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INFLATION AND THE DISPERSION OF RELATIVE PRICES: A CASE FOR 4 % SOLUTION

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Unlike earlier literature that documented positive association between inflation and the dispersion of relative prices *over time*, the empirical evidence from this study suggests that the relative price dispersion increases in response to the deviation of inflation from certain threshold/target level in either direction rather than inflation per se. The striking feature of the empirical evidence from United States and Japan is that the inflation rate at which the dispersion of relative prices is minimised turn out to be 4%; hence, supporting the proposal of 4% inflation target for both the countries.

JEL codes: E30, E31, E52.

I. Introduction

The impact of inflation on the dispersion of relative prices is considered as the central channel through which inflation affects the real sector of an economy. In new Keynesian dynamic general equilibrium models, the larger dispersion in relative prices due to inflation is considered to be the root cause of all the distortionary real effects of inflation (Green, 2005). The findings of these models advocate that monetary authorities must be committed to ensure price stability mainly because changes in aggregate prices cause larger variability in relative prices (see e.g., Woodford, 2003; Becker & Nautz, 2012).

Although vague, the effect of inflation on the structure of relative prices is a primary channel through which negative effects of inflation are transmitted to the real sector of economy (Ball &

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Romer, 2003). In this context, Blejer and Leiderman (1980) emphasised four channels through which higher variability in relative prices affect the growth rate of real output: (i) it reduces the information content of relative prices and drives a wedge between marginal rates of transformation and substitution, which results in misallocation of resources and inefficiency; (ii) it leads to large search activities which involves costs in terms of time and resources; (iii) it results in shortening of optimal contract length and more frequent revisions in existing contracts, thereby increasing the cost of contracting; and (iv) it impedes the efficient allocation of resources to the extent that the costs of higher price variability are differentially distributed among firms.

In literature, a large body of theoretical studies have demonstrated that the higher rate of inflation results in higher variability of relative prices and inflation uncertainty. Theoretically, the relationship between inflation and relative price variability is explained by models based on misperceptions or incomplete information (Lucas, 1973; Hercowitz, 1981; Cukierman, 1983). The models based on incomplete information demonstrate that under certain conditions the firms with price elastic supply adjust quantity in response to demand shocks whereas the firms whose supply is inelastic adjust prices in response to such shocks. Hence, demand shocks that generate higher inflation trigger larger variability or dispersion in relative prices.

The association between inflation and the inflation uncertainty/variability, highlighted by Friedman (1977), arises from the public's perception about the erratic policy responses by monetary authority to large changes in prices (Ball, 1992). More precisely, this relationship is built on the premise that sharp rise in inflation produces strong pressure on the policy makers and as a result, 'policy goes from one direction to other', thereby, generating wide variation in actual as well as expected inflation.[1] Under these circumstances, predicting both the longer term drift in inflation and its short term movements become more difficult.

Extending this strand of literature, Shoesmith (2000) points out that as inflation rises 'component inflation measures are not only more variable at each point in time, but also more dispersed *over time*'. He argued that component price indices are likely to drift apart more significantly during the period of high inflation (as compared to low-inflation period), hence, resulting in larger dispersion of relative prices *over time.* Note that such measure of dispersion is different from the traditional measure which is measured as variance of cross-sectional distribution of price changes at a given point of time.[2] In order to empirically verify this, Shoesmith (2000) examined the number of cointegrating relations among component price indices, as large (less) number of significant cointegrating relations implies less (large) dispersion of relative prices over time. Using Johansen's (1991) methodology, the study finds significant fall in the number of cointegrating relationship among the components of US Consumer Price Index (CPI) during the periods of high inflation.

On the contrary, Scharff (2007) found relatively less number of significant cointegrating relations among component price indices of US CPI during the period of moderate inflation, while using the same sample period as considered by Shoesmith (2000). She attributed this contradicting evidence to the fact that Shoesmith (2000) uses component indices which covers only 70% of the US CPI. She further corroborates her findings with evidences obtained from a set of countries and concludes that counting number of significant cointegrating relations in a system of component price indices is not an appropriate method to examine the response of relative price dispersion to inflation.

[1] Number of empirical studies provide evidence in favor of the hypothesis that higher rate of inflation is associated with larger inflation uncertainty (see e.g., Daal *et al*., 2005; Thornton, 2008; Herwartz & Hartmann, 2012).
[2] From the theoretical point of view, the measure of relative price dispersion measured *over time* highlights another aspect of variability associated with inflation.

In this context, we argue that the traditional approach of bifurcating the sample into high and low inflation regimes, and comparing the number of significant cointegrating relationships among component price indices across different regimes is not an appropriate approach. In particular, under such approach, the dispersion of relative prices is related to the rate of inflation per se. However, theoretically, the relative price dispersion over time is perceived to increase in response to the deviation of inflation from certain threshold/target level in either direction rather than inflation per se. That is the deviation of inflation from certain threshold level in either direction tends to increase the dispersion of relative prices over time.

A number of recent theoretical and empirical studies have demonstrated that the strength of association between inflation and relative price dispersion is largely determined by the magnitude of deviation of inflation from certain threshold level in either direction (Fielding & Mizen, 2008; Choi, 2010; Choi & Kim, 2010; Rather *et al.*, 2014). In particular, Fielding and Mizen (2008) and Choi (2010) have shown that the response of relative price dispersion to inflation is U-shaped; implying that larger deviation of inflation will result in higher dispersion of relative prices. This finding suggests that in a system of price indices, the number of cointegrating relations is expected to decline (which implies rise in price dispersion) as inflation departs far off from the threshold level in either direction. In other words, the number of cointegrating relations is likely to be less (i.e., the number of stochastic trends is expected to be more) during the periods of high as well as low inflation and maximum when inflation is closer to its threshold level.

In this paper, using the CPI data from United States and Japan, we first calculate the number of significant cointegrating relations in a system of component price indices for a sequence of rolling subsamples. Next, we compare the number of significant cointegrating relations obtained from each rolling subsample with the average inflation rate of the corresponding period. The advantage of this procedure is that it provides scope for capturing the dynamic relationship between inflation and the dispersion of relative prices. In contrary to the findings of earlier studies, the empirical results from this study indicate that inflation per se does not affect the dispersion of relative prices measured over time. Instead, this dispersion seems to be rising in response to the deviation of inflation from certain threshold level in either direction. The dispersion of relative prices tends to decline as inflation approaches the threshold level from either direction. The crucial inference that emerges from the empirical evidence is that the inflation rate at which the dispersion of relative prices is minimised turns out to be 4%; hence, supporting the proposition of 4% inflation target for United States and Japan. The rest of the paper is organised as follows: section II. provides various theoretical paradigms and the empirical evidence; section III. provides the methodology; section IV. discusses the data and empirical results; and section V. provides the concluding remarks.

II. Theoretical Paradigms and Empirical Evidence

The association between the inflation and inflation uncertainty follows from Friedman's (1977) discussion on government policy, inflation and uncertainty that 'A burst of inflation produces strong pressure to counter it. Policy goes from one direction to the other, encouraging wide variation in the actual and anticipated rate of inflation. And, of course, in such an environment, no one has single-valued anticipations. Everyone recognises that there is great uncertainty about what actual inflation will be over any specific future interval'. This resulted in great debate among the economists investigating the relationship between inflation and inflation uncertainty. Building on the seminal work of Friedman (1977) and Ball (1992) formalises a model with two types of policymakers, tough and soft, who stochastically alternate in power. The public knows that only the tough type is willing to bear the economic costs of disinflation. When inflation is already low, both types of policy makers will try to keep it low, and uncertainty about future inflation will be low. However, when inflation is high, the public can only speculate about how long it will take until a tough type is in power to bring inflation rate down. Thus under these situations the uncertainty about future inflation will therefore be higher. In literature, most of empirical studies, supporting the association between inflation and inflation uncertainty, have used a measure of uncertainty that is constructed from a measure of inflation using GARCH models (see e.g., Daal *et al*., 2005; Thornton, 2008; Herwartz & Hartmann, 2012). Extending this strand of literature, Shoesmith (2000) argued that one important aspect of this relationship is that the individual prices may diverge more or less over time, depending on the rate of inflation. In particular, price indices may drift apart more during the periods with high inflation rate than during the low inflation period in a more dynamic manner than generally understood.

On the other hand, the link between inflation and relative price variability is explained by models based on misperceptions or incomplete information (Lucas, 1973; Barro, 1976; Hercowitz, 1981; Cukierman, 1983) and the menu cost models of price adjustment (Sheshinsky & Weiss, 1977; Ball & Mankiw, 1994, 1995). Under information asymmetry, firms adjust quantity in response to unanticipated demand shocks if supply is price elastic and adjust prices if supply is price inelastic. Hence, unanticipated demand shocks that generate inflation tend to affect the variability of relative prices (Lucas, 1973; Barro, 1976). The menu cost model developed by Sheshinski and Weiss (1977) suggests that firms' price adjustment is heterogeneous as it incurs menu cost that varies across firms and thus, shocks that lead to inflation tend to increase the dispersion of relative prices. Moreover, in the presence of trend inflation shocks that raise firms' desired prices trigger larger adjustment in prices as compared with shocks that lower firms' desired prices (Ball and Mankiw, 1994). In particular, when a firm is required to lower its price in response to a shock that reduces its desired price, it is more likely that it may not respond, because a higher rate of inflation tend to reduce its relative price over time. Thus, such asymmetry in price adjustment of firms due to inflation results in higher variability of relative prices.

Following the seminal paper by Parks (1978), number of studies in literature empirically examined the relationship between inflation and variability of relative prices (see among others, Debelle & Lament, 1996; Nath, 2004; Miszler & Nautz, 2004; Fielding & Mizen, 2008; Nautz & Scharff, 2012). All these studies, while supporting this relationship, used a measure of relative price variability which is estimated as variance of distribution of individual commodity inflation rate in any 'single period of time'. Note that the measure of relative price variability constructed so does not capture the dispersion of relative prices measured 'over time' as highlighted by Shoessmith (2000). It is important that if the prices of individual commodities or commodity groups diverge more during periods of high/low inflation then Friedman's argument, regarding the loss of economic efficiency, is intensified by this additional form of uncertainty. That is not only are disaggregate measures of inflation more unpredictable about the mean at various points in time, but even more so as they significantly drift apart from one another over time.

III. The Methodology

There are plenty of empirical studies which have examined the divergence among a set of variables by counting the number of significant cointegrating vectors using Johansen (1991, 1995) cointegration test (Bernard & Steven, 1995; Siklos & Mark, 1997; Rangvid, 2001; Pascual, 2003; Mylonidis & Christos, 2010). Under this procedure, finding a single common stochastic trend among a vector of n variables, that is, finding the *n-1* number of significant cointegrating relationships implies larger convergence. On the contrary, as the number of common stochastic trends among the variables increases, the divergence among them tends to increase.[3] Thus, an increasing number of significant cointegrating relationships would constitute evidence of decreasing divergence and vice-versa (Rangvid, 2001).

To evaluate the dispersion of relative prices *over time*, we employ Johansen (1991, 1995) cointegration test to a system of component price indices for a sequence of rolling samples. Unlike, the conventional approach of simply comparing the cointegration results obtained from high- and low-inflation period, the advantage of rolling cointegration test is that it allows examining the variation in the number of cointegrating relations over time. This approach helps to understand the dynamic relationship between inflation and the dispersion of relative prices in a better manner than the procedure of simply dividing the sample into high and low inflation regime. More importantly, given the possible sensitivity of cointegration results to sample selection (Hansen & Johansen, 1999; Johansen *et al.*, 2000), the simple procedure of dividing the sample period into low and high inflation regime and comparing the cointegrating results may not be appropriate.

In the first stage, we estimated the number of significant cointegrating vectors among component price indices for a sequence of rolling subsamples. In the second stage, we compare the rank 'r' (number of significant cointegrating vectors) obtained from each rolling sample with the corresponding period average inflation rate. Here, the presence of larger (smaller) number of significant cointegrating vectors (i.e., presence of lesser number of common stochastic trends) implies lesser (wider) dispersion of relative prices over time.

Johansen (1995) demonstrates that the test procedure is unbiased if the rank tests are interpreted as a sequence. Starting from rank zero, the test procedure stops at the first insignificant test statistic. The procedure involves investigation of the k-dimensional vector autoregressive process of pth order:

$$\Delta y_t = \mu + \pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + \varepsilon_t$$

where Δ is the first-difference operator, y_t is a $(k \times 1)$ random vector of k time series variables with order of integration equal to one. μ is a $(k \times 1)$ vector of constants, π is $(k \times k)$ matrix of parameters and ε_t is a sequence of zero-mean k-dimensional white noise vectors. In this framework, the rank *(r)* of the $(k \times k)$ matrix of π provides information about the long-run relationships among the variables. The rank of π given by $0 \prec r \prec k$ implies the existence of r cointegrating vectors among the variables. Thus, there exist $(k \times r)$ matrices of α and β, each of rank r such that $\pi = \alpha\beta'$. The testing of hypothesis that the number of cointegrating vectors is at most r (where, $r = 1, 2\ldots\ldots, \ k - 1$) is conducted by using both the maximum eigenvalue and trace statistics.[4]

[3] Mylonidis and Christos (2010) demonstrate how decreasing number of cointegrating relationships or increasing number of common stochastic trends implies higher divergence among a set of variables.

[4] An alternative approach to examine the dispersion of component price indices is to compare the magnitude of adjustment coefficients obtained from the error correction framework. However, in the absence of theoretically plausible restrictions, the identification of behavioural cointegrating vectors becomes difficult in a given system. In such situations, generally, an *ad hoc* restriction is imposed to capture the errors which are mutually exclusive. Under such circumstances, the results based on the magnitude of the adjustment coefficients, obtained by imposing the ad hoc restrictions, will quite often result in misleading inferences (Kennedy, 1998; Patterson, 2000).
However, in a system where we can exactly identify the behavioural cointegrating vectors, a better method to examine the convergence would be to estimate the magnitude of the adjustment coefficients rather than

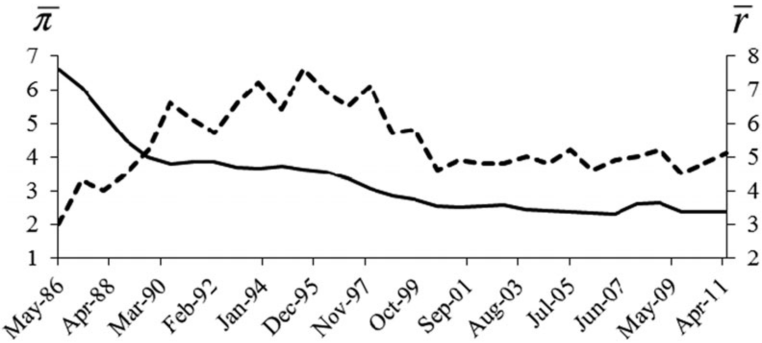

Figure 1. Average Inflation and number of cointegrating vectors (United States)

IV. Data and Empirical Results

We have used CPI data from United States and Japan as used by Scharff (2007) and Shoesmith (2000) to conduct the empirical analysis. The common feature of both these countries is that they have experienced similar evolution of inflation over time. We used monthly data on 10 major component indices of US CPI for the sample spanning from January 1978 to June 2014. These sub-price categories include: food away from home, footwear, fruits and vegetables, fuels and utilities, meat poultry fish and eggs, medical care, men's and boys' apparel, private transportation, shelter and women's and girls' apparel. Similarly, for Japan, we used monthly data on 10 major sub-price categories of the CPI basket and the sample period ranges from January 1970 to December 2014. The component price indices for Japan include: housing, fuel light and water, furniture and household utensils, cloths and footwear, medical care, transportation and communication, education, reading and recreation and miscellaneous.

Firstly, we examined the time series properties of the various component price indices using the conventional unit root test during each rolling subsample. The results from the conventional unit root tests confirmed that the time series of all the component price indices (in logarithmic form) follow I(1) process in each rolling subsample. Next, for both United States and Japan, we calculate the number of significant cointegrating vectors in a system of component price indices for a sequence of rolling sub samples over the overall sample period.[5] To this end, Johansen's cointegration test was conducted for a sequence of rolling samples having a window size of 100.

To examine the relationship between inflation and the dispersion of relative prices, we compare the number of significant cointegrating relationships obtained from each period with the corresponding period's average inflation rate. In Figure 1, we plot the average number of significant cointegrating vectors ($\overline{r}$) (the dotted line in the Figure) against the corresponding period average inflation rate ($\overline{\pi}$) (the thick line) over time for United States.[6] Similarly, in case of Japan, the average number of significant cointegrating vectors ($\overline{r}$) is plotted against the corresponding period

counting the number of cointegrating vectors. Note that for the data set used in the present study, there is no theoretically plausible restriction available, thereby making the identification of behavioural cointegrating vectors difficult.

[5] These results from cointegration tests are not presented here and can be obtained from the author upon request.

[6] As the rolling subsamples overlap, we present the average number of cointegrating vectors measured over consecutive 10 windows and the corresponding period average inflation in all the figures.

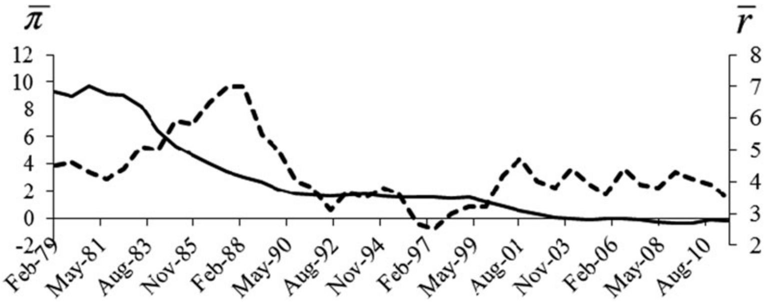

Figure 2. Average Inflation and number of cointegrating vectors (Japan)

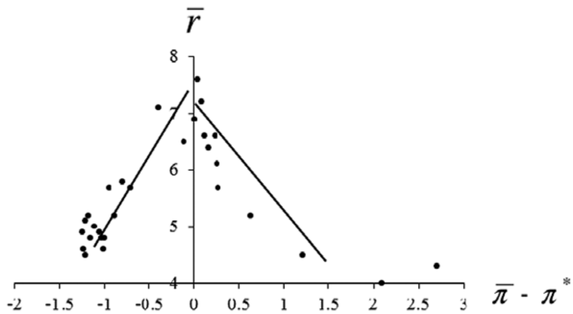

Figure 3. Inflation deviation and number of cointegrating relations (r) for United States

average inflation rate ($\bar{\pi}$) over time in Figure 2. It can be clearly seen that for both the countries initially the number of cointegrating vectors increase as inflation falls up to a certain level (4%), and subsequently as inflation falls further to lower levels the number of cointegrating vectors start decreasing. In other words, the results suggest that the number of significant cointegrating relations is comparatively less during the periods of very high or very low inflation. These results indicate that the dispersion of relative price *over time* is minimised (i.e., existence of lesser number of common stochastic trends or higher number of r) as inflation stabilises around 4%. A crucial implication of these findings is the presence of a threshold inflation level at which the dispersion of relative prices is minimised.[7]

To trace out this relationship more clearly, we sorted the time series of both inflation and the number of cointegrating vectors with respect to inflation in ascending order. Then based on the results from Figures 1 and 2, we calculated the deviation of inflation from its threshold level

[7] It is important to note that a certain proportion of dispersion in relative prices over time may be due to changes in real factors such as real income, preferences, technology, and so forth, The variability in relative prices due to such factors is crucial for efficient allocation of resources as it reflects signals purely from market and real sector of an economy. In this context, Becker and Nautz (2012) and Head and Kumar (2005) demonstrate that the rate of inflation that minimizes variability of relative prices can serve as a proxy for the threshold inflation rate that minimizes the welfare cost of inflation. Moreover, number of empirical studies provides evidence for United States and many other countries that more than half of variability in relative prices is mainly due to inflation. In particular, Ram (1990) and Rather *et al.* (2014) demonstrate that the component of relative price variability due to inflation increases as inflation rises.

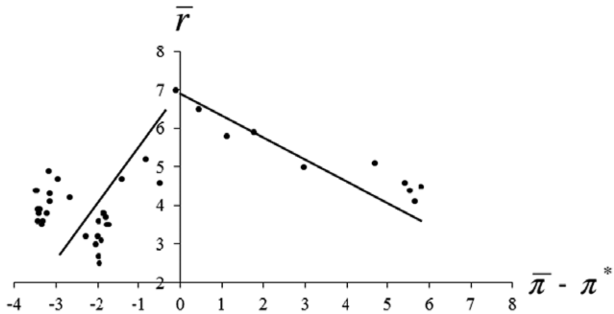

Figure 4. Inflation deviation and number of cointegrating relations (r) for Japan

$(\bar{\pi}-\pi^*)$for each period considering $\pi^* = 4\%$. In Figure 3, we present the scatter plots of average number of significant cointegrating vectors *($\bar{r}$)* against the deviation of average inflation rate from its threshold level $(\bar{\pi}-\pi^*)$ for United States. Here, the vertical axis measures the average number of significant cointegrating vectors *($\bar{r}$)* and the horizontal axis measures the corresponding period's inflation deviation$(\bar{\pi}-\pi^*)$.[8]

It is evident from Figure 3 that the number of significant cointegrating relations among component price indices is comparatively lesser when inflation deviates highly from its threshold level in either direction. The number of cointegrating relationships is found to be maximum when the deviation of inflation from its threshold level is zero. These results indicate that the dispersion in relative prices tends to decline as inflation approaches certain threshold level. This finding is consistent with the U-shaped relationship between inflation and relative price variability documented in the recent literature (Fielding & Mizen, 2008; Choi & Kim, 2010; Nautz & Scharff, 2012; Rather *et al*., 2014). More importantly, the results support the view that the dispersion of relative prices *over time* changes not with the inflation rate per se as documented by Shoesmith (2000) and Scharff (2007), but with the deviation of inflation from its threshold level.

The results for Japan are presented in Figure 4. Similar to what we have observed for United States, the scatter plot exhibits an inverted U-shape pattern indicating that there exists less number of cointegrating relations (i.e., more dispersion of relative prices) during periods of very low and very high inflation. In other words, the farther away a shock drives inflation from π^* in either direction, the more dispersed are relative prices over time. In case of Japan too, the number of cointegrating relations is found to be maximum when deviation of inflation from its threshold level is zero or inflation is around 4%. Over all, these evidences suggest that for inflation levels below the threshold level, rise in inflation tends to reduce dispersion in relative prices and on the contrary, for inflation levels above the threshold level, an increase in rate of inflation results in larger relative price dispersion.[9]

The striking feature of the empirical evidence is that the dispersion of relative prices *over time* is minimised when average inflation is around 4% for both United States and Japan. Interestingly, this finding substantiates the argument of Ball (2013, 2014) and Blanchard *et al.* (2010) in favour of raising the inflation target to 4%. Although their argument is based on the rationale that raising the inflation target would ease the constraints on monetary policy arising from zero bound on

[8] Expanding the size of rolling window, however, does not significantly alter the results.
[9] No significant change in the inferences was found when we used the sample periods considered by Shoesmith (2000) and Scharff (2007).

interest rate, the target inflation rate they propose is the same as we obtained here. On the same line, Krugman (2013) argues that a higher baseline for inflation, say 4%, than a conventional target of 2% makes liquidity trap less likely to occur and less costly even when it occurs. Hence, stabilizing the inflation around 4% will not only minimise the dispersion of relative prices but also ease the constraints/costs arising from zero bound and the liquidity trap.

Overall, the empirical results indicate that the component price indices drift apart considerably as inflation deviates farther away from certain threshold level; implying that both very low and very high inflation generate larger relative price dispersion *over time.* The results further confirm that what matters for the response of relative price dispersion is not the inflation per se as recognised in the earlier literature, but the deviation of inflation from certain threshold level.

V. Concluding Remarks

This paper examines the relationship between inflation and the dispersion of relative prices *over time* using the CPI data from United States and Japan. To measure the dispersion of relative prices, we count the number of significant cointegration relations in a system of component price indices for a sequence of rolling samples. Under this procedure, a lesser number of significant cointegrating vectors (i.e., presence of more number of common stochastic trends) implies larger dispersion among prices and vice-versa. The empirical results, unlike the earlier studies, indicate that inflation per se does not affect the dispersion of relative prices. In fact, the dispersion of relative prices seems to be rising in response to the deviation of inflation from certain threshold level in either direction. In other words, the dispersion of relative prices tends to decline as inflation approaches to the threshold level from either direction. For both United States and Japan, the empirical results confirm the sub-price indices drift apart considerably during periods of high as well as low inflation, thereby implying that both high and low rate of inflation generates larger relative price dispersion.

From the policy perspective it is important to note that unlike the dispersion of relative prices due to changes in real factors (such as income, family composition, tastes and technology), the dispersion in relative prices generated through inflation is distortionary as it is believed to reduce the information content of nominal prices, which in turn leads to misallocation and welfare loss. This effect of inflation on the structure of relative prices is considered a primary channel through which negative effects of inflation are transmitted to the real sector of economy. Thus stabilizing the rate of inflation around certain threshold level will minimise such inflation induced distortions in relative prices in an economy.

The crucial inference that emerges from the empirical analysis is that the inflation rate at which the relative price dispersion is minimised turned out to be 4% for both United States and Japan. These findings support the proposition of 4% inflation target for both the countries.

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