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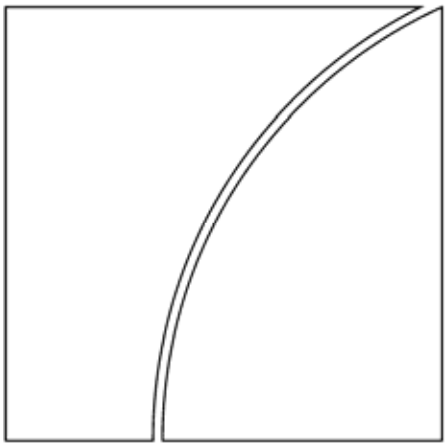

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Economic reforms and exchange rate pass-through to domestic prices in India

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Abstract

This paper examines the behaviour of exchange rate pass-through to domestic prices in India during the post-economic reforms initiated since the major devaluation of July 1991. It observes that there is no clear-cut evidence of a fall in exchange rate pass-through to domestic prices. Further, there is asymmetry in pass-through between appreciation and depreciation, and between sizes of the exchange rate change. Based on the empirical evidence provided in the literature, the paper conjectures that reductions in import tariffs, the removal of trade restrictions, the increased import penetration ratio and openness of the economy and the change in the composition of imports following the economic liberalisation could have transitorily negated the impact of lower inflation on pass-through. Part of the non-decline in long-run pass-through is due to a rise in inflation persistence. This could follow from the dismantling of price controls in an environment of periodic spurts in inflation around a non-declining inflationary trend, combined with a rise in the government deficit, which has a nexus with inflation in India.

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Economic reforms and exchange rate pass-through to domestic prices in India

Jeevan Kumar Khundrakpam[1]

1. Introduction

A macroeconomic puzzle of the 1990s was the phenomenon of low inflation despite episodes of large currency depreciation in several countries. In the cross-country context, this was shown to be the result of low global inflationary environment (for examples, see Taylor (2000), Choudhri and Hakura (2001) and Gagnon and Ihrig (2004)). Several other factors such as exchange rate volatility, import penetration, openness, import composition, trade distortions, transport costs and income have also been identified as important determinants of pass-through (for examples, see Goldfajn and Werlang (2000), Campa and Goldberg (2004) and Frankel et al (2005)). For a country undertaking major economic reforms, these identified macroeconomic variables determining exchange rate pass-through underwent substantial transformation during the transition.

In India, economic reforms were initiated on several fronts since the early 1990s and have led to: a market-determined exchange rate; full convertibility in the current account; a substantial reduction in peak and weighted average tariff rates; the abolishment of import licensing and quantitative restrictions; the encouragement of foreign investment through liberalisation and simplifying procedures; the abolishment of industrial licensing; allowing private sectors in areas earlier reserved for the public sector; decontrol of interest rates, reduction in pre-emption of banking resources and enforcing capital adequacy and prudential norms; government borrowing at market rates and the discontinuation of automatic monetisation of deficit; and the gradual liberalisation of administrative price control mechanism on a number of commodities.

Have these economic reforms affected the exchange rate pass-through to domestic prices in India? Do we observe the same declining phenomenon as in several countries? Currently, there is not much literature in India on this issue. India has been included in some of the cross-country studies (Choudhri and Hakura (2001) and Devereux and Yetman (2003)), but they do not indicate the temporal behaviour. Kapur (2004), based on Mihaljek and Klau (2001), segregates pass-through from import prices in foreign currency and pass-through from exchange rate movements to domestic prices. Using annual data for the period from 1971 to 2004, he observes that while the exchange rate pass-through to wholesale prices exhibited a slightly declining trend, pass-through from import prices to wholesale prices broadly remained unchanged, though a rising trend could be discerned since the mid-1990s. In contrast, Mallick and Marques (2005), using a panel framework on nine product groups of imports on annual data from 1981 to 2001, find that the evidence of exchange rate pass-through to import prices (the first stage of pass-through) in India is evident only after 1991, implying a higher pass-through during the post-economic reforms.

In view of the inconclusive evidence, this paper uses monthly data to further investigate whether the ongoing Indian economic reforms have been associated with a change in the exchange rate pass-through to domestic prices. To preview the results, we do not find

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evidence of decline in exchange rate pass-through to domestic prices in India, despite a lower inflationary environment. Some indicators suggest that a reduction in trade barriers, increased openness and import penetration and a change in import composition associated with economic reforms and a rise in inflation persistence could have contributed to the stable pass-through.

The rest of the paper has six sections. Section 2 provides a review of the literature to identify factors determining pass-through. In Section 3, the stylised facts are briefly described. The data and empirical framework is laid out in Section 4, and the results are presented in Section 5. Section 6 provides a conjectural explanation for the observed trend in pass-through. The final section summarises.

2. What determines pass-through? Review of selected literature

2.1 Theoretical evolution

In the traditional open-economy macroeconomic models, under the purchasing power parity (PPP) assumption, exchange rate pass-through to domestic prices is always immediate and complete. Thus, research on it was microeconomic in nature and the explanation for the evidence of incomplete pass-through was based on imperfect competition and pricing to market theory (Dornbusch (1987) and Krugman (1987)). Firms, in order to maintain market share, adjust their mark-up instead of fully passing the exchange rate movement to prices (for a survey, see Goldberg and Knetter (1996)).

Analysing pass-through from a macroeconomic perspective is related to developments in open-economy macroeconomics literature beginning with Obstfeld and Rogoff (1995), and its further extension by Betts and Devereux, (1996, 2000) and others. In this new open macroeconomic model, pass-through depends upon the pricing strategy of firms. Under producer currency pricing (PCP), the home currency price of foreign goods will move one-for-one with changes in the nominal exchange rate, ie, pass-through is complete. With local currency pricing (LCP), there is no change in the short-run prices faced by the consumer, and thus no pass-through. The aggregate pass-through, therefore, depends upon the combination of firms practising PCP and LCP (Betts and Devereux (1996) and Engle (2002)).

Another influential view, based on staggered price setting and monopolistic competition behaviour of firms, is that a credible low inflation regime leads to a lower pass-through (Taylor (2000)).

2.2 Cross-country evidence

Goldfajn and Werlang (2000), using a sample of 71 countries, find that the main determinants of pass-through are the cyclical component of output, the initial overvaluation of the real exchange rate (RER), the initial rate of inflation and the degree of openness. Among them, the RER misalignment is the most important determinant for emerging markets, while it is initial inflation for developed countries.

For pass-through to import prices, Campa and Goldberg (2004) for 25 OECD countries find the composition of imports to be more important in explaining the behaviour of pass-through than inflation and exchange rate volatility.[2]

[2] On the contrary, Otani et al (2003) in the case of Japan find that the decline in the exchange rate pass-through to import prices came mainly from declines in each product, rather than a shift in composition of imports.

On domestic prices, Choudhri and Hakura (2001) in 71 countries, consisting of both developed and developing countries, find a strong positive association between pass-through and the average inflation rate across countries. The inflation rate was found to dominate other macroeconomic variables in explaining cross-country differences in pass-through. Similarly, in 122 countries, Devereux and Yetman (2003) find a positive non-linear relationship between pass-through and mean inflation and exchange rate.

Gagnon and Ihrig (2004) test whether the change in pass-through in 20 industrialised countries is explained by change in inflation regime. Relating the estimated pass-through pre- and post-inflation regime change for each country with the corresponding inflation regimes, they find that the decline in pass-through is explained by the fall in inflation variability. Bailliu and Fujii (2004) for 11 industrialised countries also find that pass-through declines with a shift to a low-inflation environment brought about by a change in monetary policy.

Frankel et al (2005) on eight select goods of 76 countries during 1990-2001 also find decline in pass-through, which was much more rapid for developing countries than high-income countries. They find that per capita income, bilateral distance, tariffs, country size, wages, long-term exchange rate variability and long-term inflation are important determinants of pass-through.

2.3 VAR approach

Pass-through is also analysed under a recursive VAR framework by assessing the impulse responses and variance decomposition of various prices along the distribution chain obtained from shocks to exchange rate and import prices. For six industrialised countries, the pass-through is found to decline along the distribution chain, with only a modest effect on consumer prices. Further, the pass-through is found to be stronger for more open economies (McCarthy (1999)).

Several studies adopting this approach in individual countries have found lower pass-through along the distribution chain, and decline in the pass-through with lower inflation (for examples, see Bhundia (2002) for South Africa, Leigh and Rossi (2002) for Turkey and Belaisch (2003) for Brazil). Kang and Wang (2003), however, find that in Thailand and Korea the pass-through increased after the Asian crisis due to higher exchange rate volatility resulting from the adoption of a free-floating exchange rate regime and an increase in the trade to GDP ratio.

2.4 Asymmetry in pass-through

Unlike that which is assumed in the above studies, pass-through can differ between depreciation and appreciation and between large and small changes. The standard explanations for asymmetry are as follows.

Binding quantity constraints: When foreign firms face capacity constraints in their distribution networks, they limit the ability to increase sales in the importing country. An appreciation of the importing country's currency would normally induce a foreign firm to lower import price, but capacity constraints limit such expansion of sales through the lower price. Thus, foreign firms raise their mark-ups to keep import prices in the importing country's currency fixed to keep the volume of sales intact while raising the profit margin. In the case of depreciation, the same capacity constraint is not binding and does not affect the raising of import prices

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Marazzi et al (2005) find that shift in the composition of core imports provides only a partial explanation for decline in the aggregate pass-through to US import prices.

that depreciation would normally induce. Even when firms may reduce their mark-ups to absorb part of the impact of depreciation, import prices in home currency could still rise. Thus, the pass-through is higher for depreciation than appreciation. Besides capacity constraints limiting the ability to expand output, this situation can also arise when there are trade restrictions (Knetter (1994), Pollard and Coughlin (2004)).

Market share objective: However, when firms are building up market share, appreciation in the currency of the importing country will allow the firms to lower import prices to increase their market share while maintaining their mark-up. But in the case of depreciation, the exporting firms may offset the potential increase in price to maintain their market shares by reducing mark-ups. Thus, pass-through would be higher for appreciation than depreciation (Knetter (1994)).

Production switching: In this model, foreign firms use imported inputs switch between imported and domestically produced inputs depending upon the price. When the importing country's currency appreciates, foreign firms use only the input produced in their own country, and the extent of pass-through depends on the elasticity of the mark-up. In the case of depreciation, foreign firms use inputs from the currency depreciating country, and no pass-through occurs (Webber (2000)).

Menu cost: The asymmetry with respect to the size of change in the exchange rate is explained by the presence of menu cost and the type of price invoicing followed. Given that menu cost is like a fixed cost, changing invoice price is worthwhile only if the exchange rate change is above a certain threshold. If the invoice currency is the foreign firm's currency, a small change in the exchange rate does not make it worthwhile for the foreign exporter to change the price of its product in its own currency because of the menu cost involved in that change. Thus, a small exchange rate change does not affect the invoice price of imports in foreign currency, leading to a change in import prices in domestic currency to the extent of the change in exchange rate, and therefore a higher pass-through. However, when the exchange rate change is large, given the menu cost, it is worthwhile for the foreign exporter to change the invoice price in its own currency. Thus, by altering the invoice price in its own currency, the exporter absorbs part of the price variation in the importing country's currency that will follow from a large exchange rate change and reduces the extent of pass-through (Pollard and Coughlin (2004)).

In the case of invoice in the importer's currency, a small change in the exchange rate again has little impact on the invoice price because of the menu cost. Thus, the import price in the importing country's currency remains largely unchanged and there is little or no pass-through of the exchange rate change. However, the price received by the foreign exporter in its own currency changes to the extent of the exchange rate variation. But this is affordable to the foreign exporter as the exchange rate variation is small and less than the menu cost involved in changing the invoice price. However, if the exchange rate change is large, the corresponding change in foreign exporter's receivable in its currency would also be large at an unchanged invoice price and larger than the menu cost. Thus, when the exchange rate change is large, foreign exporters alter the invoice price in the importing country's currency to keep proceeds in their own currency intact, thereby leading to higher pass-through (Pollard and Coughlin (2004)).

Many empirical studies support asymmetry in pass-through, but the directions have varied. For the US, at the aggregate level of import prices, Mann (1986) found higher pass-through during appreciation than depreciation, though the difference is statistically insignificant. Marazzi et al (2005), however, find no such asymmetry. In contrast, for seven Asian countries, Webber (2000) finds more support for higher pass-through for depreciation than appreciation.

At the industry level in the US, pass-through is found to be higher for depreciation: this was found by Goldberg (1995) in the automobile industry; Kadiyali (1997) in the photographic industry and Olivei (2002) in nine out of the 34 US import industries. Pollard and Coughlin

(2004) also find asymmetry in pass-through in half of the 30 industries considered, but the direction varied between the industries.

With regard to asymmetry between the size of exchange rate changes, Ohno (1989) finds Japanese export prices to respond more to large exchange rate changes than small changes. For 30 US import industries, Pollard and Coughlin (2004) also find most firms responding positively to the size of exchange rate change. At the aggregate level in the US, Marazzi et al (2005), however, find the response to be symmetrical.

2.5 Factors explaining pass-through

From the above studies, we identify several factors affecting exchange rate pass-through. First, the higher the rate of inflation and its volatility, the higher the pass-through, as firms perceive any increase in the cost of production to be more persistent in such an environment. Conversely, improved credibility and effectiveness of monetary policy in maintaining a low inflation regime will lower the pass-through, as inflation is anchored at a lower level. Firms are thus less keen to alter prices arising from shocks on cost, as they believe that monetary policy will be successful in stabilising prices.

Second is the volatility of exchange rate. However, the literature is not unambiguous on whether the impact is positive or negative.[3]

Third, the larger the share of imports in the consumption basket (the higher the import penetration ratio) the greater the pass-through would be. Also, the greater the proportion of imported inputs in production, the greater the impact of the exchange rate on the producer's price will be. Thus, the higher the degree of openness of an economy (larger presence of imports and exports), the larger the pass-through coefficient.[4]

Fourth, the composition of imports also affects the aggregate pass-through, as the degree of pass-through differs among various categories of imports. For example, pass-through to manufactured products is found to be less as compared to energy and raw material products. Thus, a rise in the share of the former and a fall in the shares of the latter will lead to lower aggregate pass-through even when the pass-through to individual components remains the same.

Fifth, trade distortions, resulting from tariffs and quantitative restrictions, act as a barrier to arbitrage of goods between countries and lead to lower pass-through.

Sixth, in the presence of asymmetry, the pass-through would depend upon the period of appreciation and depreciation and the size of exchange rate change during various sub-periods.

Last but not least, factors such as income and transportation costs are also hypothesised to have a negative effect on pass-through.

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[3] For example, while McCarthy (1999) and Frankel et al (2005) argue for a negative relationship, Choudhri and Hakura (2001) and Devereux and Yetman (2003) support a positive relationship.

[4] The other view is that openness has a negative correlation with inflation, as openness puts a check on inflationary finance (Romer (1993)). However, this argument is not strictly concerned with pass-through of exchange rate movement to prices. Further, Alfaro (2005) argues that the correlation found by Romer might be driven by time-invariant omitted variables in the cross-section analysis. By introducing time and country dummies to capture the effect of time-invariant variables in a panel framework, he finds no negative correlation between openness and inflation. Further inclusion of exchange rate regimes in the panel estimate, on the contrary, leads to a positive and significant relationship between openness and inflation, and also lower inflation under a fixed exchange rate (which would imply lower volatility) than under a floating exchange rate.

3. Some stylised facts

In Table 1, the annualised month-to-month average inflation rate, exchange rate change and their volatility for three sub-sample periods are given. It is seen that both the average inflation rate and volatility declined, with the decline in the former being much more prominent than in the latter. However, the decline in inflation rate was primarily limited from the first half of the 1990s to the second half of the decade. The average depreciation rate also declined considerably due to increasing two-way movements in the more recent times. However, volatility appears to have increased somewhat, reflecting that India has graduated from a regime of officially determined exchange rate to a market determined exchange rate, though the central bank intervenes substantially to check extreme volatility.

Table 1

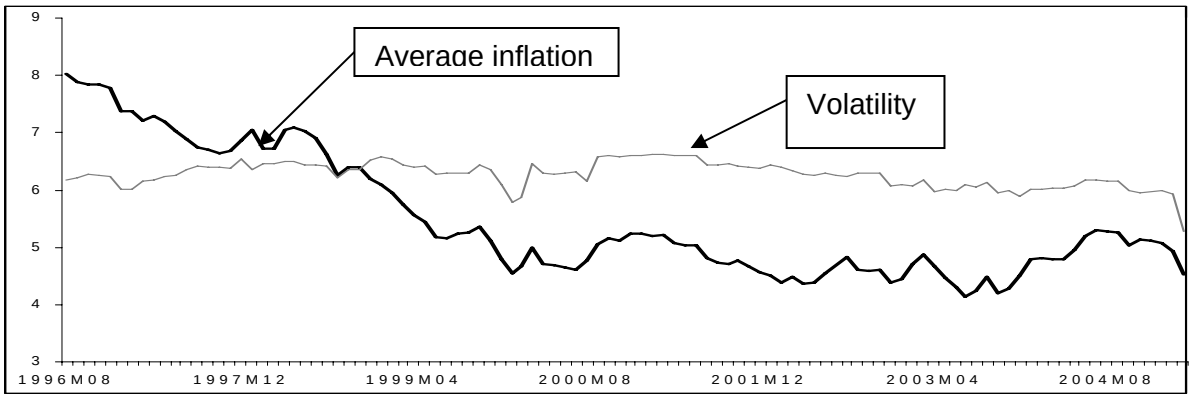
Annualised average inflation, exchange rate and their volatility

(in per cent)

Period	Average inflation	Inflation volatility	Average exchange change	Exchange volatility
1991:8 to 1995:3	9.69	6.84	5.61	19.61
1995:4 to 2000:3	5.00	6.46	2.91	20.22
2000:4 to 2005:3	4.53	5.27	0.08	21.61
Overall	6.09	6.50	2.60	20.57

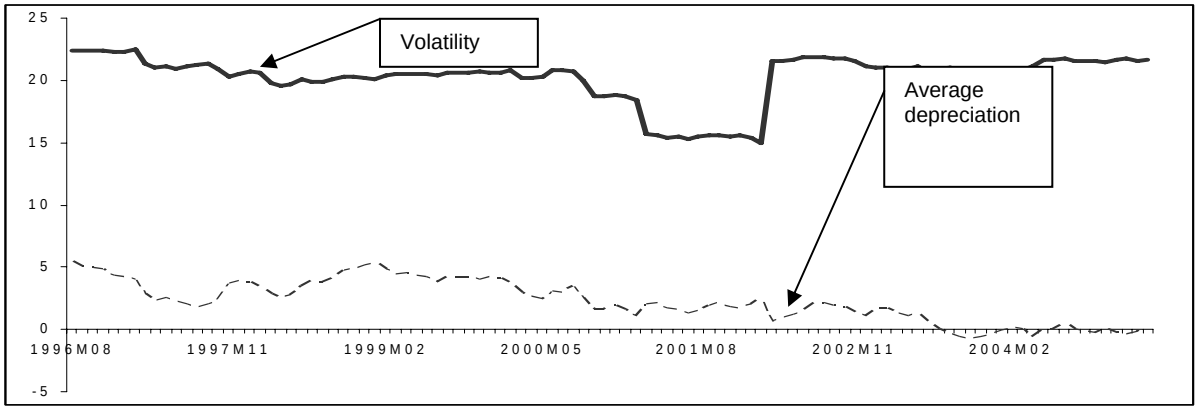
The three sub-sample periods, however, cannot capture the evolution in trend. Thus, rolling averages for five years are shown in Figures 1 and 2. From Figure 1, it is observed that the average month-to-month inflation on an overall basis has declined from over 8% to around 4 to 5%, but the declining trend ended by the mid-1990s, reflected in five years moving average flattening out for the month ending around the beginning of 2000. At the same time, one can also observe spurts in inflation in the later part of the sample period. Volatility in inflation, however, remained more or less stable.

Figure 1: Five-year rolling average inflation and volatility



Similarly, the average rate of depreciation steadily declined due to increasing two-way movements, and leading to appreciation on average in the more recent period (Figure 2). Volatility, however, has not changed much.

Figure 2: Five-year rolling average exchange rate change and volatility



4. Empirical framework

4.1 Model estimated

Drawing on the literature (for example, Bailliu and Fujii (2004)), a reduced form specification for estimating pass-through coefficient is derived from the profit maximising behaviour of an exporting foreign firm of the following type:

$$\underset{P}{Max}\pi = e^{-1}PQ - C(Q) \tag{1}$$

where 'π' is profit in the exporting firm's currency, 'e' is the exchange rate of domestic currency per the exporting firms currency, 'P' is price in domestic currency, $C(\ )$ is the cost function in the exporting firm's currency and 'Q' is the quantity demanded. The first order condition for maximisation of (1) is derived as

$$P = eC_q\mu \tag{2}$$

where 'C_q' is the marginal cost and 'μ' is the mark-up over marginal cost which depends on the price elasticity of demand of the good. Thus, the price in domestic currency 'P' can change as a result of exchange rate, change in marginal cost of the firm and mark-up. The marginal cost will change because of local input cost, while the mark-up can change due to change in demand factors in the importing country. Thus, in reduced form, the price equation is written as

$$P_t = \alpha_0 + \alpha_1 e_t + \alpha_2 P_t^* + \alpha_3 Y_t + \varepsilon_t \tag{3}$$

where 'P^*' is the exporting firm's marginal cost and 'Y' is demand conditions in the importing country. In the literature, variants of (3) are used to estimate pass-through (see Goldberg and Knetter (1996)).

For estimating pass-through at the aggregate price index level, (3) needs to be adapted taking into account the following issues. First, as the aggregate price and exchange rate are generally assumed to follow a non-stationary process, I(1), it is common to use specifications in first difference, ie, in the form of an inflation equation (see Bailliu and Fujii (2004) among others). In our case also the specification is considered in first difference, as the unit root properties discussed in Appendix A show that the series are I(1) and the cointegration tests reveal no cointegration. Second, the lagged effects of the explanatory variables need to be taken into account, thereby yielding an inflation equation of the type

$$\Delta P_t = \alpha_0 + \alpha_1 \sum_{i=0}^{n} \Delta e_{t-i} + \alpha_2 \sum_{i=0}^{n} \Delta P^*_{t-i} + \alpha_3 \sum_{i=0}^{n} Y_{t-i} + \varepsilon_t \quad (4)$$

Third, (4) being essentially an inflation equation, there is a need to account for inflation inertia. Following an adaptive inflation expectation approach, lags of inflation are included as explanatory variables, which also allow for distinguishing between short- and long-run pass-through:

$$\Delta P_t = \alpha_0 + \alpha_1 \sum_{i=0}^{n} \Delta e_{t-i} + \alpha_2 \sum_{i=0}^{n} \Delta P^*_{t-i} + \alpha_3 \sum_{i=0}^{n} Y_{t-i} + \alpha_4 \sum_{i=1}^{n} \Delta P_{t-i} + \varepsilon_t \quad (5)$$

Fourth, in India, shocks in primary commodities, particularly food prices, often affect the general price level. Following Mohanty and Klau (2001), the included food price shock variable, *fshock,* is defined as the excess of current food price inflation over the general price inflation of the previous year. The final augmented equation is of the following type.

$$\Delta P_t = \alpha_0 + \alpha_1 \sum_{i=0}^{n} \Delta e_{t-i} + \alpha_2 \sum_{i=0}^{n} \Delta P^*_{t-i} + \alpha_3 \sum_{i=0}^{n} Y_{t-i} + \alpha_4 \sum_{i=1}^{n} \Delta P_{t-i} + \alpha_5 \sum_{i=0}^{n} fshock_{t-i} + \varepsilon_t \quad (6)$$

The lagged inflation term gives the speed of pass-through to inflation. The short-run pass-through coefficient is given by α_2 and the long-run coefficient by $\alpha_2/(1-\alpha_1)$.

4.2 Asymmetry

Following Pollard and Coughlin (2004) and others, the asymmetries are estimated by interaction of the exchange rate variable with appropriate dummies in the following manner.

4.2.1 Asymmetry with respect to appreciation and depreciation

Two dummies for appreciation and depreciation, respectively, are created as

D_A = 1 for $\Delta e < 0$; 0 otherwise and D_D = 1 for $\Delta e > 0$; 0 otherwise

Interaction of the above dummies with exchange rate change in equation (6) yields

$$\Delta P_t = \alpha_0 + \alpha_{1A} D_A \sum_{i=0}^{n} \Delta e_{t-i} + \alpha_{1D} D_D \sum_{i=0}^{n} \Delta e_{t-i} + \alpha_2 \sum_{i=0}^{n} \Delta P^*_{t-i} + \alpha_3 \sum_{i=0}^{n} Y_{t-i} + \alpha_4 \sum_{i=1}^{n} \Delta P_{t-i} + \alpha_5 \sum_{i=0}^{n} fshock_{t-i} + \varepsilon_t$$

with α_{1A} and α_{1D} providing the separate pass-through coefficients for appreciation and depreciation, respectively.

4.2.2 Asymmetry with respect to size of exchange rate change

The two dummies for absolute large and small change, respectively, are:

D_L = 1 for $\Delta e >$ threshold; 0 otherwise and D_S = 1 for $\Delta e <$ threshold; 0 otherwise

Interaction of the above dummies with the exchange rate change in equation (6) yields

$$\Delta P_t = \alpha_0 + \alpha_{1L} D_L \sum_{i=0}^{n} \Delta e_{t-i} + \alpha_{1S} D_S \sum_{i=0}^{n} \Delta e_{t-i} + \alpha_2 \sum_{i=0}^{n} \Delta P^*_{t-i} + \alpha_3 \sum_{i=0}^{n} Y_{t-i} + \alpha_4 \sum_{i=1}^{n} \Delta P_{t-i} + \alpha_5 \sum_{i=0}^{n} fshock_{t-i} + \varepsilon_t$$

with α_{1L} and α_{1S} providing separate pass-through coefficients for large and small changes, respectively.

5. Data and empirical results

5.1 Source of data

We use the monthly data from the period 1991:8 to 2005:3 from the Handbook of Statistics on the Indian Economy, Reserve Bank of India (RBI). The variables used are: wholesale price index (*P*)[5], nominal effective exchange rate (*e*) defined as domestic currency per unit of foreign currency, index of industrial production (*Y*), and trade weighted foreign prices (*P**). Trade weighted foreign prices are derived using the definition of real effective exchange rate adopted by the RBI. As the real effective exchange rate (*rer*) is defined as the weighted average of nominal effective exchange rate (*e*) × [wholesale price inflation (*P*) ÷ foreign inflation (*P**)], we can derive *P* = (e × P) ÷ rer*.[6] All the series are seasonally adjusted.

5.2 Robustness tests

As a test for robustness of the pass-through coefficients, three models as specified above were estimated. The first model (model A) is specification (4). The second model (model B) is specification (5) and the third model (model C) is specification (6). The lag lengths of the variables were chosen using the general-to-specific method starting from 11 lags, as they are monthly data, and progressively dropping the insignificant lags. Table 2 presents the results.[7] It can be seen that model A is well estimated, with all the variables statistically significant at the conventional level, has reasonable explanatory power and does not suffer from a serial correlation problem. It is observed that there is substantial difference between the coefficient of foreign price (proxy of input cost) and the exchange rate. In the literature, some studies (for example, Choudhri and Hakura (2001) and Gagnon and Ihrig (2004)) consider the two to be the same, as it is assumed that foreign firms respond symmetrically to change in input cost and exchange rate. In our case, while the exchange rate pass-through is 0.066%, the pass-through of input cost is as large as 0.37%, with the Wald test [14.0(0.00)] decisively rejecting the equality of the two coefficients. One possible explanation for the difference is

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[5] We consider the WPI as it is the headline measure of inflation in India. It is computed on an all-India basis with larger coverage of commodities, higher frequency and with lesser lag of release coinciding with the release of monetary data. The CPI, which is constructed for specific groups of people and centres, and then aggregated to get the all-India index, however, is released at a lower frequency and at a higher lag. Thus WPI is more easily understood by the public than CPI and is also more easily monitored from the monetary policy point of view.

[6] Effective exchange rates in India are calculated alternatively as export-weighted and trade-weighted five-country and 36-country bilateral weights. Thus, there are four types of effective exchange rate, and here we consider the most comprehensive in terms of coverage, ie, 36-country bilateral trade-weighted.

[7] In the estimates, dummy variables were used to control for few instances of month-to-month rate of domestic and foreign inflation, which on an annualised basis were over 18%, and not explainable by the variables considered. Month-to-month inflation series are generally extremely volatile, either due to measurement error or temporary factors unrelated to underlying inflation trends. The included dummies thus control the effect of these extreme volatilities.

that change in input cost is considered to be more permanent than change in exchange rate that the foreign firms pass-through the former much more than the latter.

Table 2

Robustness tests – 1991:8 to 2005:3

Model	Constant	$\sum\Delta e$	$\sum\Delta P^*$	$\sum\Delta Y$	$\sum\Delta P_{t-1}$	$\sum$fshock	R^2	DW
Model A	0.0035	0.066	0.368	0.55			0.45	1.80
	(9.7)	(5.0)	(5.4)	(2.3)				
Model B	0.0023	0.067	0.47	0.047	0.22		0.49	2.19
	(4.9)	(3.8)	(5.8)	(1.9)	(3.6)			
Model C	0.0021	0.063	0.41	0.048	0.29	0.14	0.59	1.94
	(5.0)	(3.9)	(5.7)	(2.1)	(5.2)	(6.1)		

Note: The figures in parentheses are t-statistics.

The inclusion of the autoregressive term (model B) to incorporate inflation expectation does not alter the pass-through coefficient, though it can now differentiate between the short-run and long-run pass-through while the explanatory power improves. There are some alterations in the impact of foreign prices and domestic demand shocks. Further inclusion of food price shock (model C) leaves the coefficients of the variables more or less unaltered. However, the explanatory power and the significance level of the variables improve, the former markedly. In the remainder of the paper, we analyse pass-through using model C, though the results obtained from the other models are reported in Appendix B. This model is preferred as food price shock is understood to play an important role in the inflationary process in India. Further, it is found that the coefficients of lags of inflation are much more robust with the inclusion of food price shock than without.

5.3 Comparison of coefficients with earlier estimates

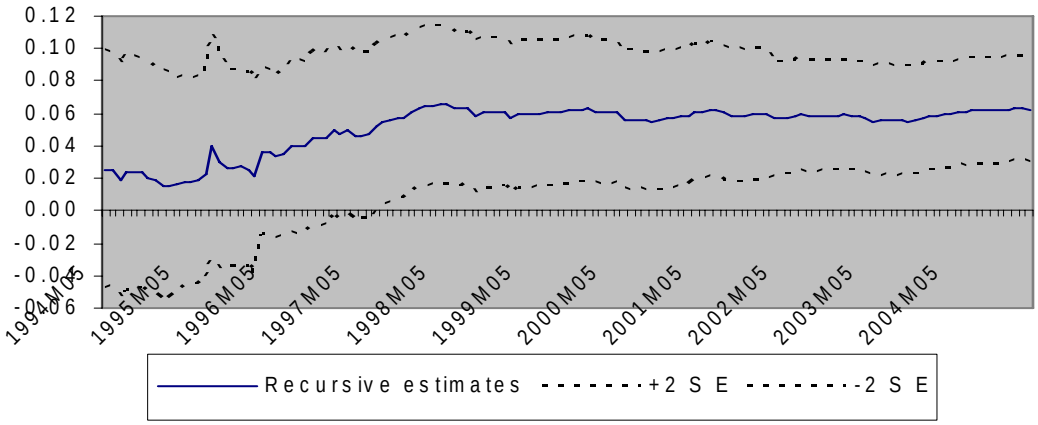
Our estimated pass-through coefficients of 0.063 in the short run and 0.09 in the long run appear to be rather small, as they imply that a 10% change in exchange rate increases final prices by 0.6 to 0.9% only. However, such magnitudes of pass-through to final prices are typically found and are very similar in size to the estimates of Choudhri and Hakura (2001) for India even though they use quarterly data for 1979 to 2000. The short-run pass-through is larger than the average pass-through of low-inflation countries and several industrialised countries, while the long-run pass-through is about the average reported by Choudhri and Hakura. As they use accumulated quarterly data, the coefficient estimated by them is expected to be larger, though it is not, since inflation gets accumulated over a quarter/year (the frequency of observation) while the accumulation in exchange rate change may not be large due to two-way movement during the same interval. Reflecting this property, Devereux and Yetman (2003), using annual data during 1970 to 2001, estimate a much higher pass-through of 0.36 for India.[8]

[8] However, they use a different specification in which inflation is explained by exchange rate and foreign price (proxy for cost) only.

5.4 Stability of coefficients in the model

The parameter stability test of Hansen (1991) shows that the coefficients are stable.[9] The recursive estimates of the coefficients, however, show an upward trend, at least in the beginning of the sample period (Figure 3). As they may indicate a gradual change in the pass-through coefficients, rolling regressions were estimated (see Appendix B for robustness check).[10]

Figure 3: Recursive estimates of coefficients



5.5 Pass-through coefficients from rolling regressions

Figures 4 and 5 show the short-run and long-run pass-through rolling regression coefficients, respectively, along with the 90% confidence intervals.[11] The short-run pass-through coefficients appear to resemble a flattened inverted V-shape, ie an upward slope in the beginning and a downward slope thereafter, while long-run pass-through seems to follow a very low upward slope.

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[9] The test statistics for Δe, ΔPt-1 and joint statistics are 0.19, 0.22 and 1.2, which is less than 5% critical values of 0.47, 0.47 and 1.2, respectively.

[10] The coefficients from the rolling regressions are often sensitive to the window size. Therefore, as a test for robustness and stability, we did rolling regressions for window sizes of five, six and seven years. We find no fundamental differences in the overall trend in the coefficient between the different window sizes and across the alternative models.

[11] The standard error of the long-term coefficients to derive the confidence interval is estimated as

$$s.e. = \sqrt{\left(\frac{1}{(1-\alpha_1)}\right)^2 Variance[\alpha_2] + \left(\frac{\alpha_2}{(1-\alpha_2)^2}\right)^2 Variance[\alpha_1] + 2\left(\frac{1}{(1-\alpha_1}\right)\left(\frac{\alpha 2}{(1-\alpha_2)^2}\right) Covariance(\alpha_1, \alpha_2)}$$

Figure 4: Short-run rolling regression coefficients

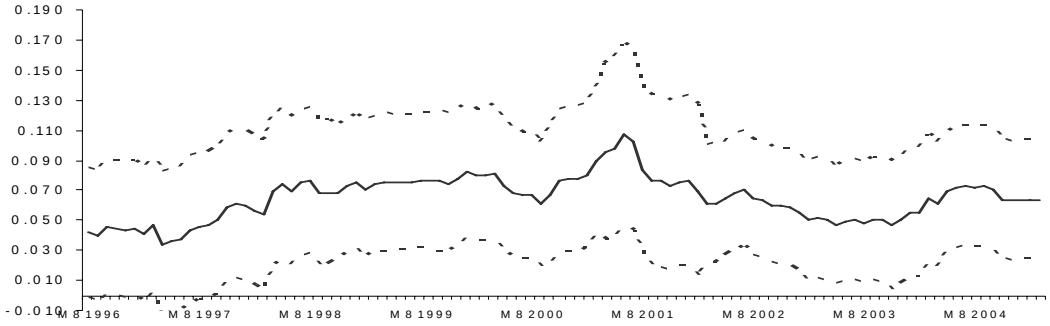


Figure 5: Long-run rolling regression coefficients

To ascertain the slope, the two types of trend were fitted:

$$1) PT = \beta_0 + \beta_1 Trend$$

where *PT* is the series of pass-through coefficients obtained from the rolling regressions. For $\beta_1 > 0$, pass-through has an overall positive slope, $\beta_1 < 0$ implies pass-through has an overall negative slope and for $\beta_1 = 0$ there is no change.

$$2) PT = \beta_0 + (\beta_2 + \beta_0) Dummy + \beta_1 Trend + (\beta_1 + \beta_3) Dummy * Trend$$

which is a kinked trend to check for the presence of a break in the slope of the pass-through coefficients over the sample period. *Dummy* takes a value of 1 from the point of significant departure from the overall trend and thereafter, and 0 otherwise.[12] The point which gave the highest R-bar square, ie, the best fit of the regression, was chosen. The increase/decrease before the identified point is given by β_1, and thereafter it is given by $(\beta_1 + \beta_3)$. For $\beta_1 > 0$, the slope is steeper after the identified point when $\beta_3 > 0$, less steep with $\beta_3 < 0$ and falls for $-\beta_3 > \beta_1$.

The results are presented in Table 3. The simple trend fits show that the slope of short-run pass-through coefficients is positive but not statistically significant at the 5% level. Further, the goodness of fit (R-bar square) is very poor (only 0.024). The trend fit with structural break gives a far better fit and indicates a positive slope up to the window ending 2001:09 and then

[12] Since this point is not known a priori, we search for it within a range of 15-85% of the sample period, which is the standard practice in the literature to locate structural breaks at an unknown point in time.

a negative slope though not statistically significant. The long-run pass-through has a more positive slope and turns steeper from 2002:03, partly due to a lower base level as reflected by the negative value of intercept coefficient of –0.038.[13] Thus, the values of slope coefficients are very small, which in the case of the short run is not even significant at the conventional level.

Table 3

Trend fits on rolling regression coefficients

Pass-through	β_0	β_1	$\beta_2+\beta_0$	$\beta_1+\beta_3$	R-bar square	Wald test $\beta_1=0$	Wald test $\beta_1+\beta_3=0$	Break at window ending
Short-run	0.059	0.00023			0.024	3.51		
	(16.3)	(1.87)				[0.06]		
Short-run	0.026	0.0008	0.075	–0.00013	0.62	165	1.36	2001:09
	(7.8)	(12.8)	(6.76)	(–1.2)		[0.00]	[0.24]	
Long-run	0.053	0.0005			0.47	92.0		
	(13.6)	(9.6)				[0.00]		
Long-run	0.038	0.0009	–0.038	0.0013	0.65	115	44.4	2002:03
	(8.5)	(10.7)	(1.86)	(6.6)		[0.00]	[0.00]	

Note: The figures in round brackets are t-statistics, while those in square brackets are p-values.

A positive slope in the trend of rolling regression coefficients, however, does not necessarily ensure that the coefficients are statistically different from each other. Therefore, Wald tests were performed between the coefficients of six pairs of windows. The compared windows are the first and the ninth (with the lowest short-run and long-run coefficient) against the windows with the highest short-run coefficient (58th), the highest long-run coefficient (95th) and the last window (104th). The results are presented in Table 4.

[13] A similar trend is indicated in all the window sizes and for the alternative specifications (see Appendix B).

Table 4

Change in pass-through coefficients and Wald test

	58th minus first	58th minus ninth	95th minus first	95th minus fifth	Last minus first	Last minus fifth
Short run	0.065	0.073	0.031	0.040	0.022	0.030
	(0.16)	(0.13)	(0.39)	(0.30)	(0.55)	(0.43)
Long run	0.079	0.093	0.085	0.098	0.062	0.076
	(0.19)	(0.13)	(0.20)	(0.15)	(0.32)	(0.23)

Note: The reported statistics in brackets are p-values of the F-test of the null hypothesis that the increase is equal to zero.

From Table 4, it is evident that the changes in the pass-through coefficients are not statistically significant for both the short and the long run between any pair of windows. The significance level goes down progressively with the increase in the distance of the windows.

5.6 Asymmetry

5.6.1 Asymmetry between depreciation and appreciation

We find that the estimated pass-through coefficients are higher for appreciation than depreciation. The respective coefficients shown in Table 5 are 0.141 and 0.037 for short-run, and 0.20 and 0.053 for long-run. The Wald tests indicate that the coefficients are statistically different from each other at the 10% level (the test statistics are 3.46 (0.063) for short-run and 3.31 (0.069) for long-run).

The higher pass-through for appreciation than depreciation is expected in the Indian case. Under the inward-looking strategy of industrialisation, most of the goods were domestically produced irrespective of quality. Thus, when foreign exporters increasingly entered the Indian market after liberalisation they would have faced some degree of competition from the locally produced substitutes. Therefore, the objective of foreign firms would likely be to hold or increase their market share. As discussed in Section 2.4, in such a situation pass-through from appreciation would be higher than depreciation as the foreign firms would be more willing to pass on the benefit of lower prices from appreciation to domestic consumers and capture a larger market share than pass on the higher prices from depreciation and lose their market share.

Table 5

Asymmetry in pass-through

	Depreciation		Appreciation		Large		Small	
Direction/size	**Short**	**Long**	**Short**	**Long**	**Short**	**Long**	**Short**	**Long**
Direction	0.037	0.053	0.141	0.20				
	(3.1)	(1.3)	(3.6)	(3.4)				
Size								
1) >sample average (10.6%)					0.053	0.076	0.148	0.21
					(3.2)	(3.1)	(2.9)	(2.8)
2) >median (16%)					0.058	0.081	0.152	0.214
					(3.5)	(3.4)	(2.2)	(2.2)
3) >24%					0.039	0.055	0.128	0.18
					(2.2)	(2.1)	(4.2)	(4.0)

Note: Figures in brackets are t-statistics.

5.6.2 Asymmetry between sizes of exchange rate change

As there are no set criteria to select threshold levels distinguishing large and small absolute exchange rate changes determined by the level of menu cost (see Pollard and Coughlin (2004), for example), three annualised rates of change were considered, viz, median (10.6%), sample average (16%) and 24% (a value higher than the sample average). As can be seen from Table 5, the estimated pass-through coefficients are found to be much higher for small than large changes, and do not vary much between the three alternative threshold levels. In the short run, pass-through from small changes range from 0.128 to 0.148, as against the range of 0.039 to 0.058 for large changes, while in the long run the respective ranges are 0.18 to 0.21 and 0.055 to 0.081. The Wald tests show that the coefficients are statistically different from each other for 16% and 24% change in exchange rate as the threshold levels.[14]

This result is again expected as most of the imports in India (over 80-90%) are priced in the exporter's currency (PCP) or invoiced in US dollars (RBI (2001)). Following the discussion in Section 2.4, under the presence of menu cost, the pass-through is expected to be higher for small exchange rate change than for large change. This follows as the menu cost (a fixed cost) involved in altering the invoice price could be larger and prohibitive as compared to the loss of market share and revenue resulting from allowing small exchange rate changes to be passed through to the domestic prices of the importing country. In contrast, the loss of market share and revenue resulting from pass-through of large exchange rate changes to the domestic prices of the importing country would be much larger than the menu cost (fixed cost) involved in altering the invoice price in the exporter's currency.

[14] The Wald test results are short-run 6.3 (0.012) and long-run 6.14 (0.013) for 24% as the threshold level, and 3.13 (0.08) and 3.02 (0.08) for 16% as the threshold level.

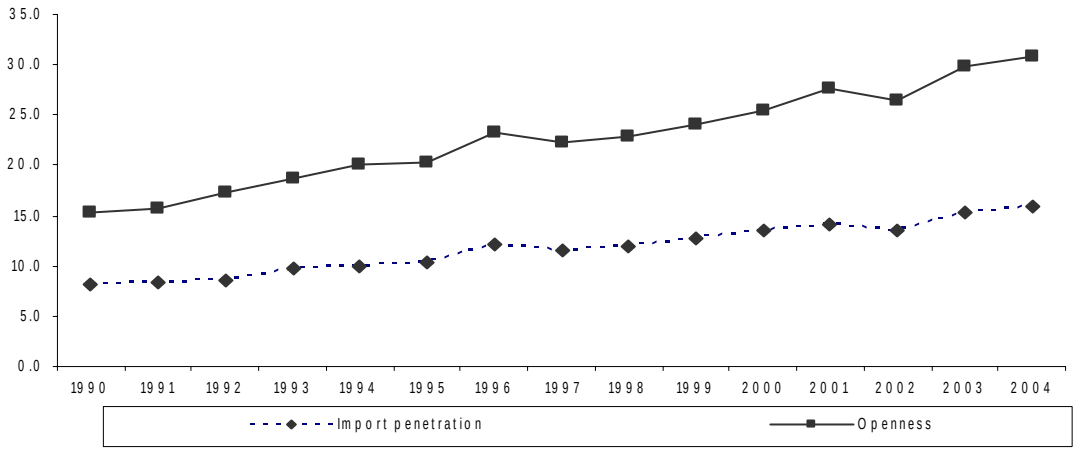
6. Why pass-through has not declined: possible factors[15]

As observed in Figure 1, the average inflation rate declined from about 10% and hovered around the range of 4 to 5% since the second half of the 1990s, even though the volatility in inflation did not decline to the same extent. The findings in the literature on the experiences of several countries indicate that this trend in inflation should have led to a decline in pass-through. For the short run, this appears to be the case, at least in the later part of the sample period. For the long run, non-decline in pass-through is more evident due to the increase in persistence in pass-though for a longer horizon, reflected in the higher coefficient of the lags of inflation. The following could be some of the transitory factors during the post-economic reform period which kept the pass-through from declining by negating the impact of lower inflation.

6.1 Openness and import penetration

It has been shown that pass-through is higher for more open economies as measured by the ratio of exports and imports to domestic income (Dornbusch, Fisher and Samuelson (1977), Goldfajn and Werlang (2000) and Kang and Wang (2003)). Further, a rise in the participation of foreign firms in the domestic economy or a rise in the import penetration ratio, measured by the share of imports in consumption, is also found to lead to higher pass-through (see McCarthy (1999)). Figure 6 shows that both the import penetration ratio and openness have increased substantially since the beginning of 1990s, roughly doubling during the sample period, which could provide a positive impact on exchange rate pass-through to domestic prices.

Figure 6: Openness and import penetration ratio



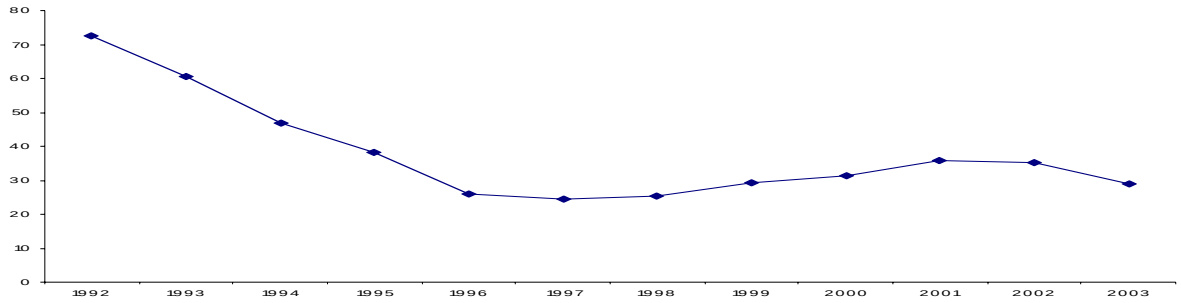
6.2 Reduction in import tariffs and removal of trade restrictions

Frankel et al (2005) argue that any theory of incomplete pass-through must posit some barrier to arbitrage of a good between country of origin and the country of purchase. They empirically find that tariffs, which are one such trade barrier, have a negative pass-through impact, particularly on the CPI of developing countries. The negative relationship between tariffs and pass-through could also arise for the following reason. Under PCP, as tariffs are

[15] We call these factors possible as they are not direct tests to explain the rise in pass-through. The required data at monthly frequencies for such a direct test are not available.

levied on the exchange rate change itself, higher tariffs would impact domestic prices more than lower ones. For instance, other things remaining the same, a 10% depreciation with a 50% tariff rate would increase domestic prices by 5%, while it would be 1% at the 10% tariff rate. Given a target price level in the domestic market to maintain its market share, a firm can pass through a much higher percentage of the depreciation at the lower tariff than at the higher tariff. In India, with economic reforms there has been a substantial reduction in tariff rates. Figure 7 shows that the weighted average import duty declined from 72.5% in 1992 to 25% in 1997, though the trend reversed somewhat during 1997 to 2001 and declined again. This fall in the level of tariffs could also restrict the decline in pass-through. Removal of various quantitative trade barriers such as quotas could also have a similar effect.

Figure 7: Weighted average import duty rate

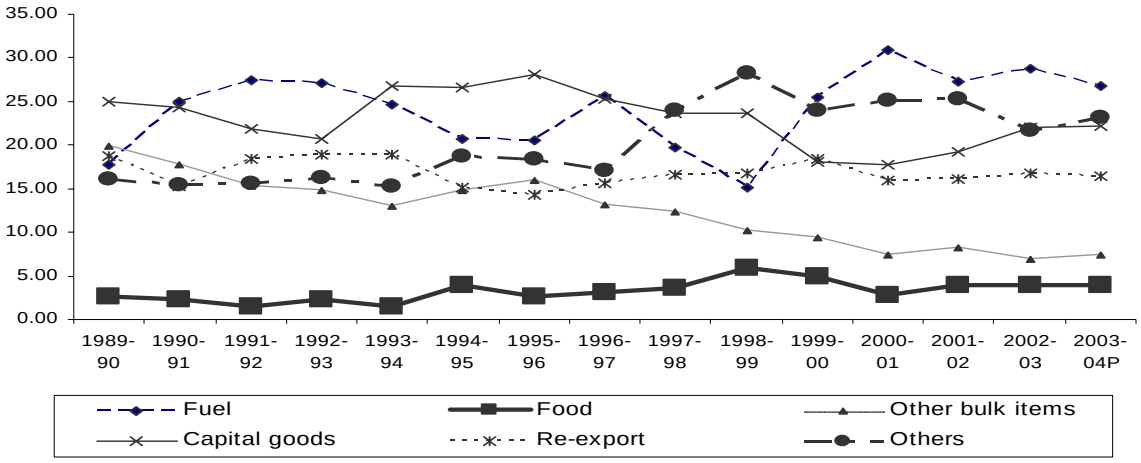


6.3 Import composition

The overall pass-through of exchange rate movements depends upon the composition of imports with various degrees of pass-through. Items such as fuel, food and raw materials have higher pass-through, while that of manufacturing is lower (see Campa and Goldberg (2004) and Otani et al (2003)). Figure 8 shows six components of commodity imports in India since the beginning of the 1990s. Among these six components, the share of energy, food and “others” have increased. An Increase in the share of energy and food would lead to increase in overall pass-through, while that of “others” is not certain as this category mostly comprises miscellaneous products not clearly defined.

The remaining three components showed a declining trend, particularly “other bulk imports”. The bulk imports consist of items which are manufactured but are of the nature of raw materials such as iron and steel, other metals, paper and paper boards, etc. The share of capital goods, which are manufactured but used as industrial inputs, increased initially and then declined with deceleration in the growth of GDP in the second half of the 1990s. As these two components are mostly manufactured intermediate goods, their impact on pass-through is uncertain. Thus, the trend in the share of manufacturing and raw materials including industrial inputs is not clear. However, the rise in the shares of fuel and food could have had a positive impact on overall pass-through.

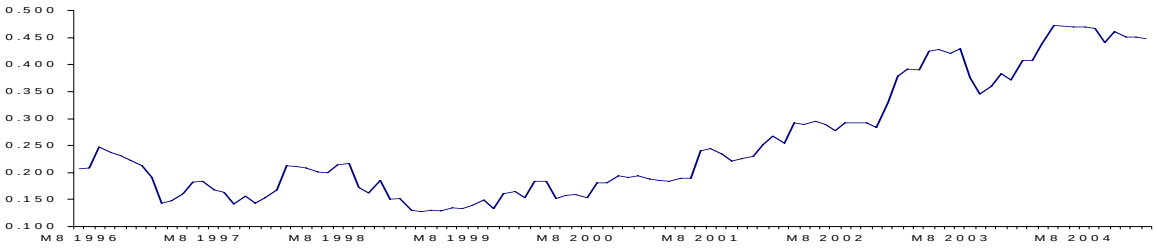
Figure 8: Import composition



6.4 Inflation persistence

An important reason for the non-decline in long-run pass-through when the short-run pass-through tended to decline in the later part of the sample period is the rise in the coefficient of lags of inflation or inflation persistence (Figure 9).

Figure 9: Lagged coefficient (inflation persistence)

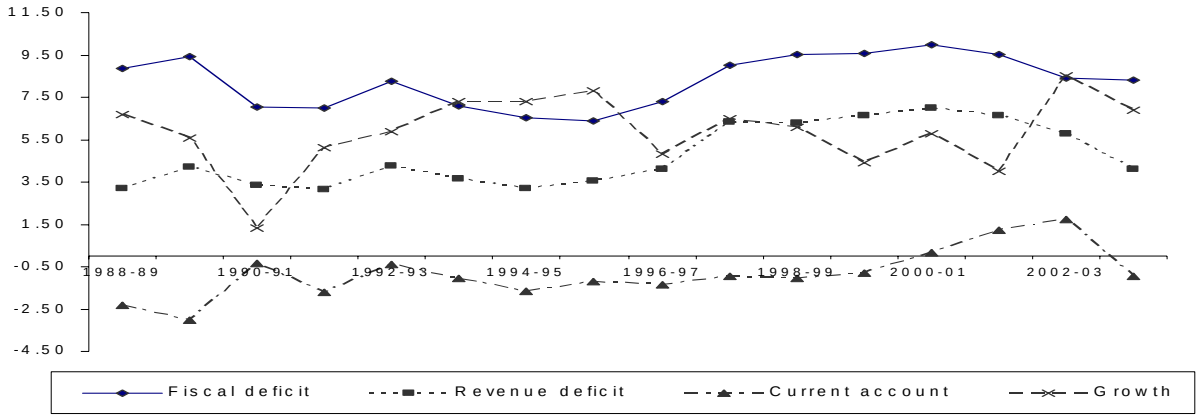


Non-decline of inflation persistence to the extent the actual inflation has fallen is consistent with some of the empirical findings in the literature for industrialised countries and Latin American countries (see, for example, O'Reilly and Whelan (2004)).

In the Indian case so far, there are plausible reasons for inflation persistence to increase. First, as noted from Figure 1, decline in inflation rate was confined to the first half of the 1990s. Since the mid-1990s, the inflation rate hovered around a trend with periodic spurts in some months of 1998-99, 2000-01 and 2004. In other words, inflation has not experienced a declining trend since the mid-1990s, which, combined with periodic spurts in the subsequent period, could raise inflation expectations and persistence. Second, the economy has been emerging out of a price control regime since the second half of the 1990s. The oil price control has been steadily liberalised by reducing the deficit in the oil pool account through reducing subsidies to the oil companies and allowing these companies to pass on the reduction in subsidies to domestic consumers by raising domestic oil prices. Similarly, administrative control on the prices of a number of commodities such as iron, steel and coal has also been liberalised, raising the prices of industrial inputs. These price decontrol measures can raise the level of inflation expectations at least in the transition. Third, the fiscal deficit of the general government (centre and states) after its consolidation and decline during the first half of the 1990s from about 9% to 6% rose since the later part of the 1990s and stood at over 8% with salary revision of public sector employees following the recommendations of the Fifth Pay Commission. The bulk of this deficit was accounted for by the revenue deficit (dis-savings) (Figure 10). In India, the nexus between government deficit

and inflation has been established by a number of studies (see Jadhav (1994) and Rangarajan and Mohanty (1998)). Thus, despite doing away with the system of automatic monetisation of the government deficit and the introduction of an auction system on government securities, due to the high government deficit, the RBI as debt manager continued to subscribe to the primary issue of government securities or injected liquidity into the system to enable market absorption (Khundrakpam and Goyal (2006)). Because of this nexus between the government deficit and inflation in India, a rise in government deficit could raise the inflation expectations of the market participants. Fourth, wage revision in the public sector would affect the wage structure of other organised sectors, leading to an upward revision of inflation expectations in the economy.

Figure 10: Growth, government deficit and current account balance



Yet, actual inflation, despite the periodic spurts, had been kept within bounds due to several factors. First, a major part of the period since the 1990s has seen a benign world inflationary situation translating into lower domestic inflation in several countries including India. Second, the pressure from the demand side eased due to lower than trend growth in output during most of the period since the second half of the 1990s. As against the average real GDP (factor cost) growth of 6.2% during 1992-93 to 2004-05, the average growth during 1997-98 to 2002-03 was 5.3%. Third, demand conditions also eased due to rising current account surpluses since 2001-02, indicating a lack of investment at a time when the overall savings rate in the economy was lower due to higher public sector dis-savings (Figure 10).[16] Fourth, supply management was facilitated by large food stocks, which could be released into the open market and through the public distribution system to keep prices in check. Fifth, allowing the free import of commodities such as sugar, edible oils and cotton, which had been restricted since the mid-1990s, also improved the supply position. Sixth, recent periods have seen more appreciation (Figure 2) which has the effect of lowering domestic prices, in contrast to the depreciation that normally induces a higher price.

7. Concluding remarks

The paper investigates the exchange rate pass-through to domestic prices in India during the post-economic reforms initiated since the major devaluation of July 1991. It finds that, unlike that which has been observed for several other countries, there is no clear-cut evidence of a decline in exchange rate pass-through to domestic prices in India during the post-economic reform period so far. Further, it observes that there is asymmetry in pass-through between

[16] The current account balance, however, has reversed its trend in the most recent period.

appreciation and depreciation, and between large and small exchange rate changes. Drawing on the empirical evidence provided in the literature, it conjectures that non-decline in pass-through in spite of a sharp permanent drop in inflation in the first half of the 1990s could be due to one-time factors associated with economic liberalisation such as a rise in the proportion of imports and exports in income and consumption, reductions in tariffs and the removal of quantitative restrictions on trade and the changing composition of imports towards fuel and food. It also finds a rise in inflation persistence since the second half of the 1990s to be an important cause for non-decline in long-run pass-through. The dismantling of price controls on a number of commodities can lead to an upward revision of inflation expectations, at least temporarily, particularly when inflation hovers around a trend with periodic spurts. Given the nexus between the government deficit and inflation in India, a rise in the government deficit since the latter part of the 1990s could have also led to a rise in inflation expectations.

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Appendix A

Unit root properties

All the series were tested for their unit root properties using ADF and PP tests. The results presented in Table A1 show that all the series are non-stationary at levels barring Y, which is inconclusive. The Zivot-Andrews (1992) unit root test, which allows for a single break at an unknown point in time, indicates that only '*P**' has a structural break, while the rest of the series do not. It implies '*P**' is stationary/non-stationary in either the pre-break or post-break period.

Table A1

Unit root test

(1990:1 to 2005:3)

	Level		**First difference**		**Z–A**	**Order**
	ADF	**PP**	**ADF**	**PP**		
Y	–2.44(t)	–3.56(t)**	–13.5*	–31.1*	NB	I(1)
e	–2.42	–2.52	–15.30	–18.14*(t)	NB	I(1)
P*	–2.57(t)	–2.14(t)	–12.1*(t)	–16.24*	B	NC
P	–2.99(t)	–3.15(t)	–10.48*(t)	–10.35*(t)	NB	I(1)

Notes: 1) Wherever it is statistically significant, a trend component is included denoted by t in parentheses; 2) Lag lengths are selected based on SBC, and the PP test is with Newey-West using Bartlett Kernel; 3) * and ** denote significance at 1% and 5%, respectively; 4) NB stands for no break and B for break; 5) NC stands for not conclusive.

Cointegration tests

Even though the series are not necessarily integrated of the same order, to ascertain the correct form of the series to be considered in the estimation, two residual-based cointegration tests among the variables were performed. The results are presented in Table A2. The Engle-Granger test indicates no cointegration in any of the models. Gregory-Hansen (1996), which allows for a structural break at an unknown point in time under three models (1) level break, 2) full structural break, and 3) a trend with a level break) also indicates no cointegrating relationships.

Table A2

Cointegration tests

Variables	Engle-Granger	Gregory-Hansen		
		Level break	Full structural break	Trend with level break
P, e, P*, Y	–2.84 (–4.16)	–3.98 (–5.28)	–5.35 (–6.00)	4.72 (–5.57)

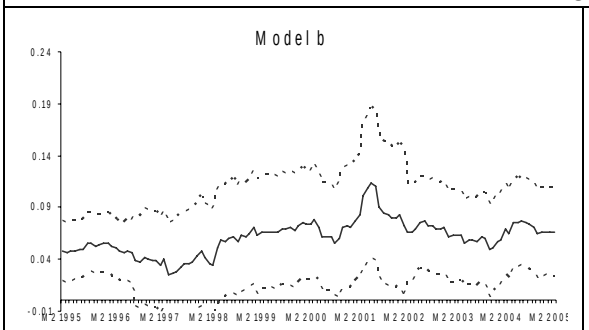
Note: Figures in parentheses are critical values at 5%.

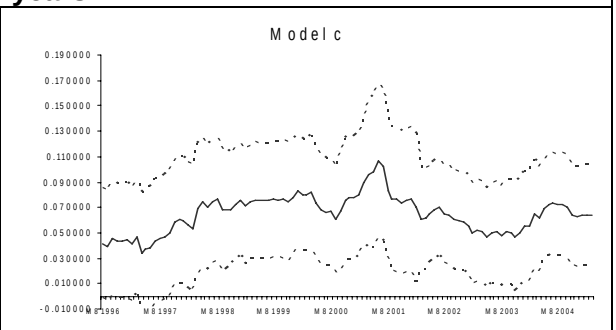
Appendix B

To confirm that the rolling regression results are robust, irrespective of window sizes and the alternative models, we perform the rolling regressions on three window sizes, viz, five, six and seven years on the three specifications: model B and model C. The summary results are presented in Figures B1 and B2 and Tables B1 and B2. From the figures, it is apparent that there are no large differences in the trend of the estimated pass-through coefficients among the models. Between the different window sizes, however, there appear to be some differences, but they broadly indicate a similar trend.

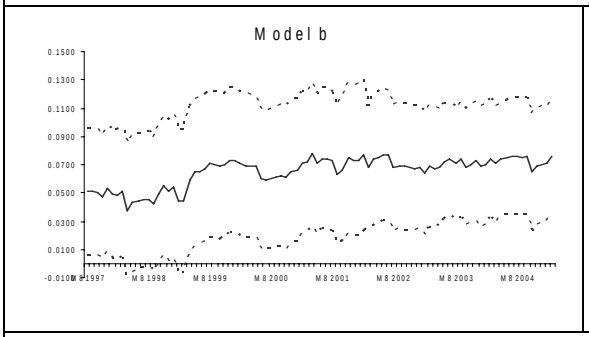
Figure B1: Short-run pass-through

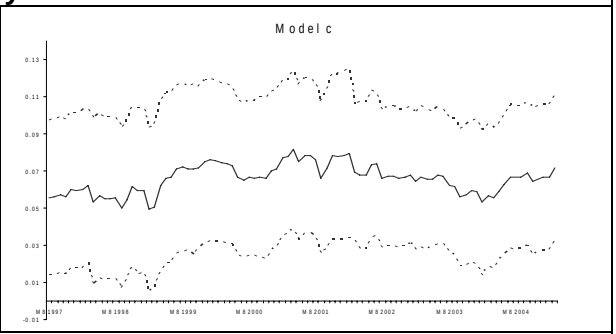
Five years





Six years





Seven years

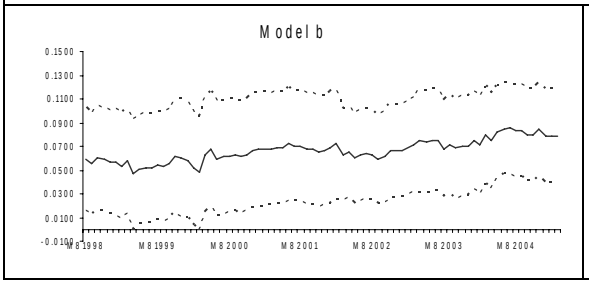
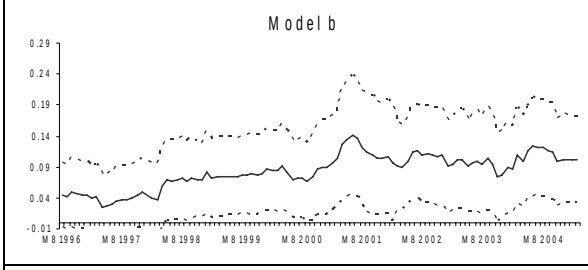
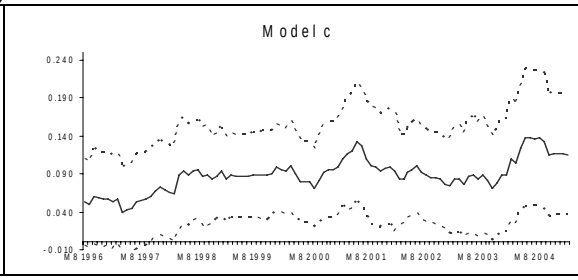


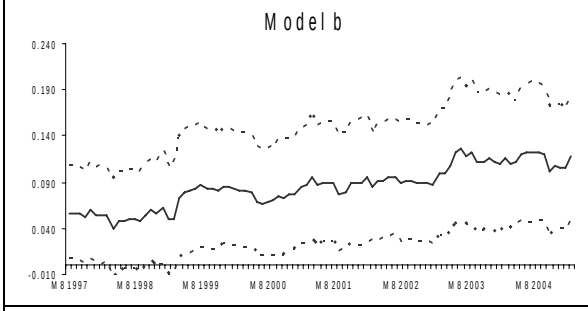
Figure B2: Long-run pass-through

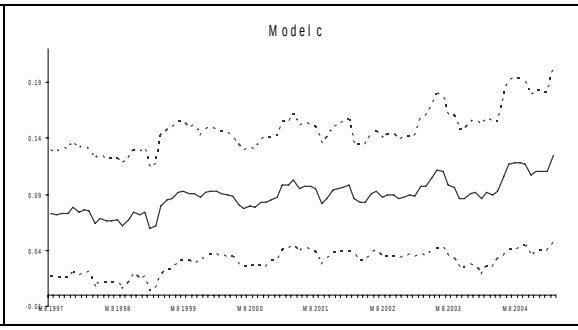
Five years



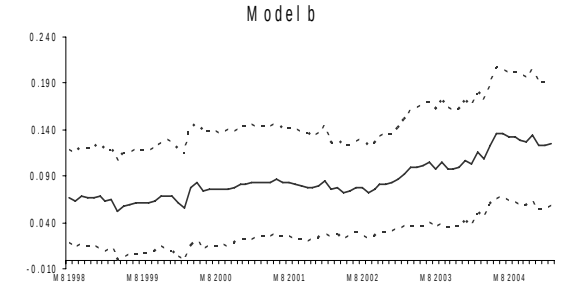


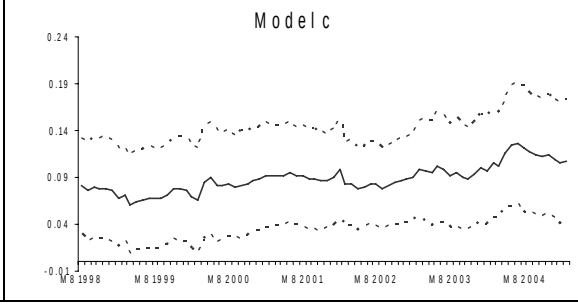
Six years





Seven years





The estimates of the trend fits on the rolling regression coefficients are presented in Tables B1 and B2. Irrespective of the window size and model, the pass-through coefficients do not show a falling trend.

Table B1

Simple trend fit on rolling regression coefficients and Wald test

Models	Five years	R-bar^2	Six years	R-bar^2	Seven years	R-bar^2
Short-run						
Model B	0.00029	0.27	0.00029	0.56	0.00034	0.74
	[0.00]		[0.00]		[0.00]	
Model C	0.00009	0.03	0.00006	0.04	0.00018	0.38
	[0.06]		[0.04]		[0.00]	
Long-run						
Model B	0.00074	0.66	0.0007	0.84	0.00083	0.78
	[0.00]		[0.00]		[0.00]	
Model C	0.0005	0.47	0.0004	0.62	0.0005	0.69
	[0.00]		[0.0]		[0.00]	

Note: Figures in brackets are p-values of the Wald test on the null hypothesis that the estimated trend coefficients are equal to zero.

Table B2

Trend fit with dummies on rolling regression coefficients and Wald test

	Model B						**Model C**					
	Short			**Long**			**Short**			**Long**		
Parameters	**Five years**	**Six years**	**Seven years**	**Five years**	**Six years**	**Seven years**	**Five years**	**Six years**	**Seven years**	**Five years**	**Six years**	**Seven years**
β_0	0.014	0.053	0.033	0.015	0.03	0.034	0.026	0.04	0.31	0.038	0.058	0.043
	(4.2)	(7.7)	(10.3)	(3.1)	(9.2)	(6.7)	(7.8)	(15.5)	(9.1)	(8.5)	(19.7)	(8.2)
β_1	0.0009	–0.00013	–0.00044	0.001	0.0007	0.0006	0.00083	0.00046	0.00051	0.00086	0.0004	0.00059
	(15.9)	(–0.75)	(8.8)	(10.7)	(15.0)	(7.6)	(12.8)	(10.9)	(9.5)	(10.7)	(9.8)	(7.3)
$\beta_2+\beta_0$	0.07	0.062	0.0071	0.134	0.164	–0.089	0.075	0.079	0.0056	–0.038	–0.139	–0.017
	(5.0)	(28.2)	(1.1)	(11.6)	(5.5)	(–8.6)	(6.7)	(9.8)	(8.0)	(–1.9)	(–3.5)	(–1.6)
$\beta_1+\beta_3$	–0.00003	0.00009	0.00063	–0.0003	–0.00044	0.0018	–0.00013	–0.00014	0.00012	0.0013	0.0021	0.0011
	(–0.2)	(36.2)	(101.6)	(–2.5)	(–1.6)	(18.5)	(–1.2)	(–1.8)	(1.9)	(6.6)	(6.2)	(10.9)
R-bar^2	0.71	0.82	0.81	0.82	0.87	0.91	0.62	0.57	0.60	0.65	0.70	0.79
Wald test $\beta_1+\beta_3 = 0$	253	0.57	77.8	114	224	57	165	118	91	115	96	53
	[0.00]	[0.45]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
Wald test $\beta_1=0$	0.07	13.7	103	6.34	2.64	344	1.36	3.2	3.5	44.4	38	118
	[0.79]	[0.00]	[0.00]	[0.01]	[0.10]	[0.00]	[0.24]	[0.07]	[0.06]	[0.00]	[0.00]	[0.00]
Break at window ending	2002:02	1999:04	2002:03	2001:03	2003:06	2002:03	2001:09	2002:03	2002:03	2002:03	2003:06	2002:03

Note: Figures in brackets are p-values of the Wald test on the null hypothesis that the estimated trend coefficients are equal to zero.