent trend in π.

Table 4: Stationarity properties of headline & core inflation

Inflation	ADF Statistics	Phillips-Perron Statistics	Inflation	ADF Statistics	Phillips-Perron Statistics
π	-1.193	-1.115	π^{T20}	-1.337	-0.890
π^{F}	-1.344	-1.155	π^{W1}	-2.135	-1.115
π^{FE}	-1.371	-1.386	π^{W2}	-0.764	-0.968
π^{T15}	-0.735	-0.883	π^{CT}	-1.398	-1.380

Table 5: Cointegration between headline and core inflation

Variables	Hypothesis	Eigenvalue	λ_{TRACE}	λ_{MAX}
π^{F}, π	r = 0	0.064	12.294	8.276
	r ≤ 1	0.032	4.018	4.018
π^{FE}, π	r = 0	0.061	10.134	7.803
	r ≤ 1	0.019	2.331	2.331
π^{T15}, π	r = 0	0.124	20.943**	16.485**
	r ≤ 1	0.035	4.458	4.458
π^{T20}, π	r = 0	0.094	17.506**	12.278**
	r ≤ 1	0.041	5.227	5.227
π^{W1}, π	r = 0	0.091	14.499**	11.914**
	r ≤ 1	0.020	2.585	2.585
π^{W2}, π	r = 0	0.059	10.996	7.552
	r ≤ 1	0.027	3.443	3.443
π^{CT}, π	r = 0	0.095	14.067**	12.206**
	r ≤ 1	0.014	1.806	1.806

* and ** indicate significance level at 1% and 5% respectively.

The λ_{TRACE} and λ_{MAX} statistics in Table 5 show that core measures derived from trimmed mean, one measure of weighted median (π^{W1}) and common trends methods are cointegrated with headline inflation. Mohanty, Rath and Ramaiah (2000) and Joshi and Rajpathak (2004) also find that core inflation from trimmed mean is cointegrated with headline inflation and the latter study provides additional empirical support for trimmed mean measure in terms of its ability to predict headline inflation. Hence, headline inflation does not diverge permanently from these core measures. However, exclusion and another weighted median based measure are not cointegrated with headline inflation. This is in consonance with the evidence of Mohanty, Rath and Ramaiah (2000). In contrast to this, Joshi and Rajpathak (2004) find several exclusion-based measures of core inflation cointegrated with headline inflation, but none of them performed well in predicting headline inflation.

The corresponding cointegrating vectors with loading coefficients and the χ^2 statistics for testing the joint restriction that core inflation is cointegrated with headline inflation with cointegrating vector (1, -1) and that the loading coefficient in core equation is zero are presented in Tables 6. The column (1) and (2) presents the unrestricted cointegrating coefficients and the corresponding loading coefficients. The unrestricted cointegrating coefficient with respect to the core measure derived from common trends method is almost –1 while that with respect to trimmed mean based core inflation is not closer to – 1. The loading coefficients only in headline inflation equation are statistically significant, indicating that there is a tendency of actual inflation adjusting to core inflation in the long run and not *vice versa*. The sign of the significant loading coefficients ensures convergence; hence, there is stability in the system.

Table 6: Results of cointegration space

Variables	Normalized cointegrating vector		χ^2 statistics
π^{T15}, π	1.00	-0.910 (0.024)*	6.441 [0.03]
Loading coefficients	-0.003 (0.110)	0.337 (0.136)*	
π^{T20}, π	1.00	-0.840 (0.029)*	7.443 [0.02]
Loading coefficients	-0.175 (0.103)	0.036 (0.133)	
π^{W1}, π	1.00	-0.594 (0.062)*	10.158 [0.006]
Loading coefficients	-0.166 (0.051)	-0.088 (0.062)	
π^{CT}, π	1.00	-1.135 (0.024)*	3.729 [0.15]
Loading coefficients	0.065 (0.037)	0.161 (0.048)*	

Figures in (#) and [#] are standard errors and p – values respectively; * and ** indicate significance level at 1% and 5% respectively; and χ^2 statistics is to test the joint restriction that cointegrating vector has the form (1, -1) and that π does not cause π^C (a particular measure of core inflation) at zero frequency.

However, the χ^2 statistics presented in the last column of the Table 6 exhibit that the joint restriction is rejected at conventional level of significance in the case of trimmed mean measures whereas the restriction is accepted in the model based core measure. This is the striking feature of the evidence from error correction model that the core measure derived from common trends approach is a powerful predictor of headline inflation. Putting the evidence based on various evaluation criteria, we find that the model based core inflation emerge as a distinct measure since it possesses all the desirable properties: unbiased to headline inflation, less volatile, relatively high correlation with growth of nominal money, cointegrated with headline inflation, and powerful attractor of headline inflation.

5. Conclusion

We have constructed core inflation using exclusion method, limited influentual methods and a common trends model over the period April 1994 to March 2005. The estimated core measures are subjected to empirical evaluation of how well they qualify certain desirable properties of core inflation. We have considered unbiasedness, less variability, close association with policy variables, cointegration with headline inflation and good attractor of headline inflation as empirical criteria for judging the estimated core measures for their usefulness.

The core measures estimated from exclusion and limited influential methods do not seem to be convincing as they fail to qualify consistently all the criteria. Nevertheless, the core inflation series derived from a common trends model have consistently passed all the empirical tests under consideration. However, it is often argued that estimation of common trends model with new observations tend to change the core inflation in the past; hence, adding difficulties for the authorities in using such model based core measures as part of communication strategy. Nevertheless, evidence suggests that overemphasizing exclusion based core inflation for its simplicity should not rule out the potential use of model based and economically interpretable core inflation for policy purpose.

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