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Inflation, inflation uncertainty and relative price variability in Bangladesh

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Abstract Developing countries are often busy with the rate of inflation and its effects on the economy. Although the monetary policy of developing countries is concerned with business fluctuations and its effects on stability, recent studies are giving importance on the relationship between inflation and relative price variability (RPV). In recent macroeconomic theory, RPV generates the fundamental distortions of inflation, which disrupts the informational content of nominal price. It has long been popularly believed that the relationship between RPV and inflation is positive and stable. Using disaggregated monthly CPI data for Bangladesh from 2002:7 to 2013:6, this study tries to tackle the following problems: (1) whether the relationship is linear? (2) whether the relationship is sensitive to the models of inflation forecasting? (3) whether the model is stable? This study finds that the relationship is not linear, which contrasts with the earlier works on RPV-inflation relationship. Our semiparametric estimations show that the relationship is U-shaped. The estimation of the parametric quadratic function shows that the model of inflation forecasting is sensitive to this relationship which makes that it is the unexpected inflation which matters for RPV. Although the equation is specified, but it is not stable over time.

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The rolling regression analysis and breakpoint test show that there is a breakpoint during the sample period. The instability also comes from the food inflationary shock and poor macroeconomic policy management. This stability of the relationship is important for determining the threshold level of inflation, which is crucial to minimize RPV.

Keywords Relative price variability · Inflation uncertainty · Threshold inflation · Structural break

JEL Classification E31 · E37 · E52

1 Introduction

Inflation and its effects on the economy are some of the most discussed issues in macroeconomics. High inflation generates not only resource misallocation, but exacerbates poverty as well. Since June 2010, an additional 44 million people fell below the USD 1.25 poverty line, as a result of higher food prices as in 2008. Low-income and lower-middle-income countries are experiencing, on average, a 5 % points increase in food price inflation compared to better-off countries (Mundial 2011). Therefore, the welfare cost of inflation is a much-debated issue in contemporary economics, although empirical evidence on this is rather weak. According to Lucas (2000), an annual inflation reduction from 10 to 0 % is equivalent to a real income increase of less than 1 % for the United States. While previous studies focus on the negative impact of inflation on aggregate demand, recent studies are giving importance to the inflationary effects on the relative price distribution in the economy. Relative price variability (RPV) contains nominal price information. As such, an increase in inflation distorts that informational content, making inflation costlier.

In the first 3 years of its recovery from the effects of the 1971 war of independence, Bangladesh experienced numerous negative supply shocks, such as droughts, floods, and a sharp rise in the OPEC oil price. In the first 2 years of independence, foreign aid and subsidized loans helped stabilize the economy, but later, the economic situation deteriorated (Hossain 2006). Additionally, shortages of agricultural and industrial raw materials intensified, causing persistent industrial unrest. Moreover, the expansionary fiscal and monetary policy, coupled with a negative supply shock, created an explosive inflationary condition. By mid-1974, the country was affected by a full-fledged famine. Despite the economic and political odds, the socio-economic condition has significantly changed over the past three decades. For example, since the introduction of the import-substitution strategy of development, Bangladesh started following an export-oriented strategy. This transformation increased the average growth rate from 2.31 % in 1972–1981 to 5.71 % in 2002–2011. Currently, Bangladesh is approaching an above 6 % rate of growth (World Bank 2015).

Despite these economic achievements, persistent dependence on foreign aid, frequent natural disasters, political upheavals, systematic corruption, and misgovernance have put economic development at risk. Additionally, the high rate of unemployment and inequality often make the high growth rate questionable (IMF

2007; Hossain 2006). Furthermore, negative supply shocks, and budget and trade deficits also exert upward pressure on the price level and erode the advantages of high growth rates (Ahmed et al. 2013; Joarder et al. 2015). However, economic vulnerability from externalities is moderate, as shown by the fact that the 2008 financial crisis did not exert a large pressure on the domestic sector. Historically, Bangladesh experienced a downward trend in the broad money supply growth rate, but an increasing inflation, and, since 1997, an increase in the consumer price index (CPI). As such, Fig. 1 shows the point-to-point inflation rate from 1996, with an average of 6.54 % and several hikes. At the end of 1998, for example, the inflation rate reached a maximum of 12.71 %. Other peaks exist in 1998, 2004, 2007, 2011, and 2013, with the major reasons being floods, domestic currency depreciation, global food price increases and balance of payment issues, political instability, etc. During this period, Bangladesh also experienced a low rate of inflation in 2000, 2009, and 2012, respectively (marked in blue). Additionally, inflation follows seasonality due to domestic rice production (for instance, *Aus*, *Aman*, and *Borro* production). As such, food has the greatest weightage in the CPI basket, followed by fuel and related components.

If both fiscal and monetary policies are designed to achieve a higher growth rate, inflation may adversely affect growth. Therefore, analyzing the RPV-inflation relationship may help formulate a prudent monetary policy aimed at reducing price variability. Consequently, the nonlinear functional relationship can dissipate the spurious asymmetry of misspecification for the RPV-inflation relationship (Choi and Kim 2010). Therefore, determining the correct functional form is important for establishing the optimal rate of inflation that is assumed to minimize RPV. Moreover, determining the actual relationship is challenging, and even if it is accurately determined, it may not support the theories on the RPV-inflation relationship. The reason is that real data often fail to support linear specification and theoretical propositions. Accordingly, Fig. 2 illustrates the relationship between RPV and inflation in the case of Bangladesh. Panels (a)–(c) show scatter, quadratic, and nonparametric plots for different samples. In most cases, the overall relationship is best described as U-shaped, as RPV initially decreases with higher levels of inflation, but starts to increase when inflation surpasses a certain threshold. The U-shaped relationship implies that the marginal effect of inflation on RPV differs according to the inflation level. For inflation below a threshold level, higher inflation lowers RPV, and the relationship is reversed above the threshold. As previously stated, the functional relationship is important for the monetary policy to determine the optimal rate of inflation. As higher inflation distorts the informational content, it is assumed that decreasing inflation can lower the RPV as well. However, this does not hold if the relationship is non-monotonic. When the operative relationship is non-monotonic, there can be a positive value of inflation for which the RPV is minimized, and reducing inflation is feasible only to a certain extent. Nonetheless, there may be a strictly positive level of inflation that is optimal for reducing price variations around the general inflation rate.

Even if the above issues are addressed, the problem of stability of the functional relationship persists. However, existing literature on the question of stability is rather limited. Most studies assume that the marginal impact of inflation on RPV is time

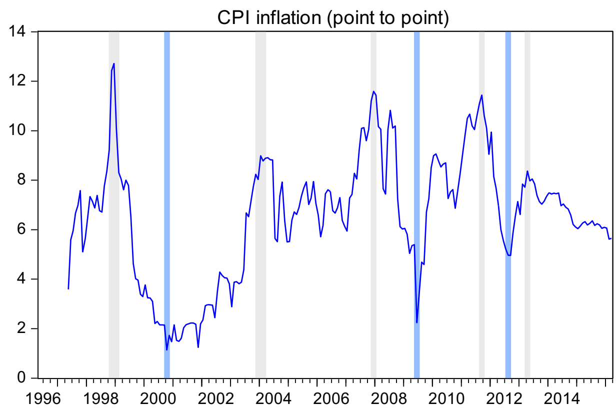
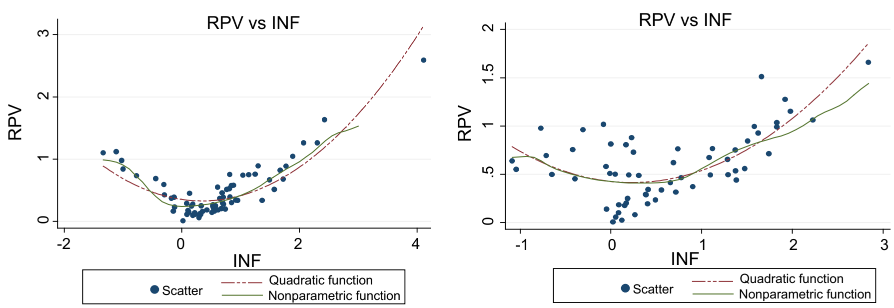


Fig. 1 CPI Inflation of Bangladesh from 1996M05 to 2016M03: *gray shades* are maximum values and *blue shades* are minimum values. *Source*: International Financial Statistics (IFS)



(a): Sample 2002M07-2008M06

(b): Sample 2008M07-2013M06

(c): Sample 2002M07-2013M06

Fig. 2 *Scatterplots*, *functional* fitted plot and *nonparametric* plot of different samples of RPV vs INF in Bangladesh. *Source*: Monthly Economic Trends (Various Issues), Bangladesh Bank and Authors' calculation

invariant. Although the empirical evidence on structural changes of inflation series is substantial, that on the RPV-inflation relationship is not. For example, Dabus (2000) finds that relationship between inflation and RPV in Argentina exhibits structural changes across different levels of inflation, and Choi (2010) shows the variation in the

RPV-inflation relationship across regimes. As previously stated, food prices and policy shocks may affect inflation, but the price level in Bangladesh may not behave as it does in developed countries. While Fischer et al. (1981), and Nautz and Scharff (2005) argue that supply shocks for food and energy prices have important effects on RPV, this may not hold true for countries such as Bangladesh with a distinct consumption basket. Consequently, this feature is not typically studied in Bangladesh. This approach necessitates a main focus of this study on the examination of supply shocks and structural changes of the functional relationship.

Following Fielding and Mizen (2008), and Choi (2010), this paper examines the relationship between RPV and inflation, in an attempt to address the issues presented above. As such, three interrelated aspects of this relationship are examined in this study: (1) the functional form of the relationship, (2) the role of different models of inflation forecasting on the relationship and the relevance of the theories, (3) the stability of the functional relationship and effects of supply shocks. Additionally, we also reexamine previous studies to assess the various aspects of inflation on RPV. To our knowledge, this is the first systematic attempt to analyze this relationship in the case of Bangladesh.

The remainder of this paper is structured as follows: Sect. 2 reviews extant literature; Sect. 3 briefly describes the data used in the current study; Sect. 4 describes the econometric analyses performed using diverse econometric tools; and Sect. 5 concludes the paper. Other relevant information is listed in the "Appendix".

2 Literature review

Theoretical and empirical literature has investigated the link between RPV and inflation and it is "established wisdom" that RPV is influenced by inflation. An influential contribution is that of Parks (1978), a frequently cited study. Other relevant studies are those of Reinsdorf (1994), Debelle and Lamont (1996), Jaramillo (1999), Chang and Cheng (2000), Miszler and Nautz (2004), Nath (2004), and Choi (2010). These studies concentrate on the positive link between RPV and inflation and their functional relationship. For instance, Nautz and Scharff (2005) emphasize that the relationship is present not only in high inflation countries but also in low inflation ones. Various studies posit that the relationship is linear (Vining and Elwertowski 1976; Blejer and Leiderman 1980; Hercowitz 1981; Domberger 1987; Van Hoomissen 1988), while others posit that it is nonlinear (Tommasi 1993; Fielding and Mizen 2008; Choi 2010). Recent explanations of the functional relationship are becoming popular, since the "so-called linear relationship" cannot explain real data. When the operative relationship is non-monotonic, there can be a positive value of inflation for which RPV is minimized, and decreasing inflation is sensible only to a certain level. As such, there may be a strictly positive level of inflation around the general inflation rate that is optimal in terms of reducing price variations (Fielding and Mizen 2008).

Glezakos and Nugent (1986) criticize Parks' (1978) model for forecasting inflation. Parks (1978) estimated the relationship between RPV and inflation and found that inflation positively affects RPV. Additionally, he also examined the effect of unexpected inflation and found a significant relationship. Consequently, he used an AR(1) model to forecast inflation. Fischer et al. (1981) also use a fourth-order autoregression of inflation for forecasting. In both cases, the performance of the forecast is poor, as Rumler and Valderrama (2010) show that, for inflation forecasting, the structural model performs better than the time series model. Using a more valid model for forecasting, one may get appropriate results from examining the effect of various aspects of inflation on RPV. Lach and Tsiddon (1993, 1994) show that expected inflation is an important explanatory factor for intra-market price variability, while unexpected inflation is important for the inter-market variability. Moreover, Rather et al. (2014) show that inflation asymmetrically affects RPV when the relationship is U-shaped. Aarstol (1999), Nautz and Scharff (2005), and others examine the effects of the various aspects of inflation, namely expected inflation, unexpected inflation, and inflation uncertainty, on RPV. Grier and Perry (1996) estimate a bivariate GARCH-M model of inflation and RPV, and show that inflation uncertainty dominates trend inflation as a predictor of RPV. This review suggests that the model generating the various aspects of inflation is an important measure for examining the relationship. The theoretical literature contains two main approaches to the inflation-RPV relationship. The first is based on the signal extraction models of Lucas (1973, 1994) and Barro (1976), extended by Hercowitz (1981) and Cukierman (1983). According to the signal extraction model, RPV should be increasing in ex-ante inflation uncertainty. As such, the coefficient of positive and negative unexpected inflation will be different for the extension of the signal extraction model, and inflation uncertainty (variance of unexpected inflation) must be positive in the RPV-inflation function for the signal extraction model. Although this theory suggests a monotonic relationship between the absolute value of price shocks and RPV, there is no reason to assume an overly complex relationship.[1] The second approach is based on the menu-cost model, which assumes that RPV is increasing in the absolute value of expected inflation (Sheshinski and Weiss 1977; Rotemberg 1983; Benabou 1992). Rotemberg (1996)'s theoretical model implies that the optimal inflation rate is non-zero (grease effect), while according to Lucas (1973, 1994) and Barro (1976), the optimal inflation rate is zero (sand effect). This grease-sand effect of inflation is important for the monetary policy. Truman (2003) suggests that a country's optimal rate of inflation should be less that 5 % per annum. Moreover, Akerlof et al. (1996, 2000), and Akerlof and Kranton (2000) suggest that macroeconomic policy should aim for a rate of inflation in the range of 1.5–4 %. This implies that disinflationary policies may not improve welfare if the benefit from lower inflation is outweighed by the cost of increased price volatility (Vo et al. 2016). Therefore, the relationship between RPV and inflation is becoming complex in view of (1) the functional relationship and the associated model of expected inflation, (2) rejection or acceptance of the theories of that functional relationship, and (3) determining the threshold level of inflation.

[1] This does not hold true for the second theoretical strand, pioneered by Ball, Mankiw, and Rotemberg (see Rotemberg 1982, 1983), which suggests a potentially complex relationship between RPV and expected inflation.

3 RPV measurement

By defining relative price as a ratio of the individual price indices of n commodities to the overall price indices, the rate of change in the *i-th* commodity's relative price is $\pi_{it} - \pi_t$, with π_{it} defined as

$$\pi_{it} = 100 \times (\ln P_{i,t} - \ln P_{i,t-1}), \tag{1}$$

where $P_{i,t}$ is the CPI of consumption item i in period t. Similarly, π_t is defined as

$$\pi_t = 100 \times (\ln P_t - \ln P_{t-1}), \tag{2}$$

where P_t is the aggregate CPI in period t, and π_t is the aggregate inflation. Additionally, P_t can be defined as

$$P_t = \sum_{i=1}^{n} w_{it} P_{it}, \tag{3}$$

where w_{it} is the weight of individual commodities in the CPI basket, which sum to 1.

While the average change rates in relative prices are zero by definition, we take the variance of these changes as a measure of the degree of relative price variability, following Parks (1978), Blejer and Leiderman (1980) and Parsley (1996), etc. This is calculated as the weighted sum of the squared deviations of the individual rates of price change around the average, that is, the RPV, which can be defined as[2]

$$RPV_t = \sqrt{\sum_{i=1}^{n} w_i (\pi_{i,t} - \bar{\pi}_t)^2}, \tag{4}$$

where $\bar{\pi}_t = (1/n) \sum_{i=1}^{n} \pi_{i,t}$ is the mean price change in period t. Note that n is the number of consumption items.

RPV is measured using all eight subcomponents of the all-commodities CPI. One of the problems of measuring RPV is that the weights of the different categories of commodities change over time. Therefore, a large span of data may cause measurement problems. Another issue is the change in the base year. The above RPV equation contrasts with the measures computed by Fischer et al. (1981), and Grier and Perry (1996), which deliberately ignore the energy, and the food and energy subcomponents, respectively, in an effort to control for supply shocks. Given the importance of market-specific shocks to the prediction of the Lucas-Barro model (Lucas 1973; Barro 1976), the use of the most comprehensive measure of RPV is appropriate as such.

[2] Note that P_t is constructed as a weighted index of all underlying prices and, therefore, it is desirable that both RPV_t and $\bar{\pi}_t$ are calculated as weighted SD and mean, respectively. However, we find that the results do not change. Moreover, prominent studies on the United States (e.g., Vining and Elwertowski (1976)) use unweighted measures. Aarstol (1999) also uses the unweighted measure of RPV.

4 Econometric analysis

Although a visual inspection of the scatter and quadratic plots in Fig. 2 could be a guideline for the underlying relationship between RPV and inflation, this may reduce the evaluation of the actual relationship for structural changes. In this section, more formal econometric techniques are utilized to provide a better understanding of the underlying relationship. We also reexamine Parks (1978) and Fischer et al. (1981) in the case of Bangladesh. Additionally, we analyze the menu-cost, Lucas-Barro signal-extraction (Lucas 1973; Barro 1976), and Hercowitz-Cukierman signal-extraction (Hercowitz 1981; Cukierman 1983) models of RPV and inflation uncertainty. Therefore, this section is comprised of: (a) a reexamination of Parks (1978) and Fischer et al. (1981); (b) a semiparametric regression analysis; (c) the relationship between RPV and inflation uncertainty; (d) rolling regression; and (e) the multivariate multiple structural test developed by Bai and Perron (1998, 2003).

4.1 Reexamination of Parks (1978) and Fischer et al. (1981)

4.1.1 Parks' (1978) preliminary estimation

Parks (1978) developed two variables used in his preliminary equation. The first is DP, a measure of aggregate inflation. It is the growth rate weighted average of the implicit price deflator for each commodity (Dp_i), where weights (w_i) are relative expenditures ($DP = \sum w_i Dp_i$). The second variable is the weighted variance (VP) of inflation in individual categories of the aggregate inflation rate ($VP = \sum w_i(Dp_i - DP)^2$ or $VP_t = \sum w_{it}(Dp_{it} - DP_t)^2$ if we consider the time dimension). Parks (1978) uses the US CPI from 1929 to 1975, as does this study. We use π_t as the aggregate inflation rate and RPV_t as RPV.[3] To establish the importance of these two measures, we regress the inflation rate on RPV. Parks (1978) assumed that the relationship must consider both inflation and deflation and used the following equations: $RPV_t = \alpha_0 + \alpha_1|\pi_t| + \varepsilon_t$ or $VP_t = \alpha_0 + \alpha_1\pi_t^2 + \varepsilon_t$. Alternatively, he examined the positive and negative price changes on RPV with the following equation: $RPV_t = \beta_0 + \beta_1|\pi_t^+|^2 + \beta_2|\pi_t^-|^2 + \varepsilon_t$, where $|\pi_t^+|^2$ (or $|\pi_t^-|^2$) represents the product of π_t and a dummy variable that takes the value 1 when π_t is positive (or negative) and 0 otherwise.

Table 1 presents the regression of inflation on RPV following Parks (1978). We use a sample of monthly observations from July 2002 to June 2013. The first row indicates that the relationship between inflation and RPV is positive and highly significant. Recent work on RPV-inflation also supports this result (Reinsdorf 1994; Miszler and Nautz 2004; Nath 2004).

If both variables are jointly determined by the same individual price shocks, there is a correlation between inflation and the error term of the equation. Therefore, the

[3] We use RPV as the square root of $VP_t = \sum w_{it}(\pi_{it} - \bar{\pi}_t)^2$ (i.e., $RPV_t = \sqrt{VP_t}$), the standard deviation and $DP_t = \pi_t$ being lagged differences of the natural logarithm of CPI multiplied by 100. Blejer and Leiderman (1980) and Parsley (1996), among others, use the same method.

Table 1 The inflation-RPV link for Bangladesh

$RPV_t = \alpha_0 + \alpha_1|\pi_t| + \varepsilon_t$

	α_0	α_1	R^2	ARCH (4)	Endogeneity test [differences in J-stats]
OLS	0.172*** [5.296]	0.460*** [14.738]	0.63	F (4,122) = 1.37 (0.24)	
TSLS	0.198*** [2.845]	0.430*** [5.157]	0.62	F (4,120) = 1.41 (0.23)	0.144 (0.70)
GMM	0.239*** [3.702]	0.371*** [4.964]	0.60		0.802 (0.37)

t-statistics are given in square brackets and p values are given in 1st brackets. ARCH test is employed to examine the ARCH effects where the null hypothesis is there is no ARCH effect. The over-identifying restriction test is the differences in J-stats where the null hypothesis is the over-identifying restrictions are valid. For TSLS and GMM, the structural determinants of inflation from Moshiri and Cameron (2000) are used as instruments. The instruments are: $\hat{y}$, $\hat{m}_t$, π_t and *oil*. The definition of these determinants are given in "Appendix"

*, ** and *** represent 10, 5 and 1 % level of significance respectively

parameter α_1 will be biased. To rectify this, we use the two-stage leasts squares (TSLS) and generalized method of moments (GMM estimation methods, with the structural determinants of inflation suggested by Moshiri and Cameron (2000). Re-estimating the equation with GMM and TSLS yields the results in rows 2 and 3 of Table 1. The GMM and TSLS estimators confirm that the RPV-inflation relationship is strong and significant. The test of endogeneity confirms that the instruments used are valid, and the variable π_t is exogenous. It is important to note that the ordinary least squares (OLS) estimator is upward biased.

Considering both the inflation and deflation rate, we estimate the equation following Parks (1978). Using the methodology previously described, Table 2 shows the estimation results.

In column 2, Model 1 exhibits only one regressor, the squared-value of π_t, whose coefficient is positive and highly significant. The model diagnostic tests also support the model specified. Model 1 is free from second-order residual autocorrelation and from heteroskedasticity, but fails to remove the ARCH effects. However, the Chow test shows that the model is not free from structural breaks (Chow 1960), as discussed in Sect. 4.5, Testing for multiple structural breaks. Model 2 also shows that the coefficients of π_+^2 and π_-^2 are positive and statistically significant, but the magnitude of deflation is more than that of inflation. As we use monthly data, the total observations with negative values are less than the ones with positive values. Consequently, if the inflation rate remains constant, the variance will decline. This variation disregards the notion of expected and unexpected inflation, which is discussed in Sect. 4.3.3.

4.1.2 Determinants of RPV movement

In order to identify the determinants of movement in RPV, Parks (1978) uses a more refined specification by constructing two additional variables. The first is a measure of real spending growth ($m_t - \pi_t$), and the second is the unexpected rate of inflation,

Table 2 The RPV-inflation link for Bangladesh (consider inflation and deflation)

Variable	Model 1	Model 2
Intercept	0.361*** [13.337]	0.336*** [12.856]
π^2	0.161*** [9.185]	
π_+^2		0.161*** [9.487]
π_-^2		0.485*** [7.704]
R^2	0.63	0.68
Adj. R^2	0.63	0.67
Standard error	0.241	0.224
F-statistics	223.709***	138.439***
AR(2) test	F(2,127) = 1.578 (0.21)	F(2,126) = 1.427 (0.24)
ARCH(1) test	F(1,128) = 3.761 (0.05)	F(1,128) = 5.476 (0.02)
Heteroskedasticity	F(1,129) = 1.482 (0.22)	F(2,128) = 0.916 (0.40)
Chow test	F(61,68) = 1.985 (0.00)	F(61,67) = 3.305 (0.00)

This table reports the parametric OLS estimation. The dependent variable is the relative price variability (RPV). The following test statistics are reported: (a) AR(2) = LM test for 2nd order residual autocorrelation when the null hypothesis is there is no serial autocorrelation; (b) ARCH(1) = LM test for 1st order ARCH test when the null hypothesis is there is no ARCH effects; (c) Heteroskedasticity test = -Breusch-Pegan Heteroskedasticity test when the null hypothesis is there is no heteroskedasticity; (4) Chow test = Chow forecast test for stability when the breakpoint is selected on the basis of residual plot. The figures in the square brackets are t values and in the first brackets are p values

*, ** and *** represents 10, 5 and 1 % level of significance respectively

($[\pi_t - E(\pi)]$), where m_t is the growth rate of the money supply. We used the growth rate of the broad money supply in our estimation; moreover, $[\pi_t - E(\pi)]$ is the unexpected rate of inflation, where $E(\pi)$ is the expected rate.[4] The way basic determinants of supply and demand combine with expectations on the rate of inflation affect the RPV resulting from an income change. The estimated equation is

$$\begin{aligned} RPV_t = {} & \alpha_0 + \alpha_1(m_t - \pi_t)^2 + \alpha_2[\pi_t - E(\pi)]^2 \\ & + \alpha_3(m_t - \pi_t)[\pi_t - E(\pi)] + \alpha_4(m_t - \pi_t) + \alpha_5[\pi_t - E(\pi)] + u_t \end{aligned}. \tag{5}$$

Table 3 reports the estimation of Eq. (5). Parks (1978) assumes that the coefficient α_1 and α_2 should be positive, where α_1 measures the real income effect, and α_2 measures the effect of the unexpected rate of inflation. The coefficient of supply shifts is α_3, and its sign should be negative. The coefficient of the supply trend term is α_4, and its predicted sign may be positive or negative. If both α_0 and α_5 are positive, then α_4 will be negative and vice versa.

The estimated results clearly resemble Parks' (1978) in that real income and unexpected inflation affect RPV positively, and their effects are significant both in linear and squared form. This allows the supply shifts to influence price level

[4] In this paper, we use both structural and statistical models to forecast the inflation rate. The deviation of the actual and the forecasted values is expressed as the unexpected rate of inflation. This is a widely accepted method to calculate the rate of unexpected inflation. For details, see footnote 7.

Table 3 Determinants of movement in RPV

Coefficient	Value	t ratio
α_0	0.191***	4.178
α_1	0.002	0.202
α_2	0.309***	6.120
α_3	−0.120**	−2.061
α_4	−0.013	−0.228
α_5	0.249***	2.669
R^2	0.78	
Adj. R^2	0.77	
Standard error	0.44	
F-statistics	88.025***	
AR(2) test	F(2,121) = 0.842 (0.43)	
ARCH(1) test	F(1,126) = 0.152 (0.69)	
Heteroskedasticity	F(5,123) = 0.989 (0.42)	
Chow test	F(61,62) = 2.305 (0.00)	

This table reports the parametric OLS estimation. The dependent variable is the square of relative price variability (VP). The following test statistics are reported: (a) AR(2) = LM test for 2nd order residual autocorrelation when the null hypothesis is there is no serial autocorrelation; (b) ARCH(1) = LM test for 1st order ARCH test when the null hypothesis is there is no ARCH effects; (c) Heteroskedasticity test = Breusch-Pegan Heteroskedasticity test when the null hypothesis is there is no heteroskedasticity; (4) Chow test = Chow forecast test for stability when the breakpoint is selected on the basis of residual plot. The figures in the first brackets are p values

*, ** and *** represents 10, 5 and 1 % level of significance respectively

variation. One of the objectives of this estimation is to check whether unanticipated inflation influences RPV, and the results show that it has a distinct effect on RPV.

4.1.3 Fischer et al.'s (1981) reexamination

Fischer et al. (1981) also examine the relationship between RPV and the inflation rate using the consumption price deflator for USA from 1930 to 1980. They compute RPV using different categories of components and regress inflation on those variables separately. Before estimation, they discuss six approaches to explain the relationship between inflation and RPV. Table 4 summarizes these approaches.

Table 4 suggests that, depending on the source of disturbance, RPV might be associated with the inflation rate itself, the absolute value of the inflation rate, or changes in the inflation rate (in absolute value). Therefore, the regression equation can be written as

$$RPV_t = \beta_0 + \beta_1\pi_t + \beta_2|\pi_t| + \beta_3 d\pi_t + \beta_4|d\pi_t| + \vartheta_t, \qquad (6)$$

where $d\pi_t$ is the inflation rate change and $|d\pi_t|$ is the absolute value of this change.

Table 5 shows the regression results of Eq. (6). We split the samples in two, based on Fig. 2 and the global financial and food crises of 2008. We also estimate the equation using alternative definitions of RPV: RPV* is calculated by subtracting

Table 4 Summary of approaches from Fischer et al. 1981. *Source*: Fischer et al. 1981

Serial no.	Approach	Exogenous factors	Function of inflation associated with RPV	Welfare implications
1	Market clearing with imperfect information	Policy disturbances	Unanticipated inflation or deflation	Misperceived aggregate disturbances produce resource misallocations
2	Menu costs	Inflation rate	Inflation or deflation	Inflation or deflation creates resource misallocations and generates unnecessary transaction costs
3	Asymmetric price response	Relative disturbances	Either inflation rate or inflation in excess of base rate	Price inflexibility leads to resource misallocations: there is too little relative price variability
4	Relative shocks same as aggregate shocks	Changes in policy	Changes in inflation rate	Given the changes in policy, relative prices should vary from efficient allocation
5	Allocative effects of macro policy	Changes in policy	Changes in inflation rate	Given the changes in policy, relative prices should vary for efficient allocation
6	Endogenous policy	Real disturbances	Same as 3	Policy may offset welfare loss associated with relative shocks by making appropriate price adjustments possible

the food and energy component from the RPV calculation.[5] Columns 2–4 show the results of linking RPV with the inflation rate and the changes in the inflation rate series in absolute values. All columns show that RPV is associated with both inflation and deflation, and changes in inflation in either direction (Models 1–3). This resembles Fischer et al.'s (1981) study and is consistent with each of the six approaches in Table 4. Although the coefficient of the inflation rate change in absolute value remains constant, other coefficients change across the models. This suggests that the relationship is not stable. Models 1 and 2 have structural breaks (in 2008 and 2004, respectively), while Model 3 does not exhibit any breaks. The first break can be attributed to the global financial and food crises and the second break to the supply shock caused by the floods in 2004. The Chow-Forecast test of Model 3 suggests that it has no structural break or parameter inconstancy between 2008M07 and 2013M06. Model 4 shows the same regression where the dependent variable is RPV*, delivering similar results to Models 1–3. This suggests that the effect of the supply shock is stronger in the RPV-inflation relationship, as the coefficients of Model 4 are larger than those of Model 1.

Fischer et al. (1981) also estimate the relationship between RPV, and actual and unexpected (actual and absolute value) rates of inflation using the following equation:

[5] This calculation is performed for two reasons: to check the stability of the RPV-inflation relationship and the effect of supply shocks on RPV. Numerous researchers have done this, such as Fischer et al. (1981), Aarstol (1999), etc.

Table 5 Fischer regression of Eq. (6)

Variable	Dependent Variable and period			
	Model 1: RPV [full sample]	Model 2: RPV [2002M07–2008M06]	Model 3: RPV [2008M07–2013M06]	Model 4: RPV* [full sample]
π_t	−0.127*** [−3.408]	−0.208*** [−6.691]	−0.055 [−1.003]	−0.497*** [−3.539]
$\lvert\pi_t\rvert$	0.630*** [16.153]	0.699*** [18.613]	0.561*** [11.108]	2.070*** [12.839]
$d\pi_t$	−0.025 [−0.945]	0.035 [1.199]	−0.078** [−2.081]	−0.026 [0.296]
$\lvert d\pi_t\rvert$	0.111*** [3.574]	0.070* [1.921]	0.145*** [3.960]	0.283** [2.506]
R^2	0.65	0.86	0.35	0.62
Adj. R^2	0.64	0.85	0.32	0.60
S.E. of regression	0.23	0.16	0.29	0.77
AR(2)	F(2,124) = 3.350 (0.04)	F(2,64) = 0.554 (0.58)	F(2,54) = 3.364 (0.04)	F(2,124) = 1.762 (0.17)
ARCH(1)	F(1,127) = 5.371 (0.02)	F(1,67) = 0.095 (0.76)	F(1,57) = 1.801 (0.18)	F(1,127) = 0.439 (0.50)
Heteroskedasticity	F(4,125) = 0.971 (0.43)	F(4,65) = 1.666 (0.17)	F(4,55) = 1.719 (0.16)	F(4,125) = 0.905 (0.46)
Chow-forecast test	F(61,65) = 3.516 (0.00)	F(31,34) = 1.620 (0.08)	F(34,22) = 0.986 (0.52)	F(61,64) = 3.264 (0.00)

This table reports the parametric OLS estimation of Eq. (6) in the main text. The dependent variable is the relative price variability (RPV) and alternative relative price variability (RPV*). Explanatory variables are: inflation rate, absolute value of inflation rate, change in inflation rate and absolute value of change in inflation rate. The following test statistics are reported: (a) AR(2) = LM test for 2nd order residual autocorrelation when the null hypothesis is there is no serial autocorrelation; (b) ARCH(1) = LM test for 1st order ARCH test when the null hypothesis is there is no ARCH effects; (c) Heteroskedasticity test = Breusch-Pegan Heteroskedasticity test when the null hypothesis is there is no heteroskedasticity; (4) Chow test = Chow forecast test for stability when the breakpoint is selected on the basis of residual plot. The figures in the square brackets are t values and in the first brackets are p values

*, ** and *** represents 10, 5 and 1 % level of significance respectively

$$RPV_t = \delta_0 + \delta_1 E(\pi_t) + \delta_2 |E(\pi_t)| + \delta_3 [\pi_t - E(\pi_t)] + \delta_4 |\pi_t - E(\pi_t)| + e_t. \quad (7)$$

Table 6 shows the estimation results of Eq. (7). The lack of significance of the expected rate of inflation in Models 1–4 suggests that anticipated inflation is neutral. The significance of the unexpected inflation (actual and absolute) confirms the view associated with market clearing, that is, the rational expectation approach. However, anticipated inflation could be associated with RPV if the inflationary shock has a delayed effect on the economy.

Both Tables 5 and 6 express that the change in inflation rate or the unexpected inflation (actual and absolute) affect RPV significantly, which confirms that RPV responds to these variables asymmetrically. Additionally, variability increases more when unexpected inflation (or the change in the inflation rate) is positive than when it is negative (i.e., the value of the coefficients is larger in each case). This view is consistent with approach 3 in Table 4.

4.2 Semiparametric regression analysis

Although the scatter plot in Fig. 2 indicates a quadratic form of the RPV-inflation relationship, it is favorable to place as few restrictions as possible on the functional form, especially when economic theory does not provide any clear guidance on this functional form. Parametric models are required because they can be precisely estimated, and the fitted parametric model easily interpreted. However, if underlying assumptions are violated, the parametric models may present a misleading picture of the relationship. In contrast, a nonparametric approach provides clear advantages, as it avoids restrictive assumptions on the functional relationship. Despite these advantages, a nonparametric approach may not be easy to interpret, and may yield inconclusive results if the number of regressors is large.

Considering these pros and cons, a semiparametric approach combines features of both parametric and nonparametric models, which are useful in interpreting the model and provide flexibility at the same time.

Following Fielding and Mizen (2008) and Choi (2010), we employ the following partially linear regression model:

$$RPV_t = X'\beta + g(\pi_t) + \varepsilon_t, \quad (8)$$

where X_t is a $(p + q) \times 1$ vector of the regressors that includes lagged terms of RPV and inflation, $X' = \{RPV_{t-1}, \ldots, RPV_{t-p}, \pi_{t-1}, \ldots, \pi_{t-q}\}$. $g(.)$ is an unknown smooth differential function that captures the contemporaneous effect of inflation on RPV and determines the underlying functional form of the relationship between inflation and RPV (see Choi 2010).

Technically, we estimate Eq. (8) using two estimators: Yatchew's (1998) and Robinson's (1988) partially linear estimators. We compare the two estimators to check the g(.) function.[6]

[6] If the true functional form is quadratic, for instance, $g(.)$ will take the form of $g(\pi_t) = \delta_1 + \delta_2\pi_t + \delta_3\pi_t^2$. See Choi (2010) for further details.

Table 6 Fischer regression of Eq. (7)

Variable	Dependent variable and period			
	Model 1: RPV [full sample]	Model 2: RPV [2002M07–2008M06]	Model 3: RPV [2008M07–2013M06]	Model 4: RPV* [full sample]
$E(\pi_t)$	−0.144	−0.299	−0.049	−0.729
	[−1.119]	[−1.202]	[−0.285]	[−1.435]
$\lvert E(\pi_t)\rvert$	0.483***	0.521**	0.546***	1.698***
	[3.603]	[2.138]	[2.850]	[3.235]
$\pi_t - E(\pi_t)$	0.168***	0.181***	0.140***	0.540***
	[4.289]	[3.309]	[2.697]	[4.292]
$\lvert\pi_t - E(\pi_t)\rvert$	0.509***	0.562***	0.413***	1.644***
	[9.133]	[7.727]	[5.510]	[9.947]
R^2	0.50	0.64	0.36	0.48
Adj. R^2	0.49	0.62	0.33	0.47
S.E. of regression	0.28	0.26	0.29	0.90
AR(2) test	F(2,123) = 0.549 (0.58)	F(2,63) = 0.710 (0.50)	F(2,54) = 0.707 (0.50)	F(2,123) = 0.266 (0.77)
ARCH(1) test	F(1,126) = 4.319 (0.04)	F(1,66) = 2.255 (0.13)	F(1,57) = 2.019 (0.16)	F((1,126) = 0.760 (0.38)
Heteroskedasticity test	F(4,124) = 4.809 (0.00)	F(4,64) = 5.267 (0.00)	F(4,55) = 1.707 (0.16)	F(4,124) = 3.930 (0.00)
Chow forecast test	F*61,64) = 1.340 (0.12)	F(31,34) = 2.996 (0.00)	F(34,22) = 1.629 (0.11)	F(61,64)) = 1.269 (0.17)

This table reports the parametric OLS estimation of Eq. (7) in the main text. The dependent variable is the relative price variability (RPV) and alternative relative price variability (RPV*). Explanatory variables are: expected inflation rate, absolute value of expected inflation rate, unexpected inflation, Absolute value of unexpected inflation. All of the expected and unexpected value of inflation is derived from Moshiri and Cameron (2000). The following test statistics are reported: (a) AR(2) = LM test for 2nd order residual autocorrelation when the null hypothesis is there is no serial autocorrelation; (b) ARCH(1) = LM test for 1st order ARCH test when the null hypothesis is there is no ARCH effects; (c) Heteroskedasticity test = Breusch-Pegan Heteroskedasticity test when the null hypothesis is there is no heteroskedasticity; (4) Chow test = Chow forecast test for stability when the breakpoint is selected on the basis of residual plot. The figures in the square brackets are t values and in the first brackets are p values

*, ** and *** represents 10, 5 and 1 % level of significance respectively

Table 7 Semiparametric regression

Sample	Robinson's estimator		Yatchew's estimator	
	Variable	Coefficient	Variable	Coefficient
2002M07–2008M06	RPV (−1)	−0.032	RPV (−1)	−0.087
		[−0.58]		[−1.05]
	π (−1)	−0.004	π (−1)	0.0008
		[−0.19]		[0.02]
2008M07–2013M06	RPV (−1)	0.173**	RPV (−1)	0.198*
		[2.01]		[1.92]
	π (−1)	0.024	π (−1)	0.005
		[0.55]		[0.11]
Full sample	RPV (−1)	0.118**	RPV (−1)	0.068
		[2.22]		[1.02]
	π (−1)	0.018	π (−1)	0.028
		[0.72]		[0.87]

Figures in square bracket are t ratios. (−1) indicates the variables in lagged in one period

*, ** and *** represent 10, 5 and 1 % level of significance respectively

Table 7 shows the semiparametric estimation of the $g(.)$ function as per Robinson (1988) and Yatchew (1998). The optimum number of lag lengths for X vector is 1, which is determined by AIC. The coefficients are in their derivative form. The results of second and third sample (2008M07–2013M06 and full sample) resemble one other, while the first sample produces a different result. As our main interest is to estimate the $g(.)$ function, Fig. 3 shows the $g(.)$ function for various values of π_t. Panel (A) shows the estimation of the $g(.)$ function using Robinson's (1988) estimator for different samples. All figures show that the function is nearly quadratic with few exceptions. The exception is the smoothing pattern of the function. The last column shows a different pattern, although we can conclude that the function is nonlinear and U-shaped. One of the tests of nonlinearity tests whether $g'(.)$ is linear.

If the function is U-shaped, there must be a point where $g'(.) = 0$ and RPV is minimized at the corresponding inflation rate. In this paper, we denote it as π^*. If the inflation rate is below π^*, then $g'(.) < 0$ and $g(.)$ is downward sloping, while when $g'(.) > 0$, $g(.)$ is upward sloping if the inflation rate is above π^*. This suggests that the g(.) function is closer to being U-shaped than V-shaped (Parks 1978; Tommasi 1993; Bomberger and Makinen 1993). Additionally, Fielding and Mizen (2008) point out that $g(.)$ is approximately U-shaped.

Panel (B) shows the estimation of the $g(.)$ function using Yatchew's (1998) estimator. Using this partial linear methodology, we estimate Eq. (8). Panel (B) shows the locally weighted scatterplot smoothing (lowess) figure using Yatchew's (1998) difference estimator. All of the figures show that the relationship between RPV and inflation is nonlinear and approximately U-shaped, which resembles Panel (A), confirming that the relationship is nonlinear and U-shaped. One of the advantages of the semiparametric regression method is that it enables us

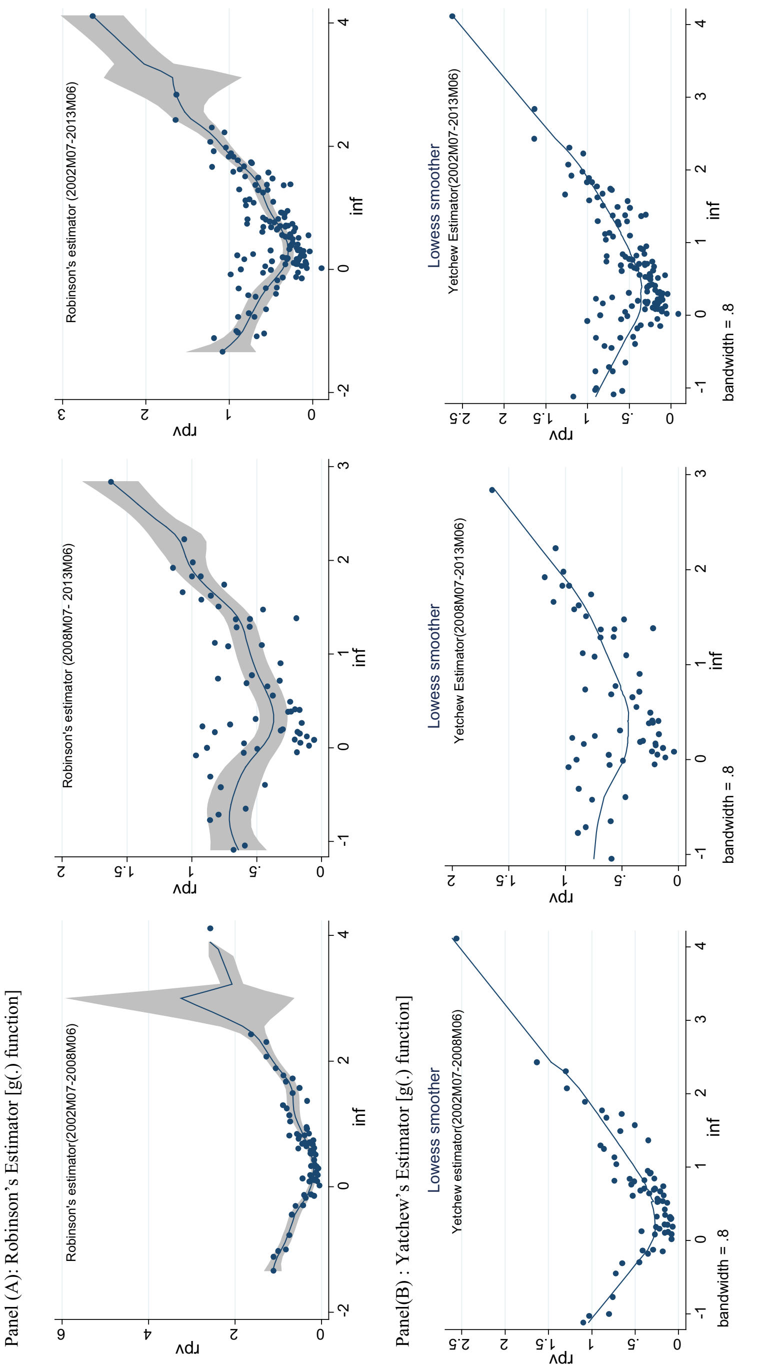


Fig. 3 The g(.) function from semiparametric regression of Eq. (8). *Panel* **a** shows the Robinson's estimator of Robinson (1988) and *panel* **b** shows the Yatchew's estimator of Yatchew (1998). *Panel* **a** shows the results with confidence intervals. Each column of two panels shows the sample: 2002M07–2008M06; 2008M07–2013M06 and 2002M07–2013M07

to track the stability of $g(.)$ by examining whether this function shifts across samples. Panel (A) of Fig. 2 shows that the second column is different from the others. Moreover, in Panel (B), the second column is also different. Visual inspection shows that the threshold level of the inflation rate, π^*, is different in each sample for both estimators, ranging from 1.5 to 3 % (yearly).

To summarize, the results of the semiparametric analysis support the visual and parametric evidence in Fig. 1. All of the estimators confirm that the relationship between RPV and inflation is U-shaped and shifts across sub-sample periods, indicating a time-varying pattern.

4.3 Relationship between RPV and inflation uncertainty

4.3.1 Model specification

Following Aarstol (1999) and Nautz and Scharff (2005), we examine the relationship between RPV and inflation uncertainty, considering both the signal extraction and menu cost features. For robustness, we employ a number of models for forecasting inflation. While in the previous section we established the nonlinear relationship between RPV and inflation through a semiparametric regression, in this section we estimate a parametric regression, following Fielding and Mizen (2008) and Choi (2010), regarding different models of inflation forecasting.[7]

Following Fielding and Mizen (2008) the basic regression equation becomes

$$RPV_t = g(E(\pi_t)) + \alpha_1 RPV_{t-1} + \alpha_2 UIN + \alpha_3 UIP + \alpha_4 \sigma_t + u_t, \tag{9}$$

where π_t is the inflation rate. The function $g(.)$ captures the menu-cost effects and can take different forms.

We briefly describe the models of generating inflation expectations. After estimation of a model, the one-period-ahead (static) forecast is denoted as $E[\pi_t]$. The same model generates the unexpected component of π_t, that is, $[\pi_t - E(\pi_t)]$. Moreover, the regression equation includes two auxiliary series, derived from $[\pi_t - E(\pi_t)]$. $[\pi_t - E(\pi_t)]^+$ equals the absolute value of unexpected inflation when it takes positive values (and 0 otherwise), denoted as unexpected inflation positive (UIP). $[\pi_t - E(\pi_t)]^-$ equals the absolute value of unexpected inflation when it takes negative values (and 0 otherwise), denoted as unexpected inflation negative (UIN). We include these series to test the hypothesis of the Hercowitz-Cukierman model (see Aarstol 1999). We also include σ_t, the conditional variance of inflation generated from a GARCH model of inflation, and a six-month variance of the unexpected inflation rate (σ_t^{six}) as a regressor, if the models are other than the

[7] As we are examining the effect of expected inflation on RPV, there are several methods to estimate expected inflation. In this study, we estimate several structural and time-series models of inflation, namely three structural models and two time-series models. These include the estimation of a preferred model of inflation, which is used to explain the inflationary situation in Bangladesh by IMF. Additionally, an open-economy model of inflation for a developing country like Bangladesh is also estimated. We decompose the forecasted inflation into expected and unexpected inflation. The used models are from Moshiri and Cameron (2000), Hossain (2002), IMF (2007), Fielding and Mizen (2008), and the authors' estimated times series model. A brief description of these is presented in the "Appendix".

univariate time series models. Finally, the regression model includes the lagged dependent variable, RPV_{t-1}. One of the tests for the function to be U-shaped is the sign of the coefficients β_2 and β_3. If $\beta_2 < 0$ and $\beta_3 > 0$, the function is U-shaped and the threshold level of inflation is positive. We include this variable to test that the variation in RPV is not due to the deterministic process of $g(.)$, but to measurement errors or changes in firm composition.

As we established that $g(.)$ is U-shaped, we follow the following structure of this function:

$$g(E(\pi_t)) = \beta_1 + \beta_2 E(\pi_t) + \beta_3 E(\pi_t)^2. \tag{10}$$

Therefore, the estimated form of the regression equation is[8]:

$$RPV_t = \beta_1 + \alpha_1 RPV_{t-1} + \beta_2 E(\pi_t) + \beta_3 E(\pi_t)^2 + +\alpha_2(UIP) + \alpha_3(UIN) + \alpha_4 \sigma_t^{six} + e_t \tag{11}$$

or

$$RPV_t = \beta_1 + \alpha_1 RPV_{t-1} + \beta_2 E(\pi_t) + \beta_3 E(\pi_t)^2 + +\alpha_2(UIP) + \alpha_3(UIN) + \alpha_4 \sigma_t + e_t. \tag{12}$$

4.3.2 *Inflation uncertainty and RPV*

Although the relationship between RPV and inflation in the previous section was estimated as per Parks (1978) and Fischer et al. (1981), the estimated equations are misspecified, as they do not consider both the quadratic specification of the regression and the unexpected inflation simultaneously. Aarstol (1999) examines the uncertainty effect of inflation on RPV. As such, we estimate the effect of inflation uncertainty by reexamining Nautz and Scharff (2005) and Aarstol (1999), checking whether inflation uncertainty affects RPV.

After the estimation of expected and unexpected inflation using the models described in the Appendix, we estimate the quadratic Eq. (11) and (12).

Nautz and Scharff (2005) estimated the following equation for Germany:

$$RPV_t = \beta_1 + \beta_2 |E(\pi_t)| + \beta_3 |\pi_t - E(\pi_t)| + u_t, \tag{13}$$

where the absolute values of $E(\pi_t)$ and $[\pi_t - E(\pi_t)]$ are used to examine the effect of expected and unexpected inflation on RPV. Table 8 shows the estimation of Eq. (13) in the case of Bangladesh, using the expected inflation from Moshiri and Cameron (2000).[9]

Furthermore, Table 8 shows that, irrespective of the estimation strategies, the unexpected inflation, rather than the expected one, is more important for RPV in

[8] We use the expected inflation rather than the actual inflation to capture the effect of inflationary expectations.

[9] We use the expected inflation from the other models of inflation, obtaining similar results.

Table 8 The impact of expected and unexpected inflation on RPV

	$RPV_t = \beta_1 + \beta_2\lvert E(\pi_t)\rvert + \beta_3\lvert \pi_t - E(\pi_t)\rvert + u_t$						
	β_1	β_2	β_3	H_0: $\beta_2 = \beta_3$	R^2	ARCH(4)	Endogeneity test [differences in J-stats]
OLS	0.166** [2.592]	0.155 [1.553]	0.497*** [6.317]	F(1,126) = 6.592 (0.01)	0.44	F(4,120) = 0.208 (0.93)	
TSLS	0.188* [1.682]	0.169 [1.483]	0.441* [1.873]	F(1,126) = 0.780 (0.37)	0.44	F(4,120) = 0.526 (0.71)	0.044 (0.83)
GMM	0.204* [1.780]	0.122 [1.055]	0.459* [1.822]	F(1,126) = 1.052 (0.30)	0.44		0.036 (0.84)

t Statistics are given in square brackets and p values are given in 1st brackets. ARCH test is employed to examine the ARCH effects where the null hypothesis is there is no ARCH effect. The over-identifying restriction test is the differences in J-stats where the null hypothesis is the over-identifying restrictions are valid. For TSLS and GMM, the structural determinants of inflation from Moshiri and Cameron (2000) are used as instrument. The instruments are: $\hat{y}$, $\hat{m}_t$, π_t and *oil*. The definition of these determinants are given in "Appendix"

*, ** and *** represent 10, 5 and 1 % level of significance respectively

Bangladesh. None of the estimation strategies show a significant impact of the expected inflation. Although our OLS results show that the equality of coefficients via a standard F-test yields a rejection at the 5 % significance level, the rest of the results lead to different conclusions.

If unexpected inflation is unexpectedly high, that is, when $[\pi_t - E(\pi_t)] > 0$, RPV is affected more (Aarstol 1999; Choi 2010). In order to examine this effect in Bangladesh, we run the following regression:

$$RPV_t = \delta_1 + \delta_2 |E(\pi_t)| + \delta_3 UIP_t + \delta_4 UIN_t + \varphi_t. \tag{14}$$

Table 9 shows the estimation results of Eq. (14), which confirm that RPV is particularly strong in Bangladesh when unexpected inflation is positive. In all estimation strategies, the hypothesis that the coefficients of positive and negative unexpected inflation are equal is rejected at the 1 % significance level.

4.3.3 Expected inflation, unexpected inflation, and RPV: the quadratic function

Hitherto, we analyzed the relationship between RPV and inflation within a simple linear framework. However, as we established earlier that the relationship between RPV and inflation is nonlinear and U-shaped, the analysis contains misspecification errors. After semiparametric estimation, we specify the relationship through Eqs. (11) and (12). Van Hoomissen (1988), Fielding and Mizen (2008) and Choi (2010), etc. examine this relationship using both parametric and nonparametric regressions. Following these studies, we also examine the relationships given in Table 10.

Table 10 shows the estimation of Eqs. (11) and (12) using the OLS method, where we have used the expected and unexpected inflations from four different types of models. The equations are similar to Aarstol's (1999), except that they are nonlinear. As previously stated, since the decomposition of inflation is derived from different models, we check the robustness of our specifications and the decomposition.

OLS 1 shows the estimation of Eq. (11) using the decomposition of inflation from Moshiri and Cameron (2000). Our preliminary check on the coefficients β_2 and β_3 ensures the positive quadratic relationship between RPV and expected inflation, where β_3 is positive and statistically significant. The coefficients of UIP and UIN are also positive and highly significant, indicating that both positive and negative unexpected inflation have a significant effect on RPV. This affects inter-market price variability, in that unexpected inflation produces an upward pressure on this variability. Additionally, the UIP coefficient being greater than that of UIN (across all models), indicates the rejection of the Hercowitz-Cukierman version of the extraction model, which predicts equal coefficients on positive and negative unexpected inflation. Therefore, the test of UIP = UIN is rejected at the 1 % significance level. This asymmetry resembles Fischer et al.'s (1981). The discrepancy is best analyzed by Hajzler and Fielding (2014), who suggest that consumer search theory can predict the positive effect of both positive and negative unanticipated inflation on RPV. The coefficient of σ_t^{six} is insignificant, which rejects

Table 9 The inflation RPV link allowing for asymmetry

$RPV_t = \delta_1 + \delta_2|E(\pi_t)| + \delta_3 UIP_t + \delta_4 UIN_t + \varphi_t$

	δ_1	δ_2	δ_3	δ_4	R^2	H_0: $\delta_3 = \delta_4$	ARCH(4)	Endogeneity test [differences in J-stats]
OLS	0.181*** [3.044]	0.174* [1.729]	0.618*** [13.352]	0.274*** [3.726]	0.55	F(1,125) = 23.303 (0.00)	F(4,120) = 0.263 (0.90)	
TSLS	0.276*** [3.008]	0.154 [1.454]	0.516*** [3.826]	0.196 [1.326]	0.53	F(1,123) = 14.464 (0.00)	F(4,118) = 0.471 (0.75)	3.909 (0.41)
GMM	0.214* [1.812]	0.154 [1.531]	0.593*** [2.863]	0.235 [0.923]	0.55	F(1,125) = 13.646 (0.00)		0.233 (0.97)

t Statistics are given in square brackets and p values are given in 1st brackets. ARCH test is employed to examine the ARCH effects where the null hypothesis is there is no ARCH effect. The over-identifying restriction test is the differences in J-stats where the null hypothesis is the over-identifying restrictions are valid. For TSLS and GMM, the structural determinants of inflation from Moshiri and Cameron (2000) are used as instrument. The instruments are: $\hat{y}$, $\hat{m}_t$, π_t and *oil*. The definition of these determinants are given in "Appendix"

*, ** and *** represent 10, 5 and 1 % level of significance respectively

Table 10 The quadratic RPV model

	Coefficient						
	OLS 1	OLS 2	OLS 3	OLS 4	OLS 5	OLS 6	TSLS
Intercept	0.208*** [3.338]	0.343*** [5.020]	0.216*** [2.954]	0.227*** [2.844]	0.265*** [3.070]	0.296*** [4.079]	0.364** [2.472]
rpv (−1)	0.108 [1.540]	0.093 [1.320]	0.164**[1.956]	0.050 [0.633]	0.007 [0.095]	0.117* [1.733]	0.129 [0.724]
$E(\pi_t)$	−0.175 [−1.247]	−0.607*** [−3.795]	−0.109 [−0.619]	−0.182 [−1.285]	−0.269* [−1.833]	−0.455** [−2.559]	−0.555* [−1.774]
$[E(\pi_t)]^2$	0.213** [2.084]	0.541*** [4.675]	0.152 [1.365]	0.229*** [2.917]	0.332*** [3.779]	0.425*** [3.139]	0.453* [1.702]
UIP	0.600*** [11.432]	0.613*** [14.705]	0.611*** [8.804]	0.588*** [11.099]	0.598*** [11.80]	0.627*** [16.039]	0.655*** [6.722]
UIN	0.219*** [20.83]	0.212*** [3.245]	0.195** [0.2.038]	0.257*** [4.062]	0.268*** [4.571]	0.234*** [4.405]	0.178* [1.735]
VAR_UN	0.618 [1.613]	0.019 [0.703]	0.040 [0.764]			0.005 [0.222]	−0.023[−0.317]
CVARI				0.051 [0.556]	0.036 [0.430]		
Threshold π	0.41	0.56	0.36	0.40	0.41	0.54	
R^2	0.60	0.65	0.59	0.61	0.63	0.66	0.65
Adj. R^2	0.58	0.64	0.57	0.59	0.61	0.64	0.63
Standard error	0.258	0.241	0.261	0.254	0.249	0.240	0.244
F-statistics	30.337***	38.007***	29.485***	31.787***	33.855***	38.904***	8.629***
AR(2) test	F(2,119) = 0.287 (0.75)	F(2,119) = 0.259 (0.77)	F(2,120) = 2.463 (0.08)	F(2,120) = 0.587 (0.55)	F(2,118) = 0.396 (0.67)	F(2,120) = 0.667 (0.52)	$\chi^2(2) = 0.432$ (0.81)[a]
ARCH(1) test	F(1,125) = 0.279 (0.59)	F(1,125) = 2.038 (0.16)	F(1,126) = 1.563 (0.21)	F(1,126) = 1.468 (0.22)	F(1,124) = 0.630 (0.42)	F(1,126) = 2.107 (0.14)	F(1,119) = 2.650 (0.11)
Heteroskedasticity	F(6,121) = 2.597 (0.02)	F(6,121) = 1.878 (0.09)	F(6,122) = 7.354 (0.00)	F92,126) = 2.157 (0.11)	F(6,120) = 1.387 (0.22)	F(6,122) = 1.980 (0.07)	F(6,115) = 2.129 (0.06)
Chow test	F(60,61) = 1.448 (0.07)	F(60,61) = 1.121 (0.33)	F(60,62) = 1.564 (0.04)	F(61,61) = 1.513 (0.05)	F(61,59) = 1.532 (0.05)	F(60,62) = 1.138 (0.30)	F(7,108) = 1.834 (0.09)

Table 10 continued

	Coefficient						
	OLS 1	OLS 2	OLS 3	OLS 4	OLS 5	OLS 6	TSLS
H_0: UIP = UIN	F(1,121) = 27.321 (0.00)	F(1,121) = 39.06 (0.00)	F(1,122) = 17.826 (0.00)	F(1,122) = 22.918 (0.00)	F(1,120) = 24.153 (0.00)	F(1,122) = 61.963 (0.00)	F(1,115) = 21.235 (0.00)
Endogeneity test (Differences in J-stats)							1.774 (0.94)

This table reports the parametric OLS estimation of Eqs. (11) & (12) in the main text. The dependent variable is the relative price variability (rpv). Explanatory variables are: lagged rpv, expected inflaton (E(inf)), square value of the expected inflation $E(\pi_t)^2$, absolute value of the unexpected inflation when it takes positive values (and zero otherwise) (UIP), absolute value of unexpected inflation when it takes negative values (and zero otherwise) (UIN), conditional variance of inflation (CVARI), 6-month variance of unexpected inflation (VAR_UN). All of the expected and unexpected value of inflation is derived from five models: (a) Moshiri and Cameron (2000) for OLS 1, (b) Hossain (2002) for OLS 2, (c) SARIMA model for OLS 3 (4) GARCH (1,1) model using Fielding and Mizen (2008) where the lag length of inflation is used 2 and 4 respectively for OLS 4 and OLS 5 and (5) IMF (2007) for OLS 6. The following test statistics are reported: (a) AR(2) = LM test for 2nd order residual autocorrelation when the null hypothesis is there is no serial autocorrelation; (b) ARCH(1) = LM test for 1st order ARCH test when the null hypothesis is there is no ARCH effects; (c) Heteroskedasticity test = Breusch-Pegan Heteroskedasticity test when the null hypothesis is there is no heteroskedasticity; (4) Chow test = Chow forecast test for stability when the breakpoint is selected on the basis of residual plot. H_0: UIP = UIN: this test is performed whether the coefficient of UIP and UIN is equal. The figures in the square brackets are t-values and in the first brackets are p values. For TSLS, the structural determinants of inflation from Moshiri and Cameron (2000) are used as instrument. The instruments are: $\hat{y}$, $\hat{m}_t$, π_t and *oil*. The definition of these determinants are given in "Appendix"

*, ** and *** represents 10, 5 and 1 % level of significance respectively

[a] Chi-square value of Breusch-Godfrey serial correlation LM test statistics in case of TSLS

the hypothesis of both the Lucas-Barro and the Hercowitz-Cukierman models. However, $E(\pi_t)^2$ is statistically significant and supports the menu-cost model.

OLS 2 and OLS 3 show the estimation results of Eq. (11) using the decomposition of inflation from Hossain (2002) and the SARIMA model, respectively.[10] The forecasted inflation from Hossain's (2002) model passes the test of the quadratic relationship, but the SARIMA model does not, although the coefficients have predictive signs. Both models reject the Hercowitz-Cukierman version of the extraction model, but Hossain's (2002) model does not reject the menu-cost model. Each of the models reject the hypothesis of the Lucas-Barro model, as the coefficient of σ_t^{six} is statistically insignificant.

OLS 4 and OLS 5 show the estimation results where decompositions use the GARCH(1,1) model, adapted from Grier and Perry (1996). Each of the models produces the same conditional variance, but the decompositions are different due to the lag structure of $E(\pi_t)$. Both columns establish the U-shaped relationship, as the coefficient of $E(\pi_t)^2$ is positive and highly significant, and both reject the Hercowitz-Cukierman version of the extraction model and the Lucas–Barro model, but not reject the menu-costs model, as the coefficients of $E(\pi_t)^2$ are statistically significant. OLS 6 shows the estimation of IMF's (2007) model of forecasting. The result is similar to the previous ones, but is the most efficient in terms of coefficient significance.

To remove the endogeneity bias, we apply the TSLS method to OLS 2, which is represented in the last column. We assume that inflation may behave as an endogenous variable due to the large global shock from food prices as food has a high weight in the CPI. We use the structural determinants of π_t and its lag as variables.[11] Our TSLS estimation also supports the previous results, where the test of endogeneity (over-identification restriction test) suggests that the instruments are valid and π_t is exogenous.

From the above estimation, we also calculate the threshold level of inflation, which ranges from 0.36 to 0.56 % monthly (4.32–6.72 % annually). This also indicates that threshold inflation is positive, which is crucial to the relationship between RPV and inflation (Van Hoomissen 1988; Hartman 1991; Choi 2010).

The above analysis generates the conclusion that none of the three models (i.e., Hercowitz-Cukierman, Lucas-Barro, and the menu-cost model) predicts the pattern of coefficients in Table 10. Table 11 shows a comparison of the estimated models with the theoretical model adopted from Aarstol (1999). Not all the specifications support the Hercowitz-Cukierman model. Additionally, our empirical data fail to explain all of the three models simultaneously. Generally, it can be concluded that our sample supports the menu-cost model, and the effect of unexpected inflation on RPV is stronger than that of expected inflation.

Although the sample does not support the Lucas-Barro model, Grier and Perry (1996) suggest to test whether the data supports this model. They estimate a regression of RPV on expected inflation and inflation uncertainty. This specification yields

[10] See the "Appendix" for the lists of inflation models.

[11] See Moshiri and Cameron (2000).

Table 11 Estimated coefficients in comparison with the predictions of the theoretical models

	Model			Coefficients					
	Menu-Cost	Barro-Lucas	Hercowitz-Cukierman	OLS 1	OLS 2	OLS 3	OLS 4	OLS 5	OLS 6
Expected inflation	+	0	0	+	+	0[a]	+	+	+
Unexpected inflation (+)	0	0	+	+	+	+	+	+	+
Unexpected inflation (−)	0	0	+	+	+	+	+	+	+
Inflation uncertainty	0	+	0	0[a]	0[a]	0[a]	0[a]	0[a]	0[a]

[a] The coefficient estimate on the variable is positive but insignificant

$$\begin{aligned} RPV_t &= \underset{[2.71]}{0.224} + \underset{[3.95]}{0.183}E(\pi_t)^2 + \underset{[2.37]}{0.435}\sigma_t \quad \text{Adj. } R^2 = 0.18, \end{aligned} \tag{15}$$

where t-statistics are given in square brackets. This equation is a restricted form of Eq. (11). The above estimation shows that the coefficients of conditional variance and expected inflation are positive and significant. This suggests that the sample supports both the Lucas-Barro model and the Hercowitz-Cukierman signal extraction model.

Tables 10 and 11 also confirm that none of the estimated models completely support all three models. The Hercowitz-Cukierman model predicts the relationship between unexpected inflation and RPV. Our results unexpectedly indicate that the coefficient of UIP is greater than that of UIN, suggesting the rejection of the Hercowitz-Cukierman signal extraction model. This implies that the relationship between RPV and inflation is incomplete. The fact that the coefficient estimate on negative unexpected inflation is smaller than the coefficient estimate on positive unexpected inflation suggests that, given the initial position of high ongoing inflation, a policy of disinflation to a lower rate of the ongoing inflation tends to promote welfare. This is because a higher RPV caused by negative unexpected inflation during the transition period tends to be outweighed by the perpetually lower RPV at the eventually lower rate of ongoing inflation. This finding is consistent with Aarstol (1999). Consequently, we can conclude that unexpected inflation influences RPV in Bangladesh.

4.3.4 An alternative measure of RPV

Various studies on RPV and inflation focus on measures of RPV that omit food and/or push prices in an attempt to control for supply shocks (Fischer et al. 1981; Grier and Perry 1996; Aarstol 1999). The value of this type of RPV is debatable, as it disregards the notion of the Lucas-Barro signal extraction model, the measurement failing to incorporate local shocks. Nevertheless, we consider this measurement to check the effect of inflation uncertainty on the new RPV. We show the results from Fischer et al.'s

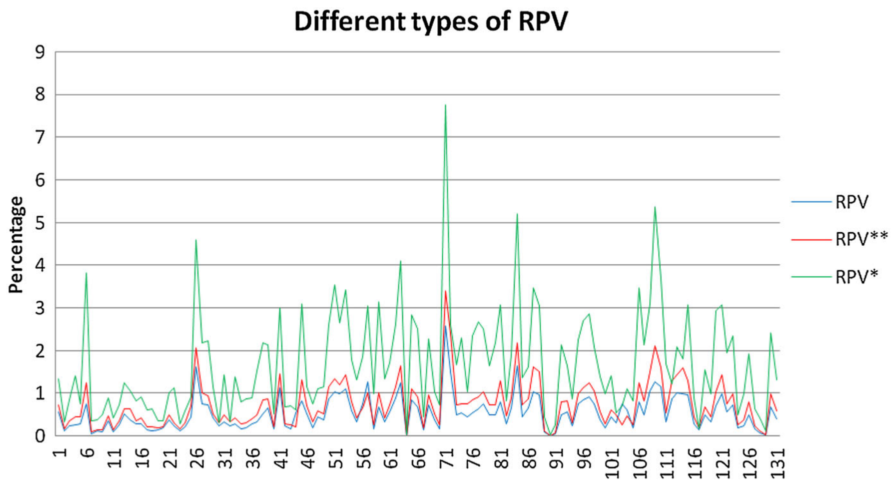


Fig. 4 Different types of RPV in Bangladesh: RPV is estimated based on Eq. (4). RPV* is estimated by excluding food and energy prices and RPV** is estimated by excluding only food prices

(1981) experiments in Table 6, where there were no changes in the event of a new RPV (henceforth RPV*). We now estimate Eq. (12) in the case of RPV* and RPV**, where RPV* excludes food and energy prices, whereas RPV** excludes only food prices.

Figure 4 shows RPV, RPV*, and RPV**. RPV is less volatile than RPV* and RPV**, which contrasts with Fischer et al.'s (1981) suggestion that supply shocks may not have significant effects on RPV. We estimate this hypothesis by excluding energy prices from the RPV calculation. Table 12 shows the estimation of the RPV* equation compared with the RPV one. There is no difference between the models in terms of coefficient signs. Additionally, both models ensure the U-shaped profile. The coefficients of UIP and UIN are positive and statistically significant, although the RPV* model fails to support the Hercowitz-Cukierman signal extraction model. The coefficient of conditional variance is also insignificant and rejects the Lucas-Barro signal extraction model.

However, the coefficients of RPV* are greater than those of RPV. This suggests that exclusion of the energy sector amplifies the effects of inflation on RPV. Generally, the inflation-RPV link is weaker if the energy sectors are excluded from the RPV measure. Additionally, when we exclude food prices from RPV, the new RPV, RPV**, is also affected by inflation and the coefficients are greater than for the RPV model, but lower than for the RPV* one. However, our result contrasts with results for developed countries (Parks 1978; Bomberger and Makinen 1993; Nautz and Scharff 2005). Fischer et al. (1981) found that energy price shocks dominate the behavior of RPV in Germany and Japan. In the case of Bangladesh, we found that RPV is still dominated by inflation even without considering energy and food price shocks. We can say that consumers are trying to smooth their consumption patterns after realizing the shock in food and energy prices. This suggests that, without energy and food, RPV* is more volatile than the general RPV. Following Bloom and Ratti (1985), we examine this case by regressing both food

Table 12 Alternative RPV and inflation quadratic relationship

Model	Dep. var.	Intercept	RPV_{t-1}	$E(\pi_t)$	$E(\pi_t)^2$	UIP	UIN	σ_t	Adj R^2	DW
1	RPV	0.265*** [3.070]	0.007 [0.095]	− 0.269* [−1.833]	0.332*** [3.779]	0.598*** [11.80]	0.268*** [4.571]	0.036 [0.430]	0.61	1.89
2	RPV*	0.860*** [3.213]	−0.014 [−0.186]	−0.568 [−1.167]	0.710* [1.958]	1.912*** [15.470]	0.917*** [5.163]	0.058 [0.219]	0.58	2.04
3	RPV**	0.385*** [3.883]	0.072 [0.958]	−0.404** [−2.402]	0.469*** [3.638]	0.834*** [11.879]	0.395*** [4.697]	−0.068 [−0.599]	0.62	1.91

Figures in square brackets are t ratios

DW Durbin Watson d-statistics

*, ** and *** represent 10, 5 and 1 % level of significance respectively

inflation and its absolute value, and energy inflation and its absolute value on RPV. We find that the positive value of food inflation dominates all other coefficients.[12] This suggests that food price inflation and non-accommodating fiscal and monetary policies generate disturbances in RPV, not considering energy prices.

4.4 Rolling regression

Table 10 shows that almost no models pass the structural change test of Chow (1960). Parameter stability would require the form of the $g(.)$ function to be close to a perfectly linear quadratic relationship. Even our reexamination of Parks (1978) and Fischer et al. (1981) does not support a constant parameter across the sample. This instability suggests that the parameterization of the quadratic relationship is too restrictive.

It is important to determine whether this "arbitrary" choosing of breakpoints is robust. The rolling regression technique captures time variation in the relationship without imposing any prior restrictions on the timing of the breakpoints. The advantage of this approach is the greater flexibility in detecting structural changes over time, by allowing for each rolling sample to have a completely different estimate (O'Reilly and Whelan 2005).

Having the quadratic specification, we estimate the following parametric model that accommodates both inflation and lagged RPV as regressors[13]:

$$RPV_t = \alpha_0 + \sum_{h=1}^{p} \alpha_h RPV_{t-h} + \beta_1 \pi_t + \beta_2 \pi_t^2 + \sum_{j=1}^{q} \delta_j \pi_{t-j} + \varepsilon_t. \tag{16}$$

In the above specification, inclusion of both π_t and π_t^* is crucial for determining the time-varying behavior of the relationship between RPV and inflation, which can be captured by their instability over rolling samples.

We have established the direction of the RPV-inflation relationship curvature. If the function is U-Shaped, the sign of β_2 must be positive and statistically significant (Choi 2010). Additionally, if β_2 approaches 0, the functional form becomes linear. In this case, the overall relationship between RPV and inflation is solely determined by β_1. Furthermore, the minimum point of a U-shaped curve provides a useful way to summarize the function. Using Eq. (16), the minimum point of the U-shaped function is reached when

$$\pi^* = \frac{-\beta_1}{2\beta_2}, \tag{17}$$

where π^* is the threshold level of inflation, calculated in Table 10. In this case, we investigate the stability of π^*. The expected sign of π^* depends on the sign of β_1. If $\beta_1 < 0$ (it must be negative for a U-shaped curve), then π^* is positive, and negative if $\beta_1 > 0$. As a result, the relationship is U-shaped around positive inflation ($\pi^* > 0$). However, if $\beta_2 = 0$, π^* is not properly defined, as it explodes. As such, the

[12] The results can be provided upon request.

[13] We add only the inflation rate rather than the expected inflation. We also use the expected inflation, which generates similar results.

stability of the relationship between RPV and inflation can be tracked by a time varying behavior of π^*. Additionally, the marginal effect of inflation on RPV in Eq. (17) is not constant as in the linear model, but varies with the inflation rate. Consequently, the marginal effect can be approximated by

$$\frac{\Delta RPV_t}{\Delta \pi_t} \approx 2\beta_2 \pi_t + \beta_1. \tag{18}$$

Apparently, the change in RPV for a one-unit change in π_t depends not only on β_1 and β_2, but also on the value of π_t.

Figure 5 reports the results of the rolling regression by showing the estimates of β_1 and β_2 from a sequence of rolling samples. We perform a rolling regression with a fixed window, in which sample size does not change across different regressions. We fixed 50 observations for the first regression, and subsequently dropped the first observation and continued to add another until the whole data is covered.[14]

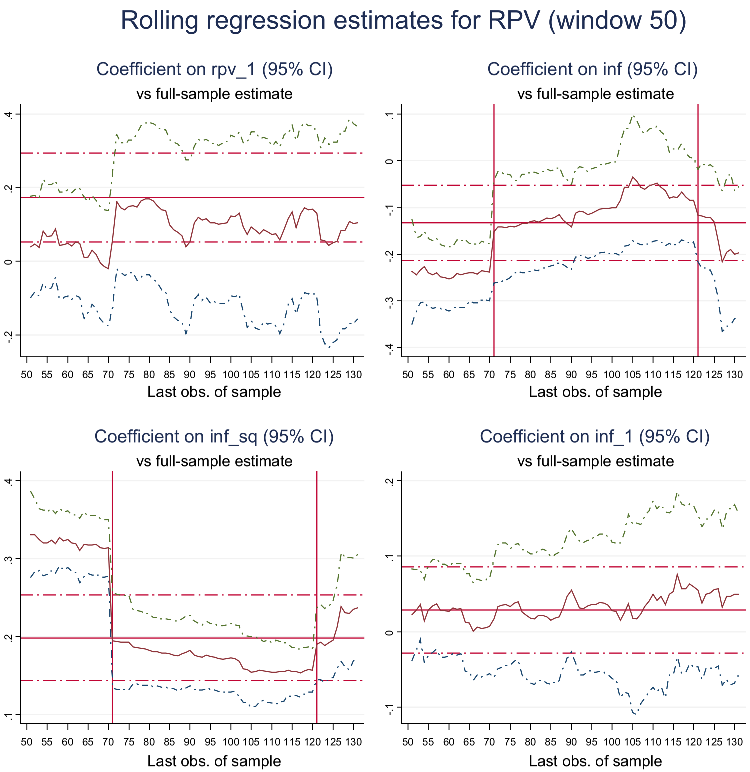
Panel (A) shows the results of the rolling regression representing coefficients of α_1, β_1, β_2, and δ_1. The solid line shows the rolling estimation, and the dashed line shows the 95 % confidence interval based on the heteroskedasticity–autocorrelation robust standard errors with pre-whitening. The numbers on the horizontal axis represent the ending point of each 50-month window. For instance, window 40 captures the subsample period of 1–40.

A number of interesting conclusions result from this plot. Both $\hat{\beta}_1$ and $\hat{\beta}_2$ exhibit significant variations over time, and share the common timing of breakpoints. For $\hat{\beta}_1$ and $\hat{\beta}_2$, the breakpoint emerges at point 71, which is the observation for June 2008, similarly to the earlier breakpoint analysis. Another breakpoint is happening at point 121 (i.e., August 2012). Historically, the first breakpoint is pertinent to this analysis. Bangladesh experienced a high price level increase after the change of political regime in January 2007, although the government was a non-political caretaker government. At the same time, the world economy experienced an increase in the price of rice, where the CPI basket of Bangladesh is mostly composed of food items. This exacerbated the price level and the country faced a price hike, altering both fiscal and monetary policies in Bangladesh to reduce the price level. It reduced the price level, although not by much.[15] The 95 % confidence intervals of these coefficient estimates are of similar magnitude across sample periods, and the upper and lower bounds move closely with the coefficients.

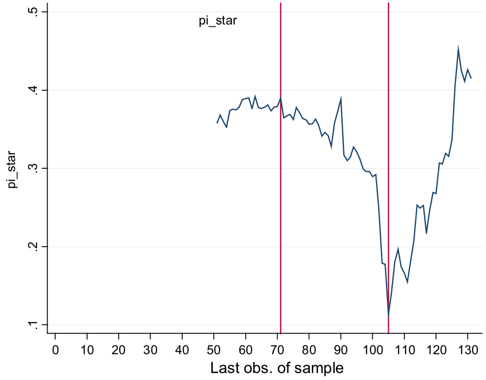
Another important feature of this estimation is the sign of $\hat{\beta}_1$ and $\hat{\beta}_2$ coefficients. The upper right-hand corner of Panel (A) shows $\hat{\beta}_1$, a value of the coefficient below

[14] We also used this methodology for the 25–45 window size with similar results.

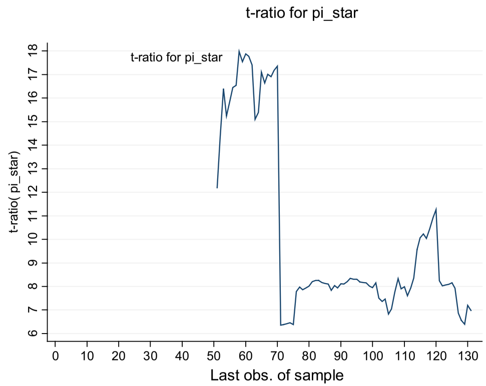
[15] From July 2011 to December 2012, Bangladesh observed a downward trend for both food and general inflation, but an upward trend for non-food inflation. Although non-food inflation is increasing, general inflation is decreasing, as a major share of income is spent on food. On the other hand, during the mentioned periods, the central bank was trying to maintain the inflation rate at 7.5 %. In their monetary policy statement (MPS) of H1-FY2013 (first half-yearly MPS), the central bank announced domestic credit cuts to achieve the targeted inflation rate in accordance with the Ministry of Finance (through fiscal budgeting). These steps ensured a relatively downward inflation trend in relation to the previous periods of high inflation.



Panel A: Rolling estimate of equation (12)



Panel (B): Rolling estimate of π^*



Panel (C): Rolling estimate of t-ratio for π^*

Fig. 5 Rolling estimation of equation (16); Window (50) zero suggesting that $\hat{\beta}_1$ is consistently negative and statistically significant for almost all observations. Having some insignificant coefficients, $\hat{\beta}_1$ shows a negative sign for all regressions. The lower left-hand corner of Panel (A) shows the coefficient of π_t^2. The sign of $\hat{\beta}_2$ is consistently positive and statistically significant, ensuring a U-shaped profile. Similar to $\hat{\beta}_1$, we also get two breakpoints for $\hat{\beta}_2$. Consequently, $\hat{\pi}^*$, as defined in Eq. (17) is consistently positive (Panel (B)). $\hat{\pi}^*$ ranges from $0.1 < \hat{\pi}^* < 0.5$ (1.2 to 6 % annually). In Table 10, we calculate the threshold level of inflation, which ranges from 0.36 to 0.56 % monthly (4.32 to 6.72 % annually). This also indicates that the threshold inflation is positive, which is crucial to the relationship between RPV and inflation (Van Hoomissen 1988; Hartman 1991; Choi 2010). Moreover, $\hat{\pi}^*$ varies significantly over time. Although it is declining until observation 105 (April 2011), it has stayed consistently above 0 over the period. The 95 % confidence interval is also pointing to a non-negligible gap of π^* from 0. In Panel (C), the t-ratio for $\hat{\beta}_2$ has been estimated, where all of the ratios are significant at the 5 % level of significance.

Overall, the estimated results suggest that the relationship between RPV and inflation is U-shaped around a positive inflation rate, which is significantly different from 0. Since some theoretical models imply that the U-shaped relationship will be centered at $\pi^* = 0$, our empirical results depart from that notion, suggesting a better theoretical model to match the data.

4.5 Testing for multiple structural breaks

Although rolling regressions predicted the breakpoint, for robustness, we implement a formal test for structural changes developed by Bai and Perron (1998, 2003). One of the advantages of this test is that it locates the structural breaks endogenously from the data, without assuming any prior knowledge of the potential break-dates and the number of breaks. The number of breaks, their timing, and the constant are estimated using a series of sequential Wald tests. We consider a multivariate setting, similar to Eq. (16), with a general error process, in which both heteroskedasticity and autocorrelation are allowed.[16] Following these guidelines, the break is assumed not to occur during the initial or the final 15 % of the sample period in testing structural breaks. Moreover, the maximum number of breaks is set to five and the minimum regime size is set to 5 % of the sample.

Table 13 reports the estimated dates for structural breaks in the relationship between RPV and inflation. The 95 % confidence interval is computed using robust standard errors based on a quadratic spectral kernel heteroskedasticity and autocorrelation (HAC) estimator with an AR(1) pre-whitening filter. Bai-Perron's multivariate breakpoint analysis identifies one structural break based on the Schwarz criterion, although the LWZ (Liu et al. 1997) criterion suggests zero break-

[16] We consider this equation based on Choi (2010). The use of Eqs. (7) or (8) prevents us from conducting a Bai-Perron test, as the number of parameters is too large to identify breakpoints. For details, see Bai and Perron (2003).

Table 13 Results of Bai-Perron Test

Breaks	# of coefficients	Sum of Sq. residuals	Log-L	Schwarz[a] criterion	LWZ[a] criterion
0	5	6.119	14.182	−2.868	**−2.697**
1	11	4.817	29.730	**−2.883**	−2.504
2	17	4.321	36.790	−2.767	−2.178
3	23	3.923	43.080	−2.639	−1.837
4	29	3.677	47.281	−2.479	−1.461
5	35	3.502	50.446	−2.303	−1.066
Estimated break dates					
1	**2008M06**				
2	2008M06	2011M12			
3	2004M09	2008M06	2011M12		
4	2004M09	2008M01	2009M008	2011M12	
5	2004M07	2006M03	2008M05	2010M03	2011M12

Compare information criteria for 0 to M globally determined breaks. Breakpoint variables: $constant, RPV_{t-1}, \pi_t, \pi_t^2, \pi_{t-1}$. Breakpoint options: trimming 0.15, max breaks: 5

[a] Minimum information criterion values displayed with bold figure. Multiple breakpoint tests

Table 14 Dummy variable regression with break dates

Variable	Coefficient	t ratio
Intercept	0.288***	8.88
RPV_{t-1}	0.132**	2.44
π_t	−0.185***	−5.01
π_t^2	0.261***	12.56
π_{t-1}	0.021	0.80
$D1_{2008M06}$	−1.400***	−4.21
$D2_{2012M08}$	−0.030	0.15
R^2	0.73	
Adj. R^2	0.72	

*, ** and *** represent 10, 5 and 1 % level of significance respectively

dates. The present study utilizes the Schwarz criterion. The identified break-date is 2008M06, which is similar to our previous analysis.

To substantiate the results of the Bai-Perron test, we consider another break-date as 2012M12. Using two break-dates, we estimate Eq. (16) with two additional dummy variables, D1 and D2. Table 14 shows the estimated results. Our results suggest that the break-date 2008M06 is statistically significant, while the break-date 2012M08 is not. We also estimate a separate regression for the break-date 2008M06. In both cases, the estimated regression ensured the U-shaped profile of the RPV-inflation relationship,[17] and that the coefficients of $\hat{\beta}_1$ and $\hat{\beta}_2$ are statistically significant. In both regressions, the values of π^* are 0.342 and 0.345 %, respectively, being almost equal (4.104 and 4.14 % yearly).

[17] The results can be provided upon request.

Overall, the results can be viewed as being supportive of the findings in the previous sections, as well as consistent with the maintained breakpoint in the sub-sample analysis. Additionally, the three different econometric tools lead to very similar conclusions about the functional relationship between RPV and inflation.

5 Conclusion

A widespread conclusion in economic literature is that RPV is positively correlated with inflation, such that higher inflation levels are linked with increased cross-sectional dispersion in relative prices. Earlier studies are based on the monotonic transformation of the RPV-inflation relationship, that is, lower inflation reduces the level of dispersion in relative prices. Many of them advocate reducing this variability by reducing high inflation, as it distorts the decision-making process of both producers and consumers (Ermisoglu et al. 2014). However, this monotonic relationship will produce inefficient results if the actual functional form of the RPV-inflation relationship is unknown. If the relationship is non-monotonic, beyond a certain threshold of inflation, the relationship will be reversed. This requires the monetary policy to consider its implications for RPV. Additionally, the decomposition of inflation also plays an important role in this relationship. Moreover, the stability of that relationship is important for the execution of the long-run monetary policy in Bangladesh.

Consequently, we use monthly personal consumption expenditure data from July 2002 to June 2013 in Bangladesh to construct inflation and relative price variability measures. In the first step, we reexamine earlier studies (Parks 1978; Fischer et al. 1981), which established the positive linear relationship between RPV and inflation, and RPV and expected inflation, and thus, the theories of the RPV-inflation relationship (menu-cost model, signal extraction model, etc.). However, based on recent literature, we extend our analysis to a core analysis of the functional relationship. In this stage, we use a semiparametric methodology to discriminate between the alternative functional forms, without imposing a particular (parametric) result in advance. With robust analysis and different samples, we find evidence of nonlinearity, with an approximately quadratic functional form at low to moderate rates of inflation, consistent with existing theoretical menu-cost models. We subsequently extend our analysis to a quadratic parametric specification. To analyze the effect of inflation decomposition, we apply various structural and time series models of inflation forecasting. The quadratic parametric specification yields an interesting result for Bangladesh. Among the models used, IMF's (2007) model for expected inflation produces the best results in terms of coefficient significance, although the expected inflation rates derived from other models support the quadratic specifications. This also implies that the RPV-inflation quadratic specification does not support the Hercowitz-Cukierman signal extraction model. Additionally, the specification weakly supports the Lucas-Barro signal extraction model. These findings also resonate with supply shocks. We found that the nature of the RPV-inflation relationship is distinct in Bangladesh from that in advanced economies, with a supply shock not producing extreme effects. Food price inflation and non-accommodating macroeconomic policies distort the stationary characteristic of RPV, as well as inflation. Our parametric methodology suggests that there is a positive optimal rate of inflation for which RPV is minimized. The threshold level of inflation ranges from 4.32 to 6.72 % annually.

Although the specification is nonlinear, it is not stable over time. This point is crucial for policy purposes. If the quadratic relationship is not stable over time, the threshold level of inflation will change, which could, in turn, affect the monetary policy. Our rolling regression estimate suggests that the threshold level of inflation decreased over time (until April 2011), and increased after that. The level and quadratic coefficients of inflation also have their expected signs,, but within the confidence interval bands. Although our rolling regression suggests two break-points, Bai-Perron's (2003) breakpoint test suggests only one.

Historically, Bangladesh has faced moderate levels of inflation. Our analysis suggests that unexpected inflation matters for RPV more than expected inflation does. This suggests that the influence of unexpected inflation disappears if a credible monetary policy stabilizes the optimal rate of inflation at a low level. In a high-inflation environment, monetary authorities can improve welfare through disinflationary policies that lower RPV. However, we establish that this policy will have no significant effect if the relationship is U-shaped. Therefore, a set of optimal rates of inflation is necessary for different time periods, as the relationship may be unstable over time. This may ensure a welfare gain by implementing disinflationary policies according to the time period. One possible caution is that the degree of price rigidity changes with the inflation regime and monetary policy framework. Consequently, a thorough understanding of the relationship between RPV and inflation is of great importance for policymaking.

Appendix

Models of inflation forecasting

1. Structural model:

The reduced form of inflation:

$$\pi_t = \beta_0 + \beta_1 \hat{y}_{t-1} + \beta_2 \hat{m}_t + \sum_{i=0}^{n} \rho_i \pi_{t-1-i} + \beta_3 oil_t + \eta_t \tag{19}$$

2. Hossain's (2002) model of inflation in Bangladesh:

The reduced form of inflation:

$$\ln(P_t/P_{t-1}) = -(1-\phi)\gamma\beta_0 + \phi\ln(ER_t/ER_{t-1}) \\ + (1-\phi)\gamma\Delta\ln m_{t-1} - (1-\phi)\gamma\Delta\ln P_{t-1} - (1-\phi)\gamma\beta_1\Delta\ln y_t + (1-\phi)\upsilon_t \tag{20}$$

This is an estimable model of inflation for a developing country like Bangladesh. Here, we use

P = General CPI, ER = TK/\$ exchange rate, m = M2/P, y = industrial production index. All series are monthly.

3. Autoregressive Integrated Moving Average (ARIMA) Model:

Using Box and Jenkins (1976) methodology, we estimate the model of inflation. After inspecting the autocorrelation and partial autocorrelation of inflation, we conclude that the series contain seasonal fluctuation. So, we estimate Seasonally Adjusted ARIMA (SARIMA) model.

4. Generalized Autoregressive Conditional Heteroskedasticity (GARCH) model:

In order to decompose inflation into expected and unexpected components, we follow a GARCH model (Fielding and Mizen, 2008). We estimated a GARCH (1, 1) model. Consider the following AR (p) model with time varying conditional variance:

$$\text{Mean equation:} \inf_t = \alpha_0 + \sum_{i=1}^{p} \alpha_{1i} \inf_{t-i} + \varepsilon_t$$

$$\varepsilon_t \sim (0, \sigma_t)$$

$$\text{Variance equation:} \sigma_t = \rho_0 + \rho_1\sigma_{t-1} + \rho_2\varepsilon_{t-1}^2 \tag{21}$$

We use the number of lag from 2 to 4. An increase in the number of lags decreases the degrees the degrees of freedom. Again, the Akaike Information Criteria (AIC) selects 2 lags while the Schwarz Information Criteria (SBC) select 4 lags of inflation. We derived UIP and UIN for all models. We also derived two series of conditional variance from GARCH (1, 1) models.

5. IMF (2007) model of inflation for Bangladesh:

$$\pi_t = \pi_{t-1} + m_{t-1} + ER_t + u_t \tag{22}$$

List of variables used in inflation forecasting

Variable	Definition	Unit	Source
$\hat{y}$	Output gap: deviation of log of output from its natural value. In this case, we use a proxy for output as our observations are monthly. We use industrial production index	Index	Monthly Economic Trends, Bangladesh Bank
y	Output. It is measured by monthly industrial production index (IPI). The natural logarithmic value is used for estimation	Number	Monthly Economic Trends, Bangladesh Bank
m_t	Growth rate of real M2 money supply. We use M2 money and corresponding price level for calculation of real money supply	Percentage	Monthly Economic Trends, Bangladesh Bank
$\hat{m}_t$	Money gap: deviation of the log of real money supply (broad money) from its natural value. We use M2 money and corresponding price level for calculation of real money supply	Number	Monthly Economic Trends, Bangladesh Bank
oil	Growth rate of monthly crude oil price in international market	Percentage	US Energy Information Administration.
ER	Exchange rate of Taka per unit of US Dollar	Ratio	Monthly Economic Trends, Bangladesh Bank
P	Consumer Price Index	Index	Monthly Economic Trends, Bangladesh Bank

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