The Determinants of Relative Price Variability

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This paper investigates how the variability of relative price changes (RPV) is associated with various aspects of inflation. Expected inflation, unexpected inflation, and, to a lesser extent, inflation uncertainty are found to be positively related to RPV. These findings are compared with the predictions of three well-known models that predict a relationship between some aspect of inflation and RPV: the menu-costs model, the Lucas-Barro signal extraction model, and the Hercowitz-Cukierman extension of the Lucas-Barro model. The findings are consistent with the first model and, to a lesser extent, the second, but are inconsistent with the third.

Introduction

The variability of relative price changes increases with inflation. Studies dating to Mills (1927) and Graham (1930) have consistently found evidence of this variability.[1] The literature, however, is divided on the question of which particular aspect of inflation is most closely related to the variability of relative price changes. The modern literature beginning with Parks (1978) has arrived at various conclusions on this matter, despite consistently defining the variability of relative price changes in period t (more commonly called "relative price variability" and denoted as RPV) as the sum of the squared deviations of rates of change of various price subindexes from the rate of change of an overall price index in period t.[2] Parks finds that RPV increases primarily with the absolute value of unexpected inflation. Tang and Wang (1993) find that RPV increases with expected inflation as well as with the absolute value of unexpected

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1 For the most part, researchers have examined this question using aggregate, economy-wide, data. The recent studies of Parsley (1996) and Debelle and Lamont (1997) also have found a positive relationship between inflation and variability in relative price changes using city-level U.S. data.

2 Given the definition, the phrase "variability of relative prices changes" is a more apt description than "relative price variability." The "relative price variability" phrase and the acronym RPV are so entrenched that I use them below.

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inflation. Fischer (1981a, 1982) finds that RPV increases with expected inflation and positive unexpected inflation, but not with negative unexpected inflation. Still others, such as Grier and Perry (1996) find that RPV increases only with *ex ante* inflation uncertainty.

Expected inflation, realized (i.e., *ex post*) unexpected inflation, and *ex ante* inflation uncertainty all have been proposed as determinants of RPV in well-specified theoretical models, as outlined below. No empirical analysis of the relationship between inflation and RPV has simultaneously included all of these aspects of inflation as explanatory variables. The empirical studies prior to Grier and Perry use only measures of expected and unexpected inflation as explanatory variables. At the same time, while Grier and Perry extend this empirical literature by including a plausible measure of *ex ante* inflation uncertainty as an explanatory variable, they omit unexpected inflation.

This study investigates the relationship between inflation and RPV using an empirical specification that includes a measure of inflation uncertainty as well as measures of both expected and unexpected inflation. Expected inflation, unexpected inflation, and, to a lesser extent, inflation uncertainty all are found to be positively and significantly related to RPV. These findings shed light on three well-known models that predict a relationship between some aspect of inflation and RPV: the menu-costs model, the Lucas-Barro signal extraction model, and the Hercowitz-Cukierman extension of the Lucas-Barro model. The findings are consistent with the menu-costs model and, to a lesser extent, the Lucas-Barro model, but are inconsistent with the Hercowitz-Cukierman model.

An additional contribution of this paper is the use of a measure of RPV constructed from price data that are more disaggregated than the data used in most previous studies. While most previous studies measure RPV using the subcomponents of the personal consumption expenditure (PCE) deflator or the subcomponents of the producer price index (PPI) at the two-digit level of disaggregation, this study measures RPV using the more numerous subcomponents of PPI at the three-digit level.

The organization of the paper is as follows. First I review the three well-known theories that have been proposed as possible explanations of the relationship between inflation and RPV. After discussing the data, I describe the statistical model of inflation that generates measures of expected inflation, unexpected inflation, and inflation uncertainty. Then I report and interpret the estimation results from regressions of RPV upon measures of expected inflation, unexpected inflation, and inflation uncertainty.

Theories Linking Inflation and RPV

Three well-developed theories predict relationships between RPV and inflation:

(1) The signal-extraction model of Lucas (1972, 1973) and Barro (1976);

(2) The extension of the signal-extraction model by Hercowitz (1981, 1982) and Cukierman (1983); and

(3) The menu-costs model of Sheshinski and Weiss (1977), Rotemberg (1983), and Bénabou (1988, 1992).

Each of these makes a distinct prediction regarding which aspect of inflation will be most closely related to RPV.[3]

The Lucas-Barro signal-extraction model predicts that RPV will increase with *ex ante* inflation uncertainty. The greater is the *ex ante* variability of aggregate demand shocks (and *ex ante* inflation uncertainty), the more various real local shocks will be misinterpreted as aggregate shocks and responded to with price changes rather than quantity changes. Realized aggregate demand shocks, on the other hand, have no effect on RPV in the Lucas-Barro model because all firms respond identically to a given aggregate demand shock.

By contrast, realized aggregate demand shocks do affect RPV in the Hercowitz-Cukierman extension of the Lucas-Barro signal-extraction model in which price elasticities of supply differ across firms. In the Hercowitz-Cukierman model, firms with high elasticities of supply adjust prices less in response to a given unexpected aggregate demand shock than do firms with low elasticities of supply. Because the discrepancies in price adjustments across firms will increase with the magnitude of the realized aggregate demand shock, the Lucas-Barro model predicts that RPV will increase with the magnitude of unexpected inflation, irrespective of whether it is positive or negative.

Finally, anticipated changes in aggregate demand have no effect upon relative prices under either version of the signal-extraction model. Anticipated changes in aggregate demand simply result in a uniform increase or decrease in all nominal prices. By contrast, the menu-costs models of Sheshinski and Weiss, Rotemberg, and Bénabou predict a positive relationship between RPV and expected inflation. These models predict that firms will use (S, s) pricing schemes if a fixed cost is incurred whenever output price is adjusted: a given firm will adjust the nominal price of its good only when the real price of its good decays to some lower bound s, at which time the price will be reset so that its real price equals some upper bound S. As inflation increases, these menu-costs models predict that the difference between the optimal s and S will increase. Provided that firms do not adjust prices synchronously, this implies that RPV will increase with the rate of inflation.[4] Symmetrically, in circumstances of deflation, RPV will increase with the rate of deflation. The predictions of menu-costs model as well as those of the Lucas-Barro and Hercowitz-Cukierman models regarding the relationship between inflation and RPV are summarized in Table 1.

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[3] Fischer (1981b) surveys the various theories linking inflation and RPV.

[4] Danziger (1987) notes that the period between observations on prices must be short compared to the period over which firms maintain a fixed nominal price for this prediction to hold. If these two periods were equal, for instance, then RPV would be zero despite (S, s) pricing policies and would be completely unrelated to expected inflation.

Table 1-Predictions of the Menu-Costs, Lucas-Barro, and Hercowitz-Cukierman Models Regarding the Relationship between RPV and Various Aspects of Inflation

	Menu-Costs Model	Lucas-Barro Model	Hercowitz-Cukierman Model
Expected Inflation	positive	0	0
Expected Deflation	positive	0	0
Inflation Uncertainty	0	positive	0
Unexpected Inflation (When Positive)	0	0	positive
Unexpected Inflation (When Negative)	0	0	positive

The Hercowitz-Cukierman model predicts that the coefficients on positive unexpected inflation and negative unexpected inflation will, in addition to being positive, be equal in size

Data

The data used are monthly producer price index (PPI) data from 1947.01–1997.10 from the Bureau of Labor Statistics (BLS) website.[5] The starting date of 1947.01 is chosen because data for most PPI subindexes begin in that month. Inflation in month t is measured as the difference between the natural logs of the PPI all-commodities price index in months t and t - 1. The rate of price change for each individual subindex within the PPI all-commodities price index is defined similarly. The variability of relative price changes in period t (RPV_t) is measured as:

$$(1)\quad RPV_t = (1/n)\sum_{i=1}^{n}(\pi_{it} - \pi_t)^2$$

where:

π_t = The overall inflation rate; and

π_{it} = The rate of change of the ith subindex.

Both π_t and π_{it} are expressed in decimal form in equation (1), i.e., 2.5 percent inflation appears as 0.025.

The particular measure of RPV used here is computed using all 64 subindexes within the PPI at the three-digit level for which continuous monthly data are available from 1947.01 to 1997.10. These 64 subindexes represent over 60 percent of the 105 subindexes within the PPI all-commodities price index which are defined by the BLS at the three-digit level. Weights are not available for the various subindexes, so this is an unweighted measure of RPV.[6]

Figures 1 and 2 depict inflation and RPV, respectively, over the sample period. The mean and standard deviation of monthly PPI inflation are 0.0027 (= 0.27 percent) and 0.0070 (= 0.70 percent), respectively, during the sample period. The mean and standard deviation of RPV are 0.00062 and 0.00083, respectively. The correlation between inflation and RPV is 0.30.

5 The website of the BLS is located at <http://stats.bls.gov>. The data used in this study were downloaded from the webpage http://146.142.4.24/cgi-bin/srgate. This website allows the retrieval of specific PPI series by inputting the appropriate series codes. The codes for the various series used in this study are listed in Appendix A.

6 These data are available from the author.

Figure 1—Monthly Inflation, 1947.01—1997.10

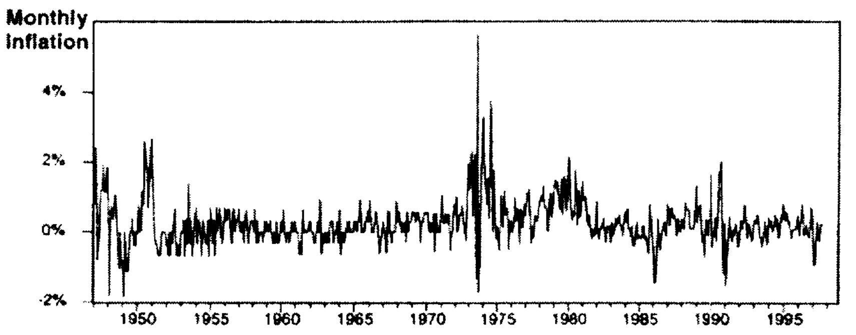
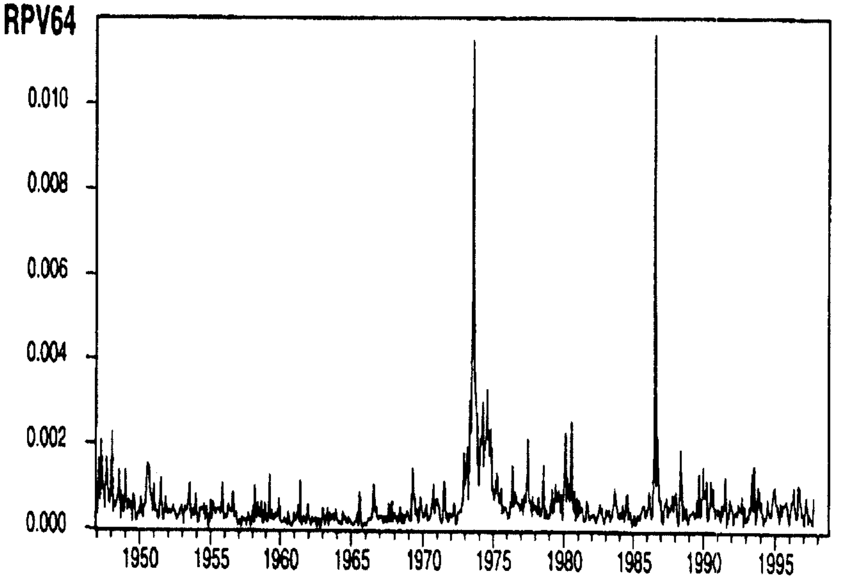


Figure 2—Relative Price Variability, 1947.01—1997.10



A Model of Inflation and Inflation Uncertainty

Estimation of preliminary ARMA models of inflation indicates that the inflation-prediction errors are heteroskedastic. In particular, the ARCH-LM statistics for testing the serial independence of the squared residuals at lags 1 through 8 are all over 100 for the best-fitting ARMA model and indicate the rejection of homoskedasticity at the 1 percent significance level. Given the finding of heteroskedasticity, I model inflation as a generalized autoregressive conditional heteroskedasticity (GARCH) process to allow the variance of the error term to evolve through time. Specifically, I model inflation using a GARCH (1,1) process in which the conditional variance of inflation in period t is a function of the conditional variance of inflation in period t-1 and the inflation prediction error in period t - 1.[7]

Let υ_t be the inflation prediction error in period t, let $\sigma^2_{\upsilon,t}$ be the variance of the inflation prediction error in period t, and let the subscript t - i indicate the ith monthly lag of the subscripted variable. Using this notation, the following GARCH (1,1) model of inflation over the 1947.01–1997.10 period minimizes the Schwarz criterion among GARCH (1,1) models with well-behaved residuals:

$$(2)\quad \underset{(4.63)}{\pi_t = 0.00199} + \underset{(7.56)}{0.586\pi_{t-1}} + \underset{(2.24)}{0.0875\pi_{t-6}} - \underset{(-2.86)}{0.287\upsilon_{t-1}} + \underset{(3.84)}{0.129\upsilon_{t-12}} + \upsilon_t$$

where:

$$(3)\quad \underset{(3.66)}{\sigma^2_{\upsilon,t} = 0.00000152} + \underset{(6.70)}{0.194\upsilon^2_{t-1}} + \underset{(24.11)}{0.759\sigma^2_{v,t-1}}.$$

(Adj. R-squared = 0.19)

(F-statistic for overall significance = 21.54)

The t-statistics of the coefficient estimates are reported in parentheses. The overall regression is significant at the 1 percent level, and all the estimated coefficients are statistically significant at the 5 percent level. This model of inflation is similar to that employed in Grier and Perry (1996).

The process reported in equation (2) generates expected and unexpected inflation. As is standard in this literature, I take the one-period-ahead forecasts from equation (2) to be expected inflation (EI) and take the difference between actual inflation and expected inflation to be unexpected inflation (UI). Similarly, equation (3) generates a series for the conditional variance (CVI) of inflation. For simplicity, CVI is referred to as *inflation uncertainty*. Finally, in order to test the prediction of the Cukierman-Hercowitz theory that the sign of unexpected inflation is irrelevant, I construct two auxiliary series from UI. The series for positive unexpected inflation (UIP) takes the value 0 in months in which UI is non-positive and is set equal to UI in months in which UI is positive. Similarly, the series for negative unexpected inflation (UIN) takes the value 0 in months in which UI is non-negative and is set equal to UI in months in which UI is negative.[8]

The variables EI, UI, UIN, and UIP are squared when used as explanatory variables. This is done in order to test the predictions of the menu-costs and Hercowitz-Cukierman models. The former model predicts that RPV will increase with the absolute

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[7] I chose a GARCH (1,1) model for inflation rather than some higher-order GARCH model because GARCH (1,1) models are predominant in the macroeconomics time-series literature on inflation.

[8] All the participants in the economy are assumed to use equations (2) and (3), based on experience prior to the sample period, to judge EI, UI, UIP, UIN, and CVI. Under this assumption, the use of these five series as explanatory variables in subsequent regressions does not create an errors-in-variables problem.

Table 2-Unit Root Tests for Inflation, RPV, $(EI)^2$, $(UI)^2$, and CVI

	Inflation (π)	RPV	$(EI)^2$	$(UI)^2$	CVI
Dickey-Fuller	-16.96	-15.03	-15.53	-15.02	-5.20
Philips-Perron	-17.80	-15.67	-16.17	-14.98	-5.58

The 1 percent critical value for rejection of the hypothesis of a unit root is -3.44 for both tests of all variables. The Phillips-Perron tests allow for an intercept term and use a truncation lag of five

magnitude of expected inflation, while the latter model predicts that RPV will increase with the absolute magnitude of unexpected inflation.

All the variables used in the regressions below appear to be stationary. Dickey-Fuller and Phillips-Perron unit root tests for each of the series π, RPV, CVI, $(EI)^2$, and $(UI)^2$ lead to the rejection of the hypothesis of a unit root at the 1 percent level. Table 2 reports the results of these tests.

Relationships between Inflation and RPV

Estimation Results

As a preliminary exercise, I verify the positive relationship between inflation and RPV found by Parks (1978) and others. A regression of RPV on inflation yields

$$\text{(4)}\quad \underset{}{RPV_t} = \underset{(12.38)}{0.000522} + \underset{(2.60)}{0.0361\pi_t} + \upsilon_t$$

(Adj. R-squared = 0.09)

The t-statistics in parentheses are based upon standard errors computed according to the procedure proposed in Newey and West (1987) to allow for residuals that exhibit both autocorrelation and heteroskedasticity of unknown form.[9] The traditional positive association between inflation and RPV clearly characterizes the data used here.

A regression of RPV upon the square of expected inflation (EI), the square of unexpected inflation (UI), and inflation uncertainty (CVI) yields

$$\text{(5)}\quad \begin{aligned} RPV_t &= \underset{(10.56)}{0.000415} + \underset{(1.67)}{1.36(EI_t)^2} + \underset{(15.36)}{2.76(UI_t)^2} + \underset{(3.61)}{1.96CVI_t} + \upsilon_t \\ \upsilon_t &= \underset{(4.11)}{0.171\upsilon_{t-1}} + \underset{(2.72)}{0.112\upsilon_{t-2}} + \varepsilon_t \end{aligned}$$

(Adj. R-squared = 0.46)

(F-statistic for overall significance = 103.11)

The t-statistics of the coefficient estimates are given in parentheses. A second-order autoregressive specification is used for the error term. Under this specification, the

[9] Serial independence is rejected at the 1 percent level for the residuals as well as for the squared residuals. The Ljung-Box Q-statistics testing the serial independence of the residuals at lags of six and 12 months are Q(6) = 244.5 and Q(12) = 292.0, respectively. The Ljung-Box Q-statistics testing the serial independence of the squared residuals at lags of six and 12 months are Q(6) = 18.3 and Q(12) = 18.4, respectively.

conditional residuals are indistinguishable from white noise.[10] The overall regression is significant at the 1 percent level. The estimates of the coefficients on inflation uncertainty and the square of unexpected inflation are both positive and statistically significant at the 1 percent level. The estimate of the coefficient on the square of expected inflation is positive and statistically significant at the 10 percent level.

When unexpected inflation is split into positive unexpected inflation (UIP) and negative unexpected inflation (UIN), the results change somewhat. This decomposition is motivated by the symmetric role of positive and negative unexpected inflation in the Hercowitz-Cukierman model and follows Fischer (1981a, 1982) where the same decomposition is performed. A regression of RPV upon the square of expected inflation (EI), the square of positive unexpected inflation (UIP), the square of negative unexpected inflation (UIN), and inflation uncertainty (CVI) yields

$$(6)\quad \begin{aligned} RPV_t &= \underset{(10.24)}{0.000403} + \underset{(3.21)}{2.76}(EI_t)^2 + \underset{(16.30)}{2.99}(UIP_t)^2 + \underset{(2.78)}{1.08}(UIN_t)^2 + \underset{(4.14)}{2.27}CVI_t + \upsilon_t \\ \upsilon_t &= \underset{(4.64)}{0.195}\upsilon_{t-1} + \underset{(2.48)}{0.103}\upsilon_{t-2} + \varepsilon_t \end{aligned}$$

(Adj. R-squared = 0.48)

(F-statistic for overall significance = 93.04)

The t-statistics of the coefficient estimates are given in parentheses. A second-order autoregressive specification again is used for the error term. Under this specification, the conditional residuals are indistinguishable from white noise.[11] The overall regression is significant at the 1 percent level and the coefficient estimates on inflation uncertainty, the square of expected inflation, the square of positive unexpected inflation, and the square of negative unexpected inflation are all positive and significant at the 1 percent level.

Some Implications of the Estimation Results

In his Nobel lecture on inflation and unemployment, Friedman (1977) suggests that inflation may induce excess variability in relative prices that, in turn, may hamper the efficiency of the price system in coordinating economic activity. Several later writers have formalized this idea. Tommasi (1994), for instance, argues that the overall productive efficiency of the economy will decrease with RPV because RPV shields inefficient firms from price competition. If welfare decreases with RPV and if causality

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10 With an AR(2) specification for the error term, the hypothesis that the residuals are serially uncorrelated is not rejected at traditional levels of significance (Q(6) = 2.90, Q(12) = 6.68). Similarly, the hypothesis that the squared residuals are serially uncorrelated is not rejected (Q(6) = .94, Q(12) = .96).

11 With an AR(2) specification for the error term, the hypothesis that the residuals are serially uncorrelated is not rejected at traditional levels of significance (Q(6) = 2.01, Q(12) = 3.72). Similarly, the hypothesis that the squared residuals are serially uncorrelated is not rejected (Q(6) = 1.21, Q(12) = 1.23).

runs from the various aspects of inflation to RPV, then the coefficient estimates reported in equation (6) have some clear policy implications.

First, a policy of aggregate demand management that stabilizes the price level would promote welfare. Second, if some inflation were unavoidable, then a policy to keep inflation stable at some average rate would promote welfare. Third, the fact that the coefficient estimate on negative unexpected inflation is smaller than the coefficient estimate on expected inflation suggests that, given an initial position of high ongoing inflation, a policy of disinflation to a lower rate of ongoing inflation would tend to promote welfare. This is because any higher RPV caused by negative unexpected inflation during the transition period would tend to be outweighed by the lower RPV enjoyed in perpetuity at the eventual lower rate of ongoing inflation.[12]

The conclusions that can be drawn from the estimation results regarding the three theoretical models described above are as follows. First, the positive and significant coefficient estimate on inflation uncertainty means that the results are consistent with the Lucas-Barro signal-extraction model. Second, the positive and significant coefficient estimate on the square of expected inflation means that the results are consistent with the menu-costs model. Third, the results are inconsistent with the Hercowitz-Cukierman model. The Hercowitz-Cukierman model predicts that the coefficients on the squares of positive and negative unexpected inflation will be the same size, but, in fact, the estimates of the coefficients on the squares of UIP and UIN are substantially different. An F-test of the equality of the coefficients on those two variables yields an F (1, 596)-statistic of 23.52 and indicates rejection of the hypothesis of equality at the 1 percent level. Fischer (1981a, 1982) and Vanderhoff (1988) find a similar disparity between the coefficients on positive and negative unexpected inflation.

Fourth and finally, the results indicate that all three models in conjunction cannot completely explain the data. The Hercowitz-Cukierman model is the only one of three that predicts any sort of relationship between inflation surprises and RPV. The fact that we reject the Hercowitz-Cukierman model despite finding significant relationships between RPV and unexpected inflation (both positive and negative) suggests that our understanding of the inflation-RPV connection is incomplete. In fact, the results suggest that the most important part of the puzzle is missing because the coefficient estimate on the square of positive unexpected inflation is the largest of all the coefficient estimates.

Robustness Tests

Here I perform two tests of the robustness of the estimation results of the previous section. The first, a Chow test of the stability of the coefficient estimates over time, is performed because studies such as Fischer (1982) find that the relationship between inflation and RPV has changed to some degree over time. The second, a test of whether the results change when dummy variables are added for oil shocks, is performed

12 This statement abstracts from other issues such as the short-run Phillips curve in order to focus solely upon the implications of the relationship between RPV and various aspects of inflation.

because Bomberger and Makinen (1993) express the concern that the widely observed relationship between inflation and RPV may be driven by a handful of oil shocks. In particular, dummy variables are added for the 1973–1974 and 1986 oil shocks because those were times of both severe changes in the price of oil and extremely high RPV (as can be seen in Figure 2).

A Chow breakpoint test of the specification in equation (6) using the midpoint of the sample (1972.05) as the breakpoint does indicate a structural break. The hypothesis that the coefficients are identical in the pre–1972.05 and post–1972.06 periods yields an F (5, 591)-statistic of 2.85 and indicates rejection at the 1 percent level. Equations (7) and (8) report the estimation results for the pre–1972.05 and post–1972.06 periods, respectively:

$$\text{(7)}\quad \begin{aligned} RPV_t &= \underset{(11.12)}{0.000297} + \underset{(2.44)}{2.23}(EI_t)^2 + \underset{(4.86)}{1.91}(UIP_t)^2 + \underset{(6.90)}{2.08}(UIN_t)^2 + \underset{(2.30)}{1.81}CVI + \upsilon_t \\ \upsilon_t &= \underset{(2.87)}{0.166}\upsilon_{t-1} + \varepsilon_t \end{aligned}$$

(Adj. R-squared = 0.32)

(F-statistic for overall significance = 29.15)

$$\text{(8)}\quad \begin{aligned} RPV_t &= \underset{(8.16)}{0.000521} + \underset{(2.48)}{3.13}(EI_t)^2 + \underset{(11.91)}{3.05}(UIP_t)^2 + \underset{(1.36)}{0.837}(UIN_t)^2 + \underset{(3.17)}{2.27}CVI_t + \upsilon_t \\ \upsilon_t &= \underset{(3.23)}{0.192}\upsilon_{t-1} + \varepsilon_t \end{aligned}$$

(Adj. R-squared = 0.47)

(F-statistic for overall significance = 55.32)

In each case a first-order autoregressive specification for the error term is adequate to whiten the residuals.[13] Both overall regressions are significant at the 1 percent level.

The estimation results for the post–1972.06 period reported in equation (8) differ from those obtained in the previous section only in that the coefficient estimate on negative unexpected inflation is insignificant. The general conclusions of the previous section continue to hold. The estimation results are consistent with the menu-costs model and the Lucas-Barro model, but are inconsistent with the Hercowitz-Cukierman model. An F-test of the equality of the coefficients on UIP and UIN in the post–1972.06 sample yields an F (1, 300)-statistic of 13.77 and indicates rejection of the hypothesis of equality at the 1 percent level.

The estimation results for the pre–1972.05 period reported in equation (7) differ from those obtained above only in that the coefficient estimates on positive and

[13] In equation (7), the hypothesis that the residuals are serially uncorrelated is not rejected at traditional levels of significance (Q(6) = 4.29, Q(12) = 8.37), nor is the hypothesis that the squared residuals are serially uncorrelated (Q(6) = 5.39, Q(12) = 9.36). Similarly, in equation (8), the hypothesis that the residuals are serially uncorrelated is not rejected (Q(6) = 3.56, Q(12) = 4.69), nor is the hypothesis that the squared residuals are serially uncorrelated (Q(6) = 0.51, Q(12) = 0.58).

negative unexpected inflation are indistinguishable in size. An F-test of the equality of the coefficients on UIP and UIN yields an F (1, 292) statistic of 0.14 and does not indicate rejection of the hypothesis of equality. As a result, the general conclusions of the previous section must be modified in one way for the pre–1972.05 period: the estimation results from the pre–1972.05 period are consistent with all three models.

A casual inspection of the Figure 2 indicates that RPV was extremely high during the 1973–1974 period in which the price of crude oil roughly doubled in 15 months and during 1986 in which the price of crude oil fell over 50 percent in seven months. This correspondence between severe changes in oil prices and the two periods of highest RPV suggests that the estimation results of the previous section may be driven by a few oil shocks. To study this possibility, dummy variables for the 1973–1974 and 1986 oil shocks are added to the specification in equation (6). Based upon an examination of the behavior of the crude petroleum subindex of the PPI (wpu0561), the dummy variable OIL73–74 takes the value 1 during 1973.05–1974.07 and 0 elsewhere, and the dummy variable OIL86 takes value 1 during 1986.02–1986.08 and 0 elsewhere.[14] A regression of RPV upon the square of expected inflation, the square of positive unexpected inflation, the square of negative unexpected inflation, inflation uncertainty, and the dummies OIL73–74 and OIL86 yields

$$\begin{aligned} RPV_t &= \underset{(11.02)}{0.000402} + \underset{(4.07)}{3.35}(EI_t)^2 + \underset{(15.14)}{2.81}(UIP_t)^2 + \underset{(2.30)}{0.853}(UIN_t)^2 + \underset{(1.58)}{1.01}CVI_t \\ &\quad + \underset{(3.20)}{.000865}(OIL73-74) + \underset{(7.32)}{.00222}(OIL86) + \upsilon_t \qquad (9) \\ \upsilon_t &= \underset{(4.34)}{0.196}\upsilon_{t-1} + \underset{(1.54)}{0.058}\upsilon_{t-2} + \varepsilon_t \end{aligned}$$

(Adj. R-squared = 0.54)

(F-statistic for overall significance = 87.79)

The t-statistics of the coefficient estimates are given in parentheses. A second-order autoregressive specification is used for the error term. Under this specification, the conditional residuals are indistinguishable from white noise.[15]

The overall regression with the oil dummies is significant at the 1 percent level. The coefficient estimates on the 1973–1974 and 1986 oil shock dummies are positive and significant as anticipated. The other estimation results reported in equation (9)

[14] The period 1973.05–1974.07 begins and ends with oil price increases exceeding 3 percent per month and includes a total of eight months with oil price increases exceeding 3 percent. For more than a half-year before and after this period, there are no other oil price increases of this magnitude. The period 1986.02–1986.08 begins and ends with oil price decrease of more than 3 percent per month and includes a total of five months with oil price decreases of more than 3 percent per month. For more than a half-year before and after this period, there are no other oil price decreases of this magnitude.

[15] With an AR(2) specification for the error term, the hypothesis that the residuals are serially uncorrelated is not rejected (Q(6) = 10.85, Q(12) = 11.52), nor is the hypothesis that the squared residuals are serially uncorrelated (Q(6) = 5.29, Q(12) = 5.32).

differ from those obtained above only in that the coefficient estimate on inflation uncertainty is insignificant. In light of this, the general conclusions reached in the previous section must be modified slightly. The results continue to be consistent with the menu-costs model and continue to be inconsistent with the Hercowitz-Cukierman model. An F-test of the equality of the coefficient estimates on UIN and UIP yields an F (1, 595)-statistic of 27.59 and indicates rejection of the hypothesis of equality at the 1 percent level. The Lucas-Barro model, however, no longer receives the support that it did above.

Conclusion

Using monthly US producer price index data, this study finds that expected inflation, unexpected inflation (particularly when positive), and, to a lesser extent, inflation uncertainty are all positively related to relative price variability (RPV). To the extent that RPV has adverse welfare consequences and to the extent that causality runs from inflation to RPV, these findings indicate that price stability has an added welfare benefit beyond those that are traditionally recognized.

The findings shed light on three theoretical models that predict relationships between various aspects of inflation and RPV. First, the findings are consistent in all specifications with the prediction of the menu-costs model that RPV will increase with expected inflation. Second, the findings are generally consistent with the prediction of the Lucas-Barro signal-extraction model that RPV will increase with *ex ante* inflation uncertainty. The one exception occurs when dummy variables are added for oil shocks. Third, the findings are generally inconsistent with the Hercowitz-Cukierman extension of the Lucas-Barro model that predicts that RPV will increase with the magnitude of unexpected inflation, irrespective of sign.

Perhaps most importantly, the findings of this paper suggest that, together, these three models can not completely explain the data. In particular, we have no explanation for why RPV is more strongly correlated with positive unexpected inflation than with negative unexpected inflation. Models incorporating the assumption of downward rigidity in prices might bridge this gap.[16] Models capturing the essence of how money is injected into (and withdrawn from) the economy might also help explain this asymmetry.

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