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MONETARY POLICY AND FOOD PRICES IN INDIA

Jeevan Kumar Khundrakpam and Dipika Das[•]

Abstract

Using a vector error correction model (VECM), the paper examines the relative response of food and manufactured prices to change in interest rate and money supply in India during the period 2001:Q1 to 2010:Q2. It finds that, in the long-run, while money supply leads to rise in the prices of both food and manufactured prices, hike in call rate has a negative effect only on manufactured prices. The impact of money supply is, however, more on food prices than on manufactured prices. There is no evidence of long-run neutrality of money, as increase in money supply leads to less than proportionate change in price. In the short-run, there is overshooting in the prices of both food and manufactured products from their respective long-run equilibrium following monetary shocks. The overshooting, however, is more in food prices than in manufactured prices. Further, in the short-run, call rate has a significant impact only on manufactured prices, while the impact of money supply is significant only on food prices. Both food and manufactured prices increase induces call rate hike. But money supply increases with rise in food prices and decreases with the rise in manufactured prices.

Keywords: food prices, manufactured prices, money supply, call rate

JEL classification: E51, P22

Introduction

In developing country context, inflation tolerance in India is fairly low. And within the overall inflation, food price inflation is least tolerated as bulk of the population spend majority of their income on food items. During the last decade, food price inflation exceeded the headline inflation measured by wholesale price

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index since around the end of 2005, barring the period September 2007 to September 2008. This gap has become all the more glaring in the more recent time. Currently, this continuing rise in food price inflation has become a major cause for concern for policy makers in India.

Traditionally, fluctuation in food prices is analysed through supply and demand gap. While this traditional approach still remains extremely important, recent literature has also given a great deal of focus on the impact of macroeconomic variables, especially monetary and financial factors, on agricultural prices [for example, Orden and Fackler (1989); Saghaian, Reed and Marchant (2002); Peng, Marchant and Reed (2004); and Asfaha and Jooste (2007)]. The basic findings of these studies are:

i) The responses of agricultural prices to a monetary shock are much faster than the corresponding responses of non-agricultural prices;

ii) Monetary shocks can lead to overshooting of agricultural prices from their long-run equilibrium in the short-run; and

iii) In economies with less developed financial markets, money supply as a monetary policy instrument has a much stronger impact on agricultural prices than interest rates.

The objective of this piece is to conduct a similar analysis in the Indian context by drawing on the literature. Specifically, it attempts to answer the following questions:

i) Do monetary policy instruments have any differential impact on food and non-food prices in India?

ii) Which of the monetary policy instrument viz., money supply and interest rate, has more impact on food and non-food prices?

The rest of the paper is organised as in the following. Section I briefly reviews some of the relevant theoretical and empirical literature. In section II, data source and some of their stylised facts are presented. The analytical framework and the empirical results are presented in section III. Section IV summarises with concluding remarks.

Section I: Theoretical and Empirical Literature

The importance of macroeconomic and financial factors in determining agricultural commodity prices was first pointed out by Schuh (1974). In the context of monetary policy, the main issue has been whether agricultural and non-agricultural prices respond proportionately to monetary policy or not. It is theoretically argued that, as agricultural prices are more flexible, they respond faster to change in money supply than non-agricultural prices, which are relatively inflexible. Agricultural prices are more flexible, because, agricultural commodities tend to be more standardised and exhibit lower transaction cost than manufactured goods. Thus, agriculture prices are characterised by short-term contracts and respond more quickly to monetary changes than the prices of other goods (Bordo, 1980). Even in the traditional explanation through supply and demand imbalances also, as agricultural production takes a much longer time, changes in demand will, in the short-term, get reflected more in price changes than change in the volume of production.

Other studies have addressed the issue in the context of broader macroeconomic environment. These studies, drawing on Dornbusch's (1976) overshooting model of exchange rate determination, establish the linkages among exchange rates, money, interest rates and commodity prices. Frankel (1986) making distinction between "fixed price" sectors such as industrial and service sector and a “flex-price” sector such as agriculture uses Dornbusch’s overshooting model and showed that monetary changes can cause agricultural prices to overshoot. In this model, decline in nominal money supply imply decline in real money supply, which by increasing interest rate depresses the flexible price of agricultural commodities. Overshooting takes place in order to generate expectation of a future appreciation sufficient to offset the higher interest rate.

The disproportionate response of agricultural and industrial prices to monetary changes, both in the short-run and long-run has been empirically validated by a number of studies for different countries. For the U.S., Orden and Fackler (1989) using a VAR and impulse response function show that an increase in money supply raises agricultural prices relative to the general price level for more than a year, implying monetary changes lead to change in real agricultural prices in both the short- and long-run. Saghaian, Reed and Merchant (2002) extended the Frankel’s closed economy model to an open economy framework by including exchange rate to account for international trading of agricultural commodities for the US economy. They found that monetary changes have both short- and long-run effects on agricultural prices.

Several other studies have also found similar evidence in many emerging and transition economies. In Hungary (Bakucs and Ferto, 2005), Slovenia (Bakucs, Bojnec and Ferto, 2006) and South Africa (Asfaha and Jooste, 2007), it was found that monetary changes not only have real short- and long-run effects on agricultural prices, but also that agricultural prices adjust faster than industrial prices to innovations in money supply. In Korea, the Philippines and Thailand also, it was found that overshooting for agricultural prices is larger than for manufactured prices and money supply affect real variables and relative prices either through overshooting or non-neutrality of money (Saghaian, Hasan and

Reed, 2002). Similarly, in Pakistan it was found that agricultural prices adjusted faster than industrial prices in the long-run due to short-run changes in money supply and exchange rate (Hye and Siddiqui, 2010).

In the above studies, though money supply had long-run impact on agricultural prices, they reject the hypothesis of long-run neutrality of money. Some studies, however, find long-run neutrality of money, but still support the evidence of overshooting in agricultural prices due to monetary changes in the short-run (for example, Cho *et al.*, 2004).

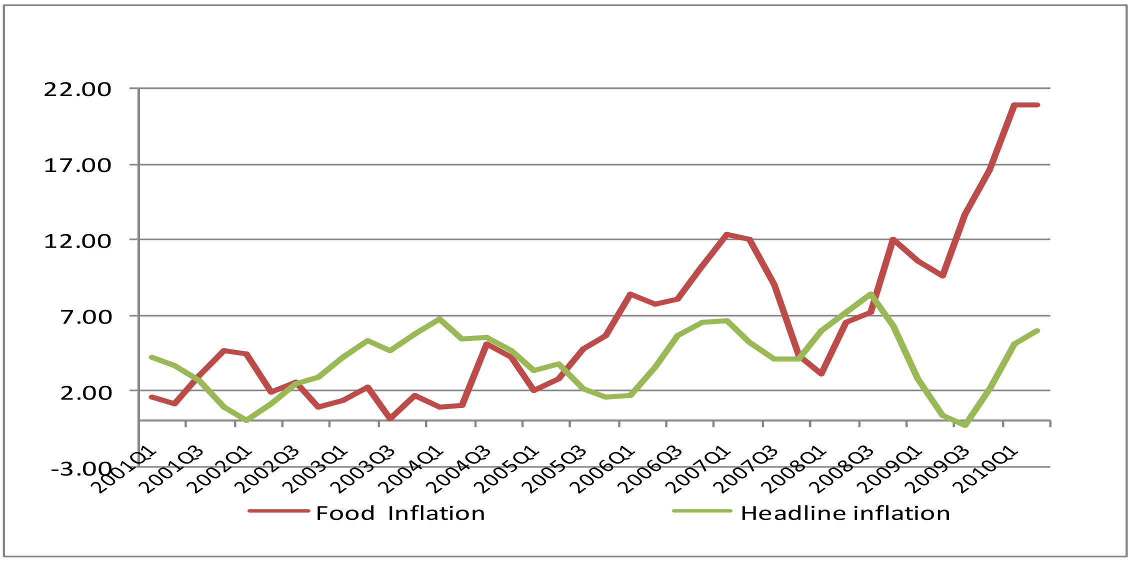
Interestingly, in the case of China, Peng, Marchant and Reed (2004) distinguish the impact of two monetary policy instruments viz., money supply and interest rate, on food prices. They find that monetary impacts on food prices in China emanates mostly from money supply, while interest rates play a very limited role due to controlled interest rate regime and underdeveloped nature of financial markets in China. With money supply being non-neutral in determining food prices in China, they conclude that the dominant monetary policy instrument which can be used to control food prices in China is the money supply instead of interest rates.

Section II: Data and Some Stylised Facts

All the relevant data are culled out from Handbook of Statistics on Indian Economy, Reserve Bank of India for the period 2001:Q1 to 2010:Q2. The variables are food prices (FP), manufactured prices (MP), exchange rate represented by nominal effective exchange (EX), weighted average call rate as the proxy for policy rate (CALL), broad money (M3) and narrow money (M1).

First, we briefly set out some of the stylised facts in increase in food prices and headline inflation measured by WPI (Chart-1). The year-on-year increase in headline and food inflation show that food inflation which mostly remained lower than headline inflation during first half of the 2000s, began to increasingly exceed headline inflation by 2005:Q4, barring the period 2007:Q4 to 2008:Q3.

Chart 1: Food and Headline Inflation (in per cent)



Second, the descriptive statistics of data measured by the coefficients of variation show that barring exchange rate and to some extent manufactured prices, other variables are much more volatile. The monetary policy variables viz., call rate and money supply variables, both narrow and broad money, are particularly volatile (Table-1). Fluctuation in food price is much more than that of manufactured prices providing some sort of evidence to the theoretical argument that food prices are relatively more flexible in nature as compared to manufactured prices.

Table-1: Descriptive Statistics of Variables

Variable	Mean	Standard Deviation	Minimum	Maximum	Coefficient of Variation
Food Prices	112.2	22.4	88.3	172.1	20.0
Manufactured Prices	102.8	13.2	84.3	127.3	12.8
Exchange Rate	88.0	3.67	80.0	96.2	4.2
Call Rate	6.1	1.91	3.2	10.4	31.3
Broad Money	2754002	1369654	1124174	5677076	49.7
Narrow Money	791438	378530	341796	1580102	47.8

Section III: Analytical Framework and Empirical Estimates

We draw on the theoretical and empirical framework provided in Saghaian, Reed and Marchant (2002), which is an extension of Dornbusch's overshooting model by including a third sector viz., agricultural prices. There are four variables in their model viz., money supply, exchange rate, agricultural and non-agricultural prices.[1] The empirical validation of the model is carried out through cointegration analysis by estimating long-run relationship between the variables and the short-run dynamics employing vector error correction model. This is carried out in the following steps. First, unit root tests are conducted to assess the stationarity of each variable. Second, co-integration tests are performed and the long-run relationships estimated to check whether the long-run neutrality of money on prices holds or not i.e., whether the coefficient of money supply in the cointegrating equation of price is equal to one or not. Because, neutrality of money on price imply that change in money supply will only have an equivalent nominal impact but no real impact. Third, the short-run dynamics are analysed by estimating a vector error correction model (VECM) to check for the presence of

[1] For detail description of the theoretical framework, please refer to Saghaian, Reed and Marchant (2002).

overshooting phenomenon. When the coefficient of the error correction term (ECT) of the relevant variable in the VECM is negative and statistically significant, it provides the evidence of overshooting. Because, the negative sign of the ECT implies reduction in the value of the variable over the horizon to return to its long-run equilibrium.

In our context, following Peng, Marchant and Reed (2004), we add another monetary policy variable viz., weighted call rate as proxy for policy rate, to the above model. By including two monetary policy variables, one representing rate variable and the other quantum variable, we attempt to compare the relative effectiveness of this two policy variables on the food and manufactured prices. Further, we also perform Granger Causality Test among the variables in the short-run by examining the joint significance of the lagged first difference of other variables in the VECM. This enables us to determine the direction of causality among the variables in the short-run. For the estimation purposes, all the variables were converted to natural logarithm by adding the prefix ‘L’ to the variable names.

Unit Root Tests

The augmented Dicky-Fuller (ADF) and Phillip-Perron (PP) tests results reported in table-2 show that all the variables, barring narrow money (M1), are stationary at 1 per cent critical level. In the case of M1, while ADF tests show it as non-stationary even after first differentiation, PP tests, in complete contrast, show the series to be stationary at 1 per cent level of significance. As a result of this inconclusive unit root properties of M1, we exclude it from our model as an alternative money supply variable.

Table 2: Unit Root Tests

Variable (X)	ADF		PP	
	Log X	∆Log X	Log X	∆Log X
LFP	1.68	-7.21(t)*	4.94	-8.12(t)*
LMP	-3.49(t)	-5.58*	-2.24(t)	-4.41*
LEX	-2.97	-7.96*	-2.91	-8.24*
LCALL	-2.89	-5.52*	-2.57	-5.34*
LM_3	-1.80(t)	-10.2*	-1.47(t)	-10.3*
LM_1	-2.52(t)	-2.71	-4.30(t)*	-12.6*

Notes: * denote significance at 5% and 1% level, respectively. The lag length in the ADF tests was chosen based on Schwarz Bayesian Criterion (SBC) with maximum lag set at 4, being quarterly data. 't' in the parentheses indicate inclusion of a trend component in the estimates, which was based on its statistical significance in the equation.

Co-integration Tests

Given that all the variables are I*(1)*, we performed Johansen's co-integration tests.[2] First the VECM lag length was selected. All the five alternative tests suggest the appropriate leg length to be unambiguously four (Table-3).

Table-3: VAR Lag Order Selection Criteria

Lag	LogL	LR	FPE	AIC	SC	HQ
0	239.8008	NA	5.01e-12	-11.83162	-11.18520	-11.60163
1	436.5777	310.7005	6.12e-16	-20.87251	-19.14874	-20.25921
2	475.0821	50.66363	3.36e-16	-21.58327	-18.78213	-20.58665
3	521.3137	48.66489	1.44e-16	-22.70072	-18.82223	-21.32079
4	575.8283	43.03783*	5.27e-17*	-24.25412*	-19.29827*	-22.49087*

* indicates lag order selected by the criterion
LR = sequential modified LR test statistic (each test at 5% level); FPE = Final predicition error; AIC = Akaike information criterion; SC = Schwarz information criterion; HQ = Hannan-Quinn information criterion.

The number of cointegrating vectors depends upon the model specification in terms of inclusion/exclusion of intercept and trend component. It was found that both the trace and max-Eigen statistics in four out of the five

[2] Two dummies, DUM1 and DUM2 were included in the tests to account for extreme volatility in exchange rate. DUM1 takes the value of 1 for 2007:Q2 to Q4 when there was sudden sharp appreciation in the exchange rate due to massive capital inflows and zero otherwise. DUM2 takes the value of 1 to account for sudden dips in exchange rate in 2006:Q3, 2008:Q1 and 2009:Q3 and zero otherwise.

models show three cointegrating vectors at least at 5 per cent significance level (Table-4).

Table 4: Number of Cointegrating Relations by Model

Data Trend:	None	None	Linear	Linear	Quadratic
Test Type	No Intercept No Trend	Intercept No Trend	Intercept No Trend	Intercept Trend	Intercept Trend
Trace	3	5	3	3	3
Max-Eig	3	5	3	3	3

Selected critical values at 0.05 level based on MacKinnon_Haug_Michelis (1999).

Among these five models, we selected the model with unrestricted intercept and no trends, which imply intercept in both the cointegrating space and in the short-run model, but no trend in either of them. This model more often than not represents the relationship among macroeconomic variables. The trace and max-Eigen statistics tests are presented in Annex-1.

Vector Error Correction Model (VECM)[3]

The normalised three cointegrating vectors representing the long-run relationships among the variables are presented in Table-5.

Table-5: Parameters in Normalised Cointegrating Vectors

Cointegrating Equations	Equation1	Equation2	Equation3
$LFP_{(-1)}$	1.0	0.0	0.0
$LMP_{(-1)}$	0.0	1.0	0.0
$LEX_{(-1)}$	0.0	0.0	1.0
$LCALL_{(-1)}$	0.023	0.162	0.015
	(0.33)	(3.31)**	(0.79)
LM3	-0.322	-0.184	0.04
	(-6.33)*	(-5.20)*	(2.92)**
Constant	0.003	-2.20	-5.09

Note: Figures in parentheses are t-statistics. * and ** denote significance at 1% and 5% level, respectively.

[3] In the VECM, we do not make a distinction on whether money supply process is endogenous or exogenous. We focus, as typically done in the literature using VAR or VECM, on the impact of a monetary shock on the two prices, irrespective of the money supply process.

The three long-run relationships are: i) impact of call rate and money supply on food prices; ii) impact of call rate and money supply on manufactured prices; and iii) impact of call rate and money supply on exchange rate. The signs of the coefficients in all the three equations are as per *a priori* expectations, except the coefficient of call rate in exchange rate equation which, however, cannot be statistically distinguished from zero. The coefficients of money supply in the price equations are negative indicating that expansionary monetary policy leads to increase in the prices of both food and manufactured products. However, the value of the coefficient is larger for food prices than manufactured prices i.e., the response of food prices to change in money supply is higher than the corresponding response of manufactured prices. One percent increase in money supply leads to 0.32 per cent and 0.18 per cent increase in food and manufactured prices, respectively. The coefficient restriction tests reject the hypothesis that these coefficients are equal to one with t-statistics (p-value) of -13.32 (0.00) and -23.03 (0.00), respectively. Thus, these results do not support the long-run neutrality of money hypothesis. Similar results were also obtained in several Asian countries such as Korea, Thailand, Pakistan and South Africa (Saghaian, Hasan and Reed, 2002; Hye and Siddiqui, 2010; and Asfaha and Jooste, 2007). Under the quantity theory of money (QTM) framework, this less than proportionate increase in food and manufactured prices to increase in money supply could follow from either a decline in velocity of money or increase in real output or a combination of both.[4] With regard to velocity of money, Pattanaik and Subhadhra (2011) find that money velocity in India declined persistently during the last six decades due to increasing monetisation of the

[4] Under QTM, MV = PY, where M = money supply, V= velocity of money, P = price and Y = real output. Assuming 'V' is constant and 'M' has no impact on 'Y', any change in 'M' gets translated into an equivalent change in 'P'.

economy. On the impact of money supply on real output, it is highly plausible that the impact is positive in the context of a fast growing economy like India. First, money supply could play a lubricating role to achieve a high growth rate by making available more finance to firms for expansion (Saghaian, Hasan and Reed, 2002). Second, increasing monetisation of the economy could lead to productivity gain. Third, in a typically supply constraint economy like India, it is quite plausible that short-run demand shocks lead to positive long-run supply response.

With regard to call rate, its coefficients in the price equations are positive, implying that monetary tightening through increase in policy rate leads to lowering in prices. However, the coefficient is statistically significant in the equation for manufactured price only. One per cent increase in call rate leads to 0.16 per cent decline in manufactured prices.

As for the impact of change in call rate and money supply on exchange rate, only the latter has a statistically significant long-run impact. One per cent increase in money supply leads to 0.04 per cent depreciation in domestic currency. This follows as increase in money supply induces currency depreciation by increasing the supply of domestic currency per unit of foreign currency.

The short-run dynamics using the residuals of the three cointegrating equations or error correction terms (ECT) in the VECM are presented in Table-6.

Table-6: Vector Error Correction Model

Variable	DLFP	DLMP	DLEX	DLCALL	DLM3
ECT1	**-0.289**	-0.009	-0.119	1.54	-0.286
	(-2.10)**	(-0.16)	(-0.52)	(1.24)	(-2.67)
ECT2	0.249	**-0.253**	0.191	-7.24	0.434
	(0.87)	**(-2.10)****	(0.41)	(-2.80)**	(1.95)
ECT3	-0.352	0.281	**-1.15**	11.73	-0.122
	(-0.72)	(1.36)	**(-1.43)**	(2.65)**	(-0.32)
C	-0.06***	-0.02	0.05	0.15	0.10*
R-Square	0.73	0.61	0.59	0.58	0.72
F-statistic	4.82	3.21	3.06	2.96	4.74
Jarque-Bera	2.12	6.48	0.36	0.72	12.6

Note: Parameter estimates of the first difference form of the variables have not been reported to conserve space, but available on request. The joint significance of these coefficients as a test for short-run causality, however, is reported in table-5.

The estimates satisfy the standard statistical diagnostics, except the normality test in one of the equations.[5] Further all the equations have a relatively much higher explanatory power as compared to those found for several countries in the literature. The coefficients of the three ECTs in the VECM measure the speed of adjustments to its long-run equilibrium after a temporary shock. All the relevant coefficients in the main diagonal indicated in bold are negative and, therefore, correctly signed. They indicate overshooting from the long-run equilibrium in the short-run, as negative signs imply reduction in the value of the variables to restore equilibrium. However, the coefficient of ECT in the exchange rate equation is not statistically significant. The estimated speed of adjustment for food prices and manufactured prices to their long-run equilibrium is -0.289 and -0.253, respectively i.e., overshooting is slightly higher for the food prices than that of manufactured prices.

[5] LM-tests for serial correlation also show that there is no serial correlation up to 15 lags. It has not been reported to conserve space, but available on request.

The short-run causality among the variables through the tests of joint significance of the other lagged endogenous variables in each of the VECM equations are presented in Table-7.

Table-7: VEC Granger Causality Tests-Wald Tests

→ Dependent Variable ↓ Independent Variable	DLFP (1)	DLMP (2)	DLEX (3)	DLCALL (4)	DLM3 (5)
DLFP		+ve (3.3)	+ve (4.1)	+ve (10.5)**	+ve (10.8)**
DLMP	-ve (1.3)		-ve (0.7)	+ve (7.9)***	-ve (18.9)*
DLEX	+ve (13.5)*	-ve (7.7)		-ve (15.5)*	+ve (4.7)
DLCALL	-ve (6.8)	-ve (8.6)***	+ve (2.3)		-ve (11.3)**
DLM3	+ve (9.2)***	+ve (5.1)	-ve (2.6)	-ve (4.4)	

Note: -ve and +ve denote the overall sign of the lag coefficients is negative and positive, respectively. Figure in the parentheses are the Chi-square statistics. *, ** and *** denote statistical significance at 1%, 5% and 10% level, respectively.

The following observation can be made. First, increase in money supply leads to rise in food prices, but not manufactured prices. In complete contrast, increase in call rate has a significant negative impact only on manufactured prices. In other words, in the short-run, interest rate channel of monetary policy is found to be more effective on manufactured prices, while quantum channel is more effective on food prices. With regard to change in exchange rate on prices, it is, however, not clear as to why currency appreciation leads to rise in food prices in the short-run, though the impact is expectedly negative on manufactured prices (Table-7, column 1 and 2).

Second, while increase in food and manufactured prices lead to hike in call rate, the response of money supply is asymmetric.[6] While rise in food prices

[6] The results that there is bi-directional causality between money supply and prices indicates the presence of both exogenous and endogenous element in the process of money supply in India, though under LAF the endogenous process would have strengthened or dominated substantially.

leads to increase in money supply, rise in manufactured prices induces a decline in money supply. This could follow if real demand for food is highly inelastic to price changes that rise in food price leads to increased nominal demand for money. In this context, Kumar *et al* (2011) find that increase in price may not adversely affect the demand for lower-value food commodities in India, indicating high price inelasticity. In contrast, price elasticity of demand for manufactured goods could be very high such that rise in prices reduces real demand for manufacturing goods, and consequently, the nominal demand for money[7].

Section IV: Summary and Concluding Remarks

This paper, drawing on the theoretical and empirical framework available in the literature, examines the relative responses of food and manufactured prices to change in interest rate and money supply in India during the period 2001:Q1 to 2010:Q2. It employs a VECM of five variables viz., food prices, manufactured prices, nominal effective exchange rate, weighted average call rate and money supply. It finds three cointegrating vectors between the variables and, thus, estimates three long-run equilibrium relationships viz., i) food prices on money supply and call rate; ii) manufactured prices on money supply and call rate; and iii) exchange rate on money supply and call rate.

In the long-run, while increase in call rate leads to fall in the prices of only manufactured products, increase in money supply leads to rise in the prices of both food and manufactured products. However, in agreement with both theoretical arguments and cross-country evidences, the impact of money supply on food prices is more than the impact on manufactured prices. Yet, increase in

[7] Though we do not have a concrete evidence on this, the general notion, given the stage of the development and consumption basket of the populace in India is that food items are far more price inelastic than manufactured goods.

money supply leads to less than proportionate increase in the prices of both the commodities, rejecting long-run neutrality of money.

In the short-run, there is overshooting in the prices both food and manufacture products from their respective long-run equilibrium following monetary shocks. In agreement with theoretical arguments and cross-country evidences again, the overshooting, however, is more in food prices than in manufactured prices. On food prices, call rate has no statistically significant negative impact, while money supply has a statistically significant positive effect. In complete contrast, on manufactured prices, money supply has no statistically significant positive effect, while call rate has a negative impact. While increase in both food and manufactured prices induces call rate hike, money supply shows an asymmetric response by way of increasing with rise in food prices and decreasing with the rise in manufactured prices. This could be attributed to demand for food being highly price inelastic while that of manufactured goods being elastic.

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Annex-1: Johansen Cointegration Test (Data Trend with Intercept and No Trend)

Trace Test				
Hypothesised No. of CE(s)	Eigenvalue	Trace Statistic	Critical Value (5%)	Probability
None*	0.966	210.8	69.82	0.00
At most 1*	0.749	86.01	47.86	0.00
At most 2*	0.469	34.84	29.80	0.01
At most 3	0.265	11.40	15.49	0.19
At most 4	0.000	0.001	3.84	0.97
Maximum Eigenvalue Test				
Hypothesised No. of CE(s)	Eigenvalue	Trace Statistic	Critical Value (5%)	Probability
None*	0.966	124.7	33.88	0.00
At most 1*	0.749	51.17	27.58	0.00
At most 2*	0.469	23.44	21.13	0.02
At most 3	0.265	11.40	14.26	0.14
At most 4	0.000	0.001	3.84	0.97

* denotes rejection of the hypothesis at 5% critical level based on MacKinnon-Haug-Michelis (1999) p-values.