On the Asymmetric U-shaped Relationship between Inflation, Inflation Uncertainty and Relative Price Skewness in the UK

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Abstract

We investigate asymmetries in the relationship between the aggregate rate of inflation and the second and third central moments of the cross-sectional distribution of relative prices using a modified Calvo pricing model with regime-dependent price rigidities. Calibration experiments using realistic parameterisations reveal that the inflation-standard deviation and inflation-skewness relationships exhibit U-shaped asymmetries around the historical mean rate of inflation. We conclude that monetary policy should target a level of inflation proximate to the (common) minima of these nonlinear relationships and that core inflation measures exclude much of the information contained in the higher moments of the distribution of price changes.

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1 INTRODUCTION

A substantial literature has explored the relationship between inflation and inflation uncertainty, which is widely held to be positive following the seminal contributions of Friedman (1977) and Cukierman and Meltzer (1986).[1] The menu-cost theory advanced by Ball and Mankiw (1995) supports a positive relationship between aggregate inflation and the standard deviation of relative price changes (often termed *relative price variability*, *RPV*). Moreover, they propose a positive association between the mean and the skewness of relative price changes (henceforth *relative price skewness*, *RPS*). They contend that firms facing menu-costs will only adjust their prices if the benefit outweighs these costs. Therefore, if the distribution of relative price changes exhibits positive (rightward) skewness, then mean inflation is likely to increase as a small group of firms engage in significant price increases while a large group of firms desire price decreases too small to be economically viable. Mutatis mutandis in the case of leftward skewness. The menu-cost model implies a fundamental interplay between RPV and RPS, as RPV should have no systematic effect on mean inflation if relative price changes are symmetrically distributed, at least in the absence of trend inflation (see Aucremanne et al., 2003 for a detailed discussion). An alternative view is offered by Tobin (1972), whose discussion of exogenous downward nominal rigidities in labour and product markets implies a negative mean-RPS relationship. This pattern will arise naturally where more rapid inflation facilitates relative price adjustments in the presence of strong downward nominal rigidity.

Until recently, the empirical literature has shed little light on these divergent views. The majority of existing studies have analysed aggregate (national) inflation rather than exploiting the informational content of disaggregated (sectoral) data. Even where studies have employed sectoral data, the focus has typically been on the mean-RPV relationship. While the analysis of time-varying higher-order moments is well-established in a range of macroeconomic and financial settings (e.g. Hansen, 1994; Harvey and Siddique, 2000;

Dittmar, 2002; Jondeau and Rockinger, 2003; Brooks, 2005), relatively little research effort has been devoted to similar analyses of the mean-RPS relationship.[2]

Recently, however, a growing literature on time-variation and nonlinearity has shed new light on the relationship between aggregate inflation and volatility. Choi (2010) provides robust evidence of time-variation in the inflation-RPV relationship in the U.S. and Japan on the basis of a modified Calvo model, finding a monotonic association in periods of high inflation but a U-shaped relationship during the 'Great Moderation'. Similarly, Fielding and Mizen (2008) and Chen, Shen and Xie (2008) observe a U-shaped mean-RPV relationship in the US and in four Asian economies, respectively. Evidence in favour of a U-shaped mean-RPV relationship has also been derived from monetary search models employing Burdett-Judd pricing (Head, Kumar and Lapham, 2010; Becker and Nautz, 2010). In this case, the asymmetry arises from the interplay of firms' market power and consumers' search intensity at various rates of inflation. Specifically, increases in the rate of inflation can be welfare-improving at low levels of inflation as they increase search-intensity more than market power. However, beyond a threshold rate of inflation the reverse is true, and increasing inflation becomes welfare-reducing. Becker and Nautz estimate the value of this threshold to be approximately 3% in Europe.

Despite the weight of evidence supporting the nonlinear mean-RPV association, we are unaware of any systematic analysis of asymmetries in the relationship between mean inflation and RPS. We therefore contribute to this literature by developing a simple framework that utilizes both the informational content of the higher-order moments and the time-variation of the cross-sectional distribution, thereby overcoming the biases that could result from the presence of informational heterogeneity across sectors (Pesaran and Smith, 1995). Moreover, the possibility that the mean-RPS relationship may exhibit various nonlinearities in turn raises the possibility of a reconciliation between the seemingly conflicting views of Ball and Mankiw (1995) and Tobin (1972) in the sense that both mechanisms may operate under different conditions.

Based on our analysis of UK sectoral CPI inflation data over the period 1997m1-2008m2, we note that the first three cross-sectional moments are highly persistent. The mean of the sectoral inflation rates is correlated with both the cross-sectional variance and skewness, although the direction of the association is contingent upon the location of average inflation rates relative to (adaptive) inflation expectations. More specifically, we observe a U-shaped relationship between both mean and RPV and between mean and RPS, and a linear association between RPV and RPS. Developing upon Choi (2010), we conduct calibration experiments on the basis of a modified Calvo (1983) model with regime-switching price rigidities. Our results convincingly verify a U-shaped relationship between mean inflation and RPV under a wide range of parameterisations. Moreover, we note that the mean-RPS relationship is linear and positive where marginal costs display only mild persistence, but that it exhibits the characteristic U-shape at the higher levels of persistence observed in the UK data. By comparing the distribution of sectoral price changes implied by Choi's model and by our regime-switching model against the distribution of actual price changes, we find that our regime-switching mechanism is important in providing an accurate characterisation of the higher moments of the distribution of sectoral price changes.

We conclude that in an optimal inflation targeting framework, the target should be set in the neighbourhood of the minimum-RPV rate of inflation such that monetary policy innovations act to maintain price stability while also minimising inflation uncertainty. Inflation in this range is also consistent with a near-symmetric (mildly left-skewed) distribution of relative price changes, indicating an economy free from pricing distortions that may otherwise exert undesirable allocative effects and reduce economic cohesion. Finally, we conclude that policymakers should not rely unduly on core measures of inflation as they obscure a great deal of valuable information contained in the higher moments of the distribution of relative price changes.

The remainder of the paper is organized as follows. Section 2 reports stylized findings evident in the UK data. Section 3 elaborates upon our calibration-based experiments and Section 4 evaluates the performance of our DGPs in terms of their ability to replicate the distribution of observed price changes and their sensitivity to the use of core CPI data. Section 5 concludes.

2 STYLIZED FACTS

Our analysis of UK sectoral inflation draws heavily on Chaudhuri, Kim and Shin (2011, henceforth CKS).[3] We utilise CKS's dataset, which is based on subsector consumer price indices and the associated weights obtained from the Office for National Statistics. Working at the three-digit level and using data spanning the period January 1997 to February 2008, CKS construct a sample of 134 monthly observations. This sample spans the period of independent monetary policy in the UK leading up to the late 2000s Global Financial Crisis. It follows, therefore, that it relates to an approximately homogeneous monetary policy regime[4].

The mean inflation rate in our sample (measured across sectors and over time periods) stands at 1.5 percent while the median is 1.6 percent.[5] Inflation uncertainty measured by the cross-sectional standard deviation shows considerable time-variation, while inflation exhibits a positive skew in just 51 out of 134 months. In addition to this time-variation, we observe considerable heterogeneity across sectors. Liquid fuels experienced the highest rate of inflation over our sample (10.5 percent) and information processing equipment the lowest (-22.3 percent). Sectoral standard deviation ranges from 0.4 percent for social protection to 26.5 percent for liquid fuels. Inflation exhibits a positive skew in 49 of the 85 subsectors under scrutiny, reflecting the tendency toward positive skewness widely noted in the literature (e.g. Aucremanne et al., 2003).

Figure 1 reproduces part of CKS's Figure 2. A pronounced U-shaped relationship between both mean and RPV and also between mean and RPS is immediately apparent, while the RPV-RPS relationship appears approximately linear. To investigate these patterns more formally, we compute the correlation between the first three cross-sectional moments evaluated at each time period. We employ a simple two-regime threshold model, distinguishing between high and low inflation regimes according to whether current inflation exceeds the mean level measured through time (the latter can be viewed as expected inflation assuming adaptive expectations). Table 1 reports the results.

– Insert Figure 1 and Table 1 here –

The relationship between the mean and the higher order moments is positive in the high inflation regime and negative in the low inflation regime. The association between mean and RPV is measured at 0.32 and -0.21 in the high and low inflation regimes, respectively. The equivalent figures for the mean-RPS relationship are 0.29 and -0.20. These findings cast doubt on the widely held belief that the association between aggregate inflation and RPV is positive. Moreover, U-shaped nonlinearity in the mean-RPS relationship suggests that menu-cost considerations may dominate when inflation exceeds expectations but that Tobin's downward nominal rigidities play a more significant role when inflation is lower than expected. Lastly, we observe an approximately linear RPV-RPS relationship that remains positive irrespective of the regime, although it is considerably stronger in the high inflation regime. Given the similarity of the nonlinear forms of the mean-RPV and mean-RPS relationships, it follows from transitive reasoning that the RPV-RPS relationship should be approximately linear.

3 A MODIFIED CALVO MODEL WITH REGIME-DEPENDENT PRICE RIGIDITIES

Choi (2010) develops a modified Calvo model that describes the U-shaped mean-RPV relationship rather well, although he does not address nonlinearities in the mean-RPS relationship. In this section, we extend Choi's approach to accommodate this more general case. Choi considers the following calibration-based data generating process:

$$\log\left(\frac{P_{it}}{W_t}\right) = \lambda_i \log\left(\frac{P_{i,t-1}}{W_{t-1}}\right) - \lambda_i\left(\frac{1-\beta\rho}{1-\lambda_i\beta\rho}\right)\log\left(\frac{W_t}{W_{t-1}}\right),\ i=1,\dots,N;\ t=1,\dots,T, \tag{1}$$

where P_{it} denotes sectoral prices, β the discount factor, λ_i the degree of price rigidity (with $\lambda_i = 1$ indicating perfect rigidity), and W_t the common marginal cost generated by the following AR(1) process:

$$\log\frac{W_t}{W_{t-1}} = \rho\log\frac{W_{t-1}}{W_{t-2}} + \epsilon_t,\ \epsilon_t \sim N\left(0,\sigma_\epsilon^2\right)\ \text{with}\ |\rho| < 1. \tag{2}$$

The sectoral prices derived under equations (1) and (2), which we denote DGP1, are generated by:

$$p_{it} = \lambda_i p_{i,t-1} + \gamma_i w_t - \delta_i w_{t-1}, \;\; i = 1, ..., N; t = 1, ..., T, \tag{3}$$

where

$$p_{it} = \ln P_{it} \;\;,\;\; \gamma_i = \frac{1-\lambda_i}{1-\lambda_i \beta\rho} \;\;,\;\; w_t = \ln W_t \;\;,\;\; \delta_i = \frac{\lambda_i (1-\lambda_i)(1-\beta\rho)}{1-\lambda_i \beta\rho} \;\;,\;\; p_{i0} = 0 \;,$$

and where the degree of price rigidity is assumed to be uniformly distributed as $\lambda_i \sim U(a, b)$ with $0 \leq a < b \leq 1$.

Choi (2010) implicitly assumes that the degree of price rigidity is independent of the characteristics of the firm or sector. It is, however, well established that pricing behaviour depends upon a range of factors including the competitive structure of the industry or sector (Martin, 1993; Head, Kumar and Lapham, 2010). We introduce this form of heterogeneous regime-dependent rigidity in a simple manner as follows, yielding DGP2:

$$p_{it} = (\lambda_{1it} p_{i,t-1} + \gamma_{1it} w_t - \delta_{1it} w_{t-1}) 1_{\{p_{i,t-1} < w_{t-1}\}} + (\lambda_{2it} p_{i,t-1} + \gamma_{2it} w_t - \delta_{2it} w_{t-1}) 1_{\{p_{i,t-1} \geq w_{t-1}\}} \tag{4}$$

where

$$\gamma_{jit} = \frac{1-\lambda_{jit}}{1-\lambda_{jit}\beta\rho} \;\text{ and }\; \delta_{jit} = \frac{\lambda_{jit}(1-\lambda_{jit})(1-\beta\rho)}{1-\lambda_{jit}\beta\rho} \;\text{ for } j = 1, 2,$$

and where $1_{\{p_{i,t-1} < w_{t-1}\}}$ is a Heaviside function taking the value of unity if the price is lower than the marginal cost and zero otherwise. The regime-dependent rigidities are uniformly distributed such that $\lambda_{1it} \sim U(a, b)$ if $p_{i,t-1} < w_{t-1}$ while $\lambda_{2it} \sim U(c, d)$ if $p_{i,t-1} \geq w_{t-1}$, with $a < c$ and $b < d$. In this framework, price rigidities are generally stronger in the second regime when price exceeds marginal cost. This implies that firms operating with market power or in a weakly competitive environment will set their prices in a relatively rigid fashion. By contrast, in the first regime, price is less than marginal cost. Such a situation may arise due to strategic pricing behaviour (e.g. price competition), slow adjustment in the wake of a cost shock or due to government subsidies. We assume greater price flexibility in this regime, reflecting the imperative to maintain a positive income stream over anything but the (very) short-run. We also assume that $|\partial\lambda_{jit}/\partial d_{it}| < 0$ where $d_{it} = p_{i,t-1} - w_{t-1}$. This introduces an inertial mechanism akin to menu-costs, whereby price rigidities are lower in those sectors which experienced large differences between price and marginal cost in the previous period.[6]

Using these DGPs, we conduct a range of calibration-based experiments. In each case, we employ 1000 replications with $(N, T) = (100, 200)$ and we follow Choi's (2010) selection of common parameter values of $\beta = 0.99$ and $\sigma_\epsilon = 0.01$. We consider 17 different values of the AR coefficient in equation (2), i.e. $\rho = \{0.1, 0.15, ..., 0.85, 0.9\}$ and draw the price rigidities uniformly from the range of values recorded in Table 2.[7]

– Insert Table 2 here –

The lower and upper bounds for DGP1 are the same as those set by Choi (2010). Under DGP2, we employ the same range of values in a global sense but now we distinguish between lower and higher rigidities. Importantly, we allow the ranges of λ_{1i} and λ_{2i} to overlap. This provides greater generality than a strict separation between their values. Consider a simple case with two firms identified as A and B with heterogeneous pricing rigidities λ_{1A} and λ_{1B}. Let us assume that $\lambda_{1A} < \lambda_{1B}$. It would be excessively restrictive to assume that in the second regime $\lambda_{2A} > \lambda_{1B}$. Rather, if the degree of price rigidity in the second regime is to exceed that in the first regime, it is sufficient that $\lambda_{2A} > \lambda_{1A}$ and $\lambda_{2B} > \lambda_{1B}$.

For each DGP, we estimate the following nonlinear U-shaped (more accurately V-shaped) relationships between the higher moments (RPV and RPS) and the aggregate inflation rate:

$$RPV_t = c + \theta_1 \pi_t 1_{\{\pi_t \leq \bar{\pi}\}} + \theta_2 \pi_t 1_{\{\pi_t > \bar{\pi}\}} + e_{1t}, \tag{5}$$

$$SK_t = c + \varphi_1 \pi_t 1_{\{\pi_t \leq \bar{\pi}\}} + \varphi_2 \pi_t 1_{\{\pi_t > \bar{\pi}\}} + e_{2t}, \tag{6}$$

where $\pi_t = N^{-1}\sum_{j=1}^{N}\pi_{jt}$ and $\bar{\pi} = T^{-1}\sum_{t=1}^{T}\pi_t$. Without loss of generality, we set $\bar{\pi} = 0$ in all cases.

Figure 2 plots the average coefficients across replications, denoted $\bar{\theta}_1$, $\bar{\theta}_2$, $\bar{\varphi}_1$ and $\bar{\varphi}_2$. We observe a strong asymmetric mean-RPV relationship under both DGPs. Irrespective of the value of ρ, the relationship is positive in the higher inflation regime and negative in the lower regime, although the differential decreases somewhat as ρ increases. An inverse association between the degree of asymmetry and the value of ρ is intuitively plausible, as increasing the degree of persistence of firms' marginal costs reduces one source of uncertainty that they must consider when setting their prices.

– Insert Figure 2 here –

As noted by Choi (2010), a U-shaped mean-RPV relationship suggests that the association between inflation and inflation uncertainty does not depend upon the rate of inflation as such, but rather upon the deviation from a threshold rate of inflation. This threshold may represent a societal norm for inflation in which the economic system is relatively stable in the neighbourhood of the threshold and where deviations from it reflect shocks which are naturally associated with increased volatility. Interestingly, while his results suggest that the adoption of inflation targeting monetary policy reduces the minimum RPV rate of inflation, it is not always the case that this threshold coincides with the inflation target. Therefore, under certain circumstances, inflation targeters may have to trade lower inflation against greater uncertainty. More specifically, this may occur if the declared inflation target is set below the minimum RPV rate of inflation.

Moving on to the mean-RPS relationship, under DGP1 we observe a negative association that differs little between regimes, a finding that contrasts sharply with the majority of empirical studies that report a positive association (e.g. Ball and Mankiw, 1995; Balke and Wynne, 2000). It seems that Choi's DGP cannot reproduce the asymmetric mean-RPS relationship evident in Figure 1, indicating that it is only partially consistent with observed empirical regularities. Indeed, given the common observation of a positive linear RPV-RPS relationship and the robustness of Choi's U-shaped mean-RPV relationship, one would expect to observe a U-shaped mean-RPS relationship on the basis of transitivity. Its absence may, therefore, indicate that Choi's DGP is misspecified.

By contrast, under the extended framework of DGP 2, we observe a positive linear mean-RPS relationship for $\rho < 0.7$ with a U-shaped relationship emerging thereafter. Importantly, the degree of persistence of marginal costs required for asymmetry is consistent with the UK data over our sample period. Using a simple AR(1) model of aggregate wage inflation, we estimate the AR coefficient as 0.706 with a standard error of 0.129.[8]

– Insert Figure 3 here –

In conjunction with the stylised findings summarised above, our calibration exercises have a number of significant implications. Firstly, the positive mean-RPS association observed for low to moderate values of ρ suggests that the Ball-Mankiw view dominates where marginal costs are relatively flexible and becomes weaker as ρ increases. The switch to a negative association for $\rho > 0.85$ suggests that the downward nominal rigidities discussed by Tobin become more significant as marginal costs become increasingly persistent. This is an intuitively pleasing result because it follows that menu-cost considerations should be more pressing when marginal cost shocks are unpredictable (i.e. ρ is small) but that as marginal costs become more persistent, so firms become more constrained in their ability to make downward nominal adjustments and, therefore, become increasingly reliant on real adjustments.

It is, however, the transitional range $0.7 \leq \rho \leq 0.85$ that is of particular interest as this is where the U-shaped mean-RPS relationship emerges and it is also consistent with the UK data during our sample period. Within this range, our results support our stylized finding that Tobin's negative mean-RPS relationship dominates in the low inflation regime while the positive relationship associated with the menu-cost model dominates in the high inflation regime. This implies that when marginal costs are sufficiently persistent that neither menu-costs nor downward nominal rigidities dominate price-setting behaviour, then the inflationary regime plays a key role in price-setting. In a high inflation setting, firms are less constrained in their ability to make rapid relative adjustments to overcome nominal rigidities, but the menu-costs associated with price changes become increasingly important. The linkage to the widely discussed 'grease' and 'sand' effects of inflation is readily apparent (Aucremanne et al., 2003).

At an over-arching level, our results indicate that both the Ball-Mankiw and Tobin explanations may be applicable under different feasible parameter sets. This is an important result, because each model has different implications for the conduct of monetary policy. In the former, low and stable inflation is desirable as it reduces the losses associated with regular costly price adjustments. By contrast, in the latter, the central bank may prefer somewhat more rapid inflation as it would prevent the build-up of potentially harmful pricing distortions due to downward nominal rigidity.

Taken together, these results suggest that the optimal level of inflation will be near the turning point of the mean-RPS relationship, which also coincides with the minimum RPV rate of inflation due to the linearity of the RPV-RPS relationship. Under the common assumption of a positively sloped linear mean-RPV relationship, the pursuit of low inflation and low inflation uncertainty go hand-in-hand. However, as noted by Choi (2010), where the mean-RPV relationship is U-shaped, a central bank may reduce economic welfare by pursuing disinflationary policies if they are associated with rising uncertainty. Similarly, given that RPS measures the degree of symmetry of price adjustments, it follows that large absolute values of RPS indicate that price adjustment is concentrated in certain sectors which may be indicative of undesirable economic (allocative) distortions.

At present, inflation targets tend to be set in a rather ad hoc manner, taking into account rough notions of mismeasurement of the price index, quality enhancements and productivity gains. However, our results provide a more formal method of determining an appropriate inflation target. If the inflation target is to act as a nominal anchor, it follows that this can be most readily achieved if inflation uncertainty is minimised and the distribution of relative price changes is approximately symmetric. The latter is likely to be especially important where the patterns of expenditure of different economic and social groups show considerable heterogeneity that may expose some groups more than others to inflation in specific sectors, such as liquid fuels and staple foods. Therefore, provided that it would not move the economy too close to the zero lower bound on nominal interest rates, the inflation target should be set within the neighbourhood of the turning points of the mean-RPV and mean-RPS relationships.[9]

The usefulness of our analytical framework is not limited to the case of countries that wish to adopt inflation targeting: it also provides a means of assessing the appropriateness of existing inflation targets. In the case of the UK, it is interesting to note that Choi (2010) reports a 95% confidence interval for the minimum-RPV rate of inflation of 0.8-1.9%. This range is generally consistent with Figure 1, above, which also suggests a turning point lower than 2% but still within the Bank's 1-3% permissible range. This suggests that if the Bank of England were to target a lower rate of inflation than it currently does, inflation uncertainty may fall.[10] By contrast, if a central bank targets a rate of inflation below the turning point of the U-shaped mean-RPV and mean-RPS relationships, then it is likely that increasing the target may lead to welfare gains due to reduced levels of volatility and pricing distortions. Finally, we contend that in a system with no declared inflation target, the common turning point may provide a useful proxy.

4 EVALUATION OF THE REGIME-SWITCHING CALVO MODEL

We have demonstrated that our regime-switching Calvo model can replicate an important stylized finding from the UK sectoral data, namely the asymmetric U-shaped patterns that characterise the relationships between the mean and the higher moments of the distribution of sectoral price changes. We now turn to the question of whether our calibrated model can successfully replicate the distribution of actual price changes in the UK data. Furthermore, we will also briefly discuss the implications of our findings for the debate over the relative merits of headline and core measures of inflation.[11]

4.1 Replicating the Distribution of Sectoral Price Changes

To establish a benchmark, we first evaluate the degree to which Choi's DGP can approximate the distribution of sectoral price changes. To this end, we estimate the parameters $\hat{\lambda}_i$, $\hat{\gamma}_i$ and $\hat{\delta}_i$ from equation (3) for each sector.[12] To replicate the properties of the dataset as closely as possible, we then generate prices for 85 sectors over 134 months. The distribution of standardised price changes at each time period is then estimated using the kernel density estimator $\left(\hat{f}_t\right)$ with 1,024 grids. Finally, we take the average of these distributions through time $\left(\text{i.e. } \bar{f} = (1/134)\sum_{t=1}^{134} \hat{f}_t\right)$ and compare $\bar{f}$ with the average distribution of actual price changes $\left(\bar{f}_0\right)$, which is also estimated using the kernel density estimator.

In the case of DGP2, we employ the same procedure but consider the following time-invariant specification of equation (4) because of difficulties estimating the time-varying parameters given the available data:

$$p_{it} = \left(\lambda_{1i} p_{i,t-1} + \gamma_{1i} w_t - \delta_{1i} w_{t-1}\right) 1_{\{p_{i,t-1} < w_{t-1}\}} + \left(\lambda_{2i} p_{i,t-1} + \gamma_{2i} w_t - \delta_{2i} w_{t-1}\right) 1_{\{p_{i,t-1} \geq w_{t-1}\}}, \tag{7}$$

As before, we estimate $\left(\hat{\lambda}_{1i}, \hat{\gamma}_{1i}, \hat{\delta}_{1i}\right)$ and $\left(\hat{\lambda}_{2i}, \hat{\gamma}_{2i}, \hat{\delta}_{2i}\right)$ for each sector and filter $\hat{\lambda}_{1i}$ and $\hat{\lambda}_{2i}$ for outliers before comparing the calibrated and actual distributions.

Figure 4 compares the calibrated distributions based on DGPs 1 and 2 against the distribution of actual price changes. In both cases, the calibrated and actual distributions are very close to one-another, implying that both models can replicate the properties of the real data rather well given data-based parameters. The main difference between the two DGPs is that DGP1 exhibits a mild right skew and is slightly platykurtic, while DGP 2 is more centred but somewhat leptokurtic.

– Insert Figure 4 here –

It is not obvious from visual inspection which DGP more closely matches the distribution of actual price changes. Indeed, the absolute distance of the calibrated distributions from the actual distribution appears to be similar in both cases. The nonparametric estimates of the first four moments for each of the three distributions reported in Table 3 provide a basis for more detailed comparison. These reveal that DGP 1 is closer to the actual distribution in terms of both mean and variance, while DGP2 is closer in terms of skewness and kurtosis.

– Insert Table 3 here –

We also compute the following three formal divergence criteria: (i) the uniform norm (D_U), (ii) the Hilbert norm (D_H), and (iii) the generalised entropy measure (D_E). These are defined respectively as

$$D_U\left(\bar{f},\bar{f}_0\right)=\frac{\int\left\{\bar{f}\left(x\right)-\bar{f}_0\left(x\right)\right\}^2dx}{\int\bar{f}\left(x\right)^2dx+\int\bar{f}_0\left(x\right)^2dx}\quad,\quad D_H\left(\bar{f},\bar{f}_0\right)=\frac{\sup_x\left|\bar{f}\left(x\right)-\bar{f}_0\left(x\right)\right|}{\sup_x\bar{f}_0\left(x\right)}$$
$$\text{and}\quad D_E\left(\bar{f},\bar{f}_0\right)=\int\bar{f}_0\left(x\right)g\left(\frac{\bar{f}\left(x\right)}{\bar{f}_0\left(x\right)}\right)dx$$

where $\bar{f}$ is the calibrated distribution, $\bar{f}_0$ the actual distribution, and where $g\left(y\right)=\left(\gamma-1\right)^{-1}\left(y^\gamma-1\right)$ with $\gamma>0$ and $\gamma\neq1$.[13] All three quantities are non-negative and return a value of zero if $\bar{f}=\bar{f}_0$.

Table 4 reports the divergence measures for each DGP. We find that DGP 1 presents a smaller value than DGP 2 for the uniform norm, while DGP 2 presents a smaller value for the entropy measure. In the case of the Hilbert norm, the difference between the DGPs is negligible.

– Insert Table 4 here –

We must conclude, therefore, that both DGPs describe the actual distribution to a similar degree of accuracy, but that DGP2 gives a better characterisation of the higher moments. The observation that DGP2 better captures the higher moments of the distribution of sectoral price changes is intuitively plausible given our earlier finding that DGP1 is incapable of replicating the U-shaped mean-RPS association observed in the raw data.

4.2 Core Inflation

There is considerable debate among economists about the relative merits of headline and core measures of inflation. A widely held position is that by excluding the highly volatile food and energy sectors, core inflation may achieve a higher signal to noise ratio (e.g. Bryan and Ceccheti, 1994). An obvious question, therefore, is whether similar U-shaped mean-RPV and mean-RPS relationships characterise both headline and core inflation data, or whether the asymmetries are driven by the more volatile sectors. Figure 5 reproduces the graphs from Figure 1 in the case of core CPI.[14]

– Insert Figure 5 here –

It is immediately apparent that the U-shaped mean-RPV and mean-RPS relationships are greatly attenuated when working with core CPI. Indeed, while the negative association in the low inflation regime remains, there appears to be no significant relationship between the mean and either RPV or RPS when inflation is above the historical average. This is an interesting finding which suggests that the Ball-Mankiw menu-cost approach may largely break down once the highly volatile food and energy sectors are removed. In essence, the menu-cost model assumes that desired price changes are only affected if they exceed a threshold value. It follows, therefore, that removing the most volatile sectors from the dataset will reduce the proportion of cases in which the desired price change exceeds the menu-cost threshold. When viewed in this way, Figure 5 suggests that the desired price changes in the non-food and non-energy sectors in the UK over our sample period may generally have been too small to be enacted in the presence of non-negligible menu costs. It follows similarly that removing the most volatile sectors would not completely undermine the Tobin-style mechanism; rather it would weaken it and reduce the speed with which firms can achieve downward relative price adjustments. This is consistent with the patterns reported in Figure 5(b).

This line of reasoning cautions against the use of core inflation for policy purposes, especially in economies where menu-costs are thought to be important. While core CPI *may*, in some cases, provide clearer signals about inflationary trends than headline CPI, it does so at the expense of valuable information about the higher moments of the distribution of sectoral price changes. The Ball-Mankiw and Tobin models emphatically stress the fundamental role played by these higher moments in facilitating price adjustments. Moreover, the higher moments of inflation may exert a strong influence on inflation expectations. It follows, therefore, that policymakers should attach greater weight to headline inflation data than to simplified, truncated measures. Given that inflation-targeting monetary policies are set with recourse to inflation forecasts, it follows that

CKS's work on forecasting the time-varying distribution of sectoral inflation rates may provide policymakers with an invaluably enriched information set.

5 CONCLUSION

This paper extends the modified Calvo (1983) model developed by Choi (2010) by allowing for regime-dependent price rigidities. Using this framework, we examine asymmetries in the relationships between the mean of sectoral inflation rates and the higher order moments of their distribution. Our calibration exercises reveal an asymmetric U-shaped relationship between both mean and RPV and between mean and RPS under plausible parameterisations. More specifically, where sectoral price rigidities are relatively strong and marginal costs relatively persistent, an economy in which firms are not constrained to engage in marginal cost pricing at all times may deliver U-shaped asymmetries in both the mean-RPV and mean-RPS relationships.

Our findings have important implications for monetary policy, as an inflation targeting central bank may face a trade-off between low inflation on the one hand and increased volatility and absolute skewness of relative price changes on the other. We conclude by suggesting three avenues for continuing research. Firstly, as with Choi (2010), we limit our attention to the case of common marginal costs across sectors. The extension to heterogeneous costs may better reflect sectoral skills differentials, for example, and may yield deeper insights into pricing behaviour. Secondly, we develop DGP2 on the basis of the obvious case of imperfect competition. However, our specification represents a very simplistic case that could be readily refined and extended. Moreover, price rigidities may depend on a wide range of other factors such as the mean level of inflation (Devereux and Yetman, 2002). Finally, in light of the pervasive international evidence of U-shaped relationships between aggregate inflation and volatility adduced herein and by Becker and Nautz (2010), Chen, Shen and Xie (2008), Fielding and Mizen (2008) and Choi, further research into optimal policy design and coordination in a cross-country setting should be viewed as a priority.

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NOTES

1. We note in passing that while Friedman argues that higher inflation leads to higher uncertainty, Cukierman and Meltzer stress that it is uncertainty which is inflationary.

2. A notable exception is Balke and Wynne (2000), which analyses both the mean-RPV and mean-RPS relationships, finding positive correlations in both cases. More recently, Aucremanne et al. (2003) have carefully analysed sectoral data in Belgium, and have also found the Ball-Mankiw model preferable to the Tobin model. Our work extends beyond the scope of these studies due to our handling of nonlinearity in the mean-RPS relationship.

3. CKS employ a semi-parametric functional autoregressive approach to model and forecast the time-varying distribution of UK sectoral inflation rates. Their model performs admirably in a range of forecast evaluation tests.

4. The policy framework is homogeneous in the sense that the central bank has manipulated the short-term interest rate to achieve a declared inflation target throughout this period. Some subtle changes have been made to the target itself during this time, but there has been no fundamental change in the policy stance of the Bank of England (in particular, note that our sample period ends prior to the enactment of quantitative easing at the Bank of England). While it would be interesting to consider earlier periods in order to evaluate the effect of inflation targeting on the mean-RPV and mean-RPS relationships, no comparable data are available for the UK prior to 1997.

5. Due to space constraints we do not report detailed descriptive statistics – they are available on request.

6. Note that according to Klenow and Kristov (2008) the simple time-dependent Calvo model cannot account for the distribution of price changes in the firm level data. The Calvo model treats the magnitude of price changes as endogenous but the timing/frequency of individual price changes as exogenous. In the state-dependent sticky price models (SDP), firms decide to change the price subject to menu costs. The SDP models by Midrigan (2011), Gertler and Leahy (2006) and Dotsey, King and Wolman (2006) show better performance in terms of matching the price changes with inflation. Our modified Calvo model can be viewed as an amalgamation of the time-dependent and the state-dependent models as we specifically allow for regime dependent price rigidities.

7. To ensure that our results are not dependent on the assumption that price rigidities are uniformly distributed, we repeated our analysis while drawing the price rigidities from a variety of different distributions, including a truncated normal distribution. The calibration results show little sensitivity to such changes. Details are available on request.

8. Using a threshold specification, we estimate AR coefficients of 0.715 and 0.699 respectively in the low and high inflation regimes, although we cannot reject the null hypothesis of symmetry. Note, however, that Choi (2010) sets $\rho = 0.9$.

9. As noted above, these turning points coincide in our analysis. This is likely to be a fair approximation

of reality given the linearity of the RPV-RPS relationship. Further support for this view is provided by Becker and Nautz (2010), who show that the minimum point of the mean-RPV relationship in their search model is modestly higher but very near to the welfare maximising rate of inflation.

10. An alternative interpretation is that the Bank of England has been operating according to a de facto target that is lower than its declared target, and that the observed turning point reflects this de facto target (or perhaps public perceptions of the latter). In cases like this, where the turning point does not coincide with the declared target, policymakers may be acting with discretion. For example, the Bank of England may aim for a level of inflation slightly below the centre point of its 1-3% target band if policymakers are inflation-averse. If this is the case, however, one must ask whether the declared target should be revised downward.

11. A good example of the differing views regarding headline and core inflation can be found by comparing Smith (2004) and Crone et al. (2011). While the former argues in favour of using core inflation to forecast the rate of inflation in the US, the latter holds that core inflation is not necessarily the best predictor of total inflation and that the forecasting performance depends on the inflation measure and forecasting horizon. The debate continues.

12. w_t is only available at quarterly frequency, so we generate monthly values by interpolation under the assumption that $\rho = 0.7$. In a small minority of cases we observe $\hat{\lambda}_i > 1$, which is inconsistent with the theoretical condition $0 \leq \lambda \leq 1$. Hence, we apply a simple filtering rule by replacing any values larger than 0.99 with 0.99. Note that we have obtained qualitatively similar results for different values of ρ between 0.7 and 0.85.

13. For our purposes, we set $\gamma = 1/2$.

14. We construct core CPI by excluding both the food sectors (1-9) and the energy sectors (25-28), and then adjusting the weighting scheme accordingly.

Table 1: Threshold Regression Estimates

Model	β^+	β^-	$H_0 : \beta^+ = \beta^-$
$\mu_t = \left(\alpha^+ + \beta^+\sigma_t^+\right) + \left(\alpha^- + \beta^-\sigma_t^-\right) + \epsilon_t$	0.32 (0.00)	-0.21 (0.00)	[0.00]
$\mu_t = \left(\alpha^+ + \beta^+ s_t^+\right) + \left(\alpha^- + \beta^- s_t^-\right) + \epsilon_t$	0.29 (0.00)	-0.20 (0.00)	[0.00]
$\sigma_t = \left(\alpha^+ + \beta^+ s_t^+\right) + \left(\alpha^- + \beta^- s_t^-\right) + \epsilon_t$	0.47 (0.00)	0.19 (0.04)	[0.00]

Note: μ_t, σ_t and s_t denote the mean, standard deviation and skewness of the cross-sectional distribution of sectoral inflation rates. To examine the asymmetric relationships between the moments, we decompose them according to the high and low inflation regimes as follows: $\sigma_t^+ = \{\sigma_t | \mu_t \geq \mathbf{E}[\mu_t]\}$, $\sigma_t^- = \{\sigma_t | \mu_t < \mathbf{E}[\mu_t]\}$, $s_t^+ = \{s_t | \mu_t \geq \mathbf{E}[\mu_t]\}$, $s_t^- = \{s_t | \mu_t < \mathbf{E}[\mu_t]\}$. Coefficient standard errors are reported in rounded parentheses while the p-values of the F-test for the null of symmetry are shown in square parentheses.

Table 2: Distribution of Price Rigidities

	DGP1	DGP2	
	λ_i	λ_{1i}	λ_{2i}
Lower bound	0.60	0.60	0.70
Upper bound	0.99	0.90	0.99

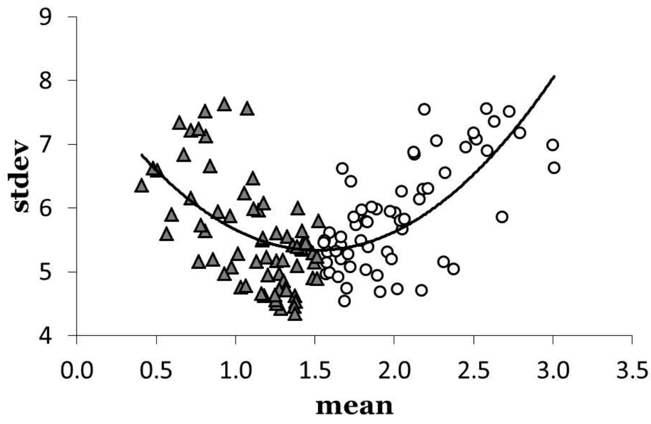
Table 3: Comparison of Central Moments

Moments	Actual	DGP 1	DGP 2
Mean	-0.007	-0.006	0.015
Variance	1.172	1.212	0.894
Skewness	-0.521	-0.911	-0.844
Kurtosis	6.017	3.845	6.230

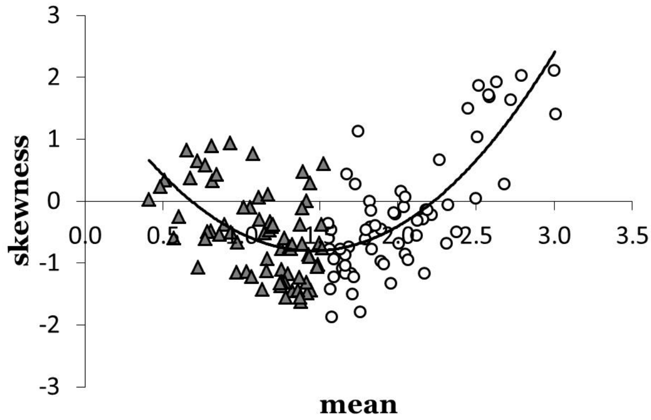
Table 4: Comparison of Divergence Criteria

	DGP1	DGP2
Uniform Norm	0.0769	0.0935
Hilbert Norm	0.0024	0.0025
Generalised Entropy Measure	0.0051	0.0041

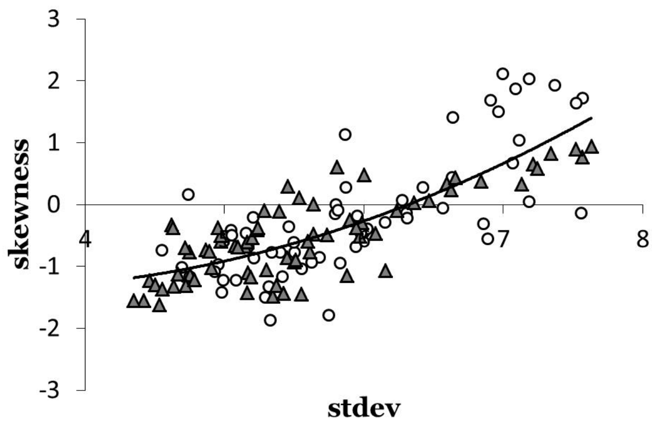
Figure 1: The Relationship between Cross-Sectional Moments



(a) Mean:RPV



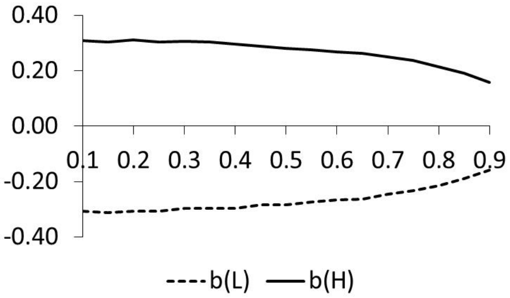
(b) Mean:RPS



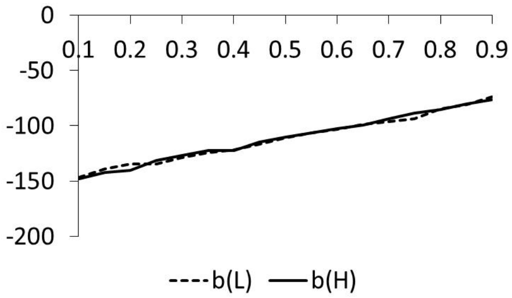
(c) RPV:RPS

Note: Triangular (circular) markers indicate that inflation is below (equal to or above) average.

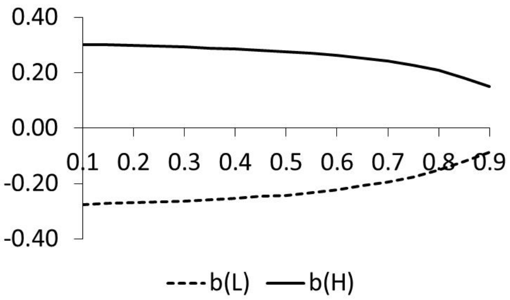
Figure 2: The Relationship between Aggregate Inflation and Higher Order Moments



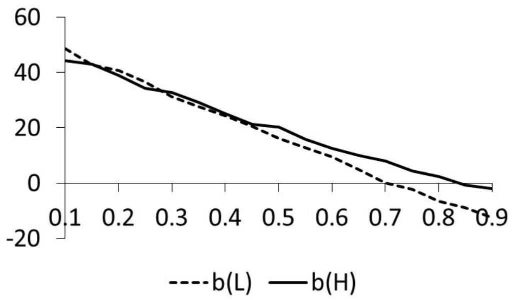
(a) DGP 1: RPV



(b) DGP 1: RPS



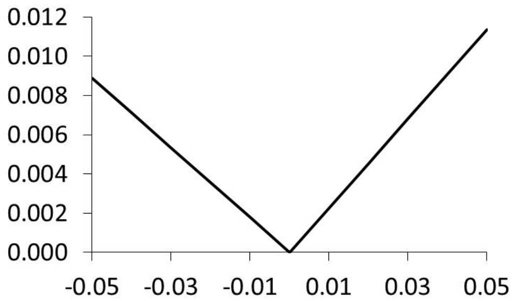
(c) DGP 2: RPV



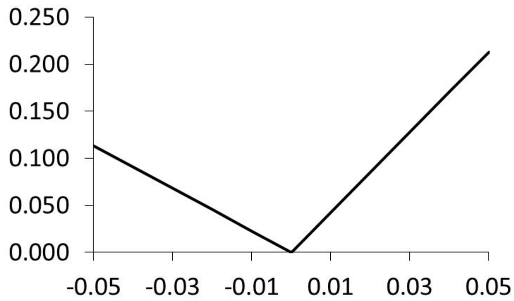
(d) DGP 2: RPS

Note: b(L) and b(H) denote the low inflation and high inflation regime-specific coefficients, respectively. The horizontal axis records the value of the AR(1) coefficient, ρ.

Figure 3: U-Shaped Relationships between Aggregate Inflation and Higher Order Moments



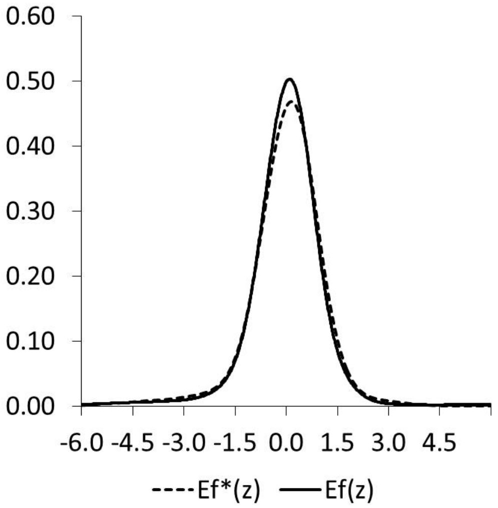
(a) DGP 2 ($\rho = 0.75$): RPV



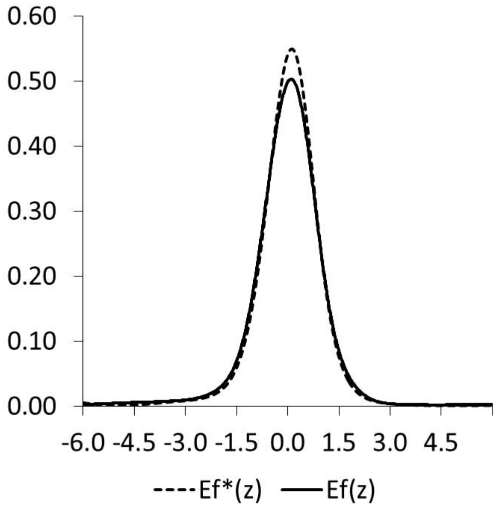
(b) DGP 2 ($\rho = 0.75$): RPS

Note: The horizontal axis records the aggregate rate of inflation relative to the mean.
The vertical axis shows the estimated response based on equations (5) and (6).

Figure 4: The Distribution of Standardised Price Changes



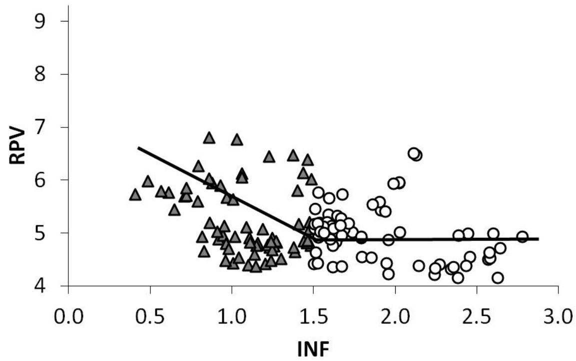
(a) DGP 1



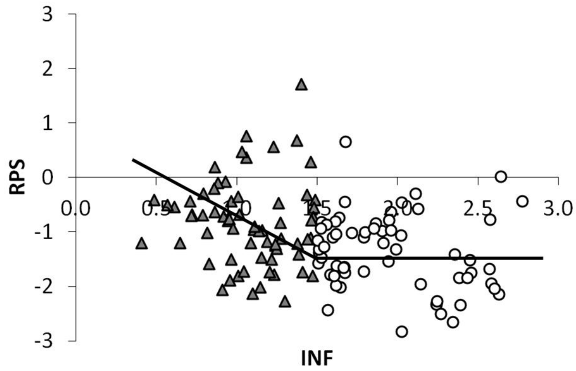
(b) DGP 2

Note: The distribution of standardised price changes is estimated by the kernel density estimator. The solid line represents the average of the empirical distribution from the actual data, while the dashed line represents the average of the calibrated distribution from the model.

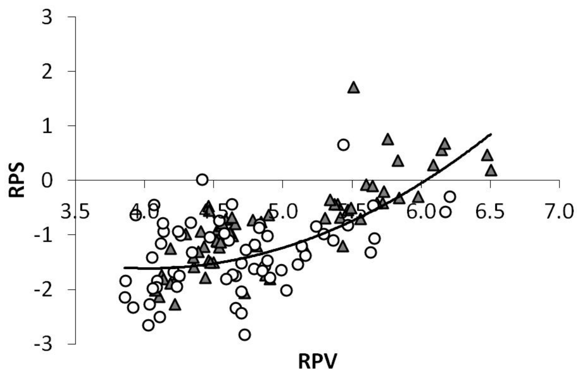
Figure 5: The Relationship between Cross-Sectional Moments using Core Inflation Rates



(a) Mean:RPV



(b) Mean:RPS



(c) RPV:RPS

Note: Triangular (circular) markers indicate that inflation is below (equal to or above) average.