Inflation and Relative Price Variability: New Evidence from U.S. Cities

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Abstract

The impact of inflation on relative price variability (RPV) is an important channel for real effects of inflation. This paper explores the inflation-RPV linkage employing a modified version of Hansen's panel threshold model to recent U.S. data. We find two significant inflation thresholds implying three different regimes where the relation between inflation and RPV differs. There are both positive and negative effects of inflation on RPV. The smallest effect of inflation on RPV is ensured if inflation is low but well above zero. If monetary policy aims at minimizing inflation's impact on relative prices, our estimates suggest that U.S. inflation should range between 1.8% and 2.8%.

Keywords: Real effects of inflation; relative price variability; inflation thresholds

JEL classification: E31, C23

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1 Introduction

There is a growing consensus that inflation has real effects on the economy through its impact on the variability of relative prices (RPV). Yet, the causes and implications of the inflation-RPV linkage are still controversial. While the bulk of empirical work finds that inflation *increases* RPV, there are also contributions suggesting inflation and RPV are *negatively* associated. This paper tries to shed more light on the empirical relationship between inflation and RPV in the United States.

Theories generating real effects of inflation through its impact on RPV are typically based on menu costs or imperfect information about the price level.[1] In both types of model, inflation increases RPV and, thus, distorts the informativeness of nominal prices. As a result, inflation-induced increases in RPV impede the efficient allocation of resources and unambiguously lead to welfare losses. These models on the real effects of inflation have been very influential for recent macroeconomics. In particular, in standard New-Keynesian DSGE models, increased relative price variability is "the root of all evil" caused by inflation, see Green (2005, p.132).

In line with these theoretical predictions, several studies have provided evidence in favor of a positive impact of inflation on RPV for various countries, see e.g. Parsley(1996), Debelle and Lamont (1997), Jaramillo (1999), Aarstol (1999), Chang and Cheng (2002), Konieczny and Skrzypacz (2005), and Nautz and Scharff (2005). However, there are notable exceptions challenging this dominant view on the inflation-RPV linkage. Reinsdorf (1994) finds a negative relationship for U.S. data from the disinflationary early 1980s. Following Lastrapes (2006), the positive relationship between U.S. inflation and RPV breaks down in the mid-eighties. Silver and Ioannidis (2001) show that RPV decreases in inflation for several European countries during the pre-EMU period. In these cases, conclusions about the welfare effects of inflation are no longer obvious. For example, if inflation decreases RPV, this may indicate inflation-enhanced, possibly welfare improving price convergence (Fielding and Mizen, 2000). In a similar vein,

[1]Given inflation, price adjustments or menu costs increase RPV by making it optimal for (heterogenous) firms to change their prices infrequently even if their real prices erode, see e.g. Rotemberg (1983). In incomplete information models introduced by Lucas (1973), noisy price information leads to misperceptions of relative price changes and inefficient supply.

inflation may decrease RPV because it ameliorates the distorting effect of nominal downward wage rigidities, see e.g. Akerlof, Dickens and Perry (1996).

A common feature of empirical contributions on the inflation-RPV linkage is that they restrict the attention to *linear* relationships.[2] However, both the mixed evidence provided by the empirical literature and the manifold aspects emphasized by theoretical models suggest that the relationship between inflation and RPV is more complex. For example, the increasing effect of inflation on RPV implied by menu cost and imperfect information models could be solely present in episodes of high inflation (Bomberger and Makinen, 1993). By contrast, the decreasing effect of inflation on RPV related to nominal downward wage rigidities should be particularly relevant if inflation is very low. If the impact of inflation on RPV results from a combination of different - partly offsetting - effects, the relation can be expected to be nonlinear. In particular, the marginal impact of inflation on RPV may differ for high and low inflation regimes. In order to account for possible regime dependent effects of inflation on RPV, this paper employs a modified version (Bick, 2007) of the threshold model introduced by Hansen (1999).

Threshold models nest the linear case, such that they can be viewed as a first, natural step to generalize linear relationships. More interestingly, the existence of inflation thresholds can be related to the recent discussion about the acceptable range of inflation. Although the Federal Reserve has never officially stated a target range of inflation, most analysts believe that the Fed has *implicit* upper and lower limits of its inflation objective, compare e.g. Thornton (2006).[3] In view of the important role of the inflation-RPV linkage for the inflation transmission mechanism, the identification of inflation thresholds could provide useful information about the appropriate location and width of an inflation targeting band.

Applying the threshold model to inflation rates from U.S. cities, we determine the number of

[2]In contrast to the traditional reduced form equations employed in the empirical literature, Lastrapes (2006) includes all individual prices in a linear VAR and estimates the cross-sectional distribution of impulse responses of these prices to measure the influence of e.g. monetary shocks on price dispersion. Fielding and Mizen (2006) suggest to examine the functional form of the inflation-RPV linkage with non-parametric methods.

[3]Explicit inflation targeting bands or critical values of inflation are used by many central banks, including the Bank of England and the European Central Bank, to facilitate the communication of monetary policy. Inflation targets may help anchoring inflation expectations and increase the transparency and accountability of the central bank. Mishkin and Westelius (2006) show that the announcement of an inflation targeting band can be interpreted as an inflation contract ameliorating the inflation bias of discretionary policy.

inflation thresholds, the threshold levels, as well as the marginal impact of inflation on RPV in the various regimes. Building on earlier work of Parsley (1996) and Debelle and Lamont (1997), our sample covers the years 1999 to 2005 when average U.S. inflation was low and stable. Yet, the panel data provides us with a sufficient variation of inflation rates in a range which should be of particular interest for assessing the current low inflation environment.

The remainder of the paper is structured as follows. Section 2 introduces the data and presents results from a linear panel regression. Section 3 applies the threshold model to the inflation-RPV linkage revealing regime dependent effects of inflation. Section 4 summarizes our main results and offers some conclusions.

2 Inflation and RPV in U.S. Cities

2.1 The Data Set

Our empirical analysis uses price data of the eight major CPI subcategories published by the Bureau of Labor Statistics (BLS) for a panel of 14 U.S. cities.[4] Specifically, we follow Debelle and Lamont (1997) and use the CPI-U index representing the expenditures by all urban consumers which can be downloaded from http://www.bls.gov/cpi/home.htm. The sample period starts in January 1998 when two new major groups were introduced in the CPI-U and ends in August 2005.[5]

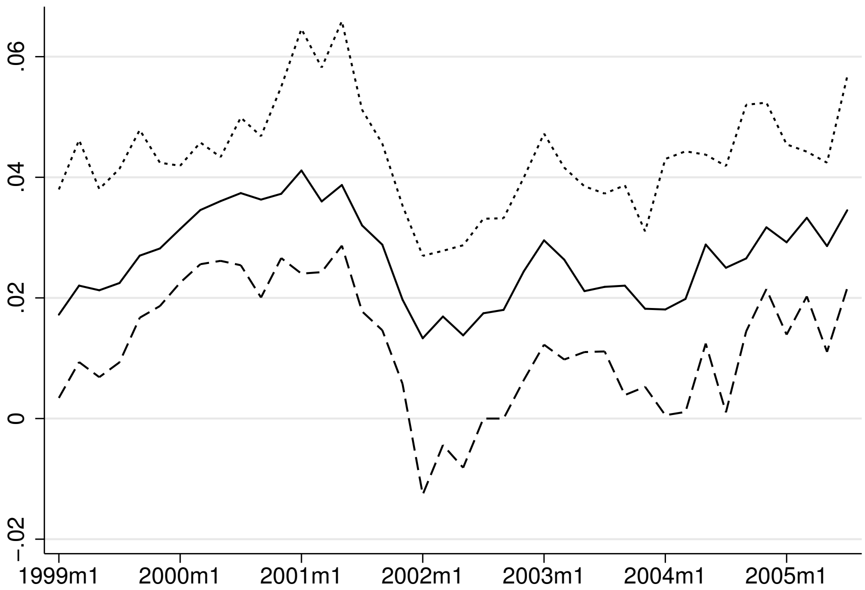
The frequency and the timing of the CPI publication differs across cities: for eleven cities data is released every second month. Only for three cities (Chicago, Los Angeles, and New York), price data is available on a monthly basis. Since the estimation of Hansen's (1999) panel threshold model requires a balanced panel, we took only the data of every second month for these three cities. Specifically, we selected the observations of the odd months because this choice implied that the number of observations in our sample from odd and even months is exactly the same.[6]

[4]These subcategories are food and beverages, housing, apparel, transportation, medical care, recreation, education and communication, and other goods and services.

[5]Replacing "entertainment", these new groups are "education and communication" and "recreation". Moreover, for three cities (Atlanta, Seattle and Washington) data were published only twice a year before 1998.

[6]Note that the lack of synchronicity in the data is not a problem because both, the traditional linear equation and the threshold model will contain no lagged variables. Data for Atlanta, Detroit, Houston, Miami, Philadelphia,

Figure 1: Inflation Rates Across U.S. Cities



Notes: Minimum, mean and maximum of yearly CPI-U inflation rates of 14 U.S. cities from 1999.01 to 2005.08. Source: BLS.

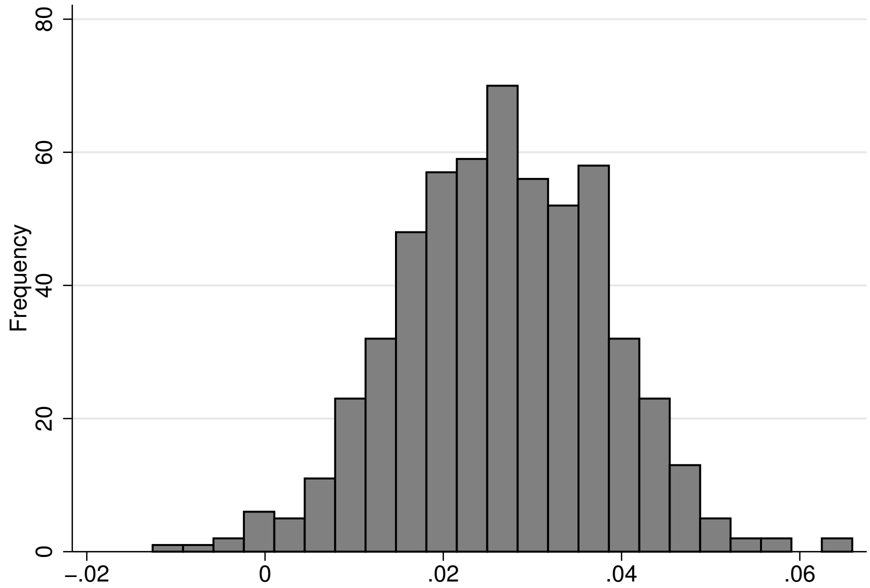
After these data adjustments we are left with $40 \times 14 = 560$ observations of yearly inflation rates, a sufficient sample size for applying panel threshold models.

U.S. inflation has been low and stable over the last years. Since 1999 the average inflation rate across U.S. cities has fluctuated around 2.7%. Figure 1 further displays the minimum and the maximum of the city-specific inflation rates indicating that inflation in U.S. cities exceeded 6% and went even below zero at least for some cities in some periods. This illustrates that inflation differentials between U.S. cities have been modest but far from negligible. Typically, inflation rates varied in a range of 3 to 4 percentage points.[7] Figure 2 reveals more information about the distribution of city inflation rates from a time-less perspective. Note that our sample provides us with a sufficient variation of inflation rates. In particular, 25% of the observed inflation rates were below 1.88% or above 3.50%, respectively.

San Francisco and Seattle is released only in even months; for Boston, Cleveland, Dallas and Washington only in odd months.

[7] Note, however, that the *persistence* of inflation differentials between U.S. cities has been relatively low. Therefore, inflation differentials between cities did not lead to significant price level divergence.

Figure 2: Distribution of City Inflation Rates in the U.S.



Notes: Yearly CPI-U inflation rates of 14 U.S. cities from 1999.01 to 2005.08.
Source: BLS.

2.2 Relative Price Variability

Following the empirical literature, we define relative price variability (RPV_{it}) for city $i = 1, \ldots, 14$ in period $t = 1, \ldots, 40$ as

$$RPV_{it} = \sqrt{\sum_{j=1}^{8} w_j (\pi_{ijt} - \pi_{it})^2} \tag{1}$$

where $\pi_{ijt} = \ln P_{ijt} - \ln P_{ijt-6}$ is the yearly inflation rate for subcategory $j = 1, \ldots, 8$ and P_{ijt} is the level of the corresponding price index. $\pi_{it} = \sum_{j=1}^{8} w_j \pi_{ijt}$ denotes the inflation rate for city i and w_j refers to the weight of the j-th subcategory in the aggregate index such that $\sum_{j=1}^{8} w_j = 1$.[8] Silver and Ioannidis (2001) introduce the coefficient of variation as an alternative measure of relative price variability. However, this RPV measure is not applicable in our sample because it includes inflation rates below zero.

[8]In the following, the RPV measure will take into account that subcategory weights are adjusted on a yearly basis. The subcategory weights can also be downloaded from http://www.bls.gov/cpi/home.htm. They are only available as averages over all cities covered in the CPI.

Table 1: The Linear Relation Between Inflation and RPV

$RPV_{it} = \alpha_i + \beta\|\pi_{it}\| + \varepsilon_{it}$ (2)	
$\widehat{\beta}$	-0.017 (0.03)
R^2	0.00
Exogeneity-Test (p-value)	0.70

Notes: Standard errors are given in parentheses. Inflation and relative price variability (RPV) for 14 U.S. cities; p-value refers to the $F-$statistic of the Davidson-MacKinnon (1993) exogeneity test. Sample period 1999.1 - 2005.08; data source: BLS.

2.3 The Linear Relation Between Inflation and RPV

Following Parsley (1996) and Debelle and Lamont (1997), we start the analysis of the empirical relation between inflation and RPV with a least squares panel regression of RPV on the absolute value of aggregate inflation with city-specific fixed effects α_i:[9]

$$RPV_{it} = \alpha_i + \beta|\pi_{it}| + \varepsilon_{it}. \tag{2}$$

The results summarized in Table 1 show that the findings of Debelle and Lamont (1997) obtained for similar data up to the mid-eighties cannot be confirmed for the recent period of low and stable inflation.[10] In line with Lastrapes (2006), the coefficient of inflation in the linear RPV equation (2) is clearly insignificant and the R^2 is virtually zero. According to the Davidson-MacKinnon (1993) exogeneity test,[11] endogeneity of inflation is not an issue in Equation (2), compare Jaramillo (1999). In the next section, we shall investigate whether relaxing the linearity assumption can reveal more information about the inflation-RPV linkage.

[9] Panel unit root tests for inflation and RPV indicate rejection of the null of an individual or common unit; these results are not presented but are available from the authors.

[10] With a view to the general inflation trends of the 1970s and the 1980s, Parsley (1996) and Debelle and Lamont (1997) include time dummies to control for shocks that hit all cities in a uniform manner. However, for the current low inflation period, including time dummies does not alter our results in a significant way.

[11] Using lagged inflation rates as instruments, a test of overidentifying restrictions could not reject the validity of the instruments.

3 Inflation Thresholds and the Inflation-RPV Linkage

3.1 The Threshold Model

Since the early findings of Parks (1978) it has been repeatedly suspected that the empirical evidence in favor of a positive link between U.S. inflation and RPV might be only due to a few high inflation periods, see Jaramillo (1999). In the following we will employ a modified version of Hansen's (1999) panel threshold model to explore whether the linear model presented in the preceding section is misspecified because the marginal impact of inflation on RPV depends on the inflation level.

To that aim, consider the following threshold model for the inflation-RPV linkage:

$$RPV_{it} = \alpha_i + \sum_{k=0}^{K-1} (\delta_{k+1} + \beta_{k+1}\pi_{it})\, I(\gamma_k < \pi_{it} \leq \gamma_{k+1}) + \beta_{K+1}\pi_{it} I(\gamma_K < \pi_{it} \leq \gamma_{K+1}) + \varepsilon_{it}, \quad (3)$$

where $\gamma_0 = -\infty$, $\gamma_{K+1} = \infty$ and I is the indicator function. Equation (3) allows for K inflation thresholds and, thus, $K + 1$ regimes. In each regime, the marginal impact of inflation (β_k) on RPV may differ. Equation (3) generalizes the panel threshold model introduced by Hansen (1999) by allowing for regime dependent intercepts (δ_k). According to Bick (2007), ignoring intercepts can lead to biased estimates of both, the thresholds and the corresponding marginal impacts.

Given the observed inflation differentials across cities, it is an important feature of the panel threshold model that at each point of time different cities are allowed to be in different inflation regimes. Hansen (1999) provides tests for the number of thresholds and estimates the threshold values, i.e. the critical inflation levels where the impact of inflation on RPV changes. The modified threshold model seems to be a natural first step to analyze potential non-linearities in the relation between inflation and RPV. In particular, the standard specification (2) uses the *absolute value* of inflation as regressor implying a specific nonlinearity in the inflation-RPV linkage which can be interpreted as an inflation threshold of zero. If (2) is correctly specified, the more general threshold model (3), where RPV responds to inflation and not to its absolute value, should identify a significant inflation threshold at zero.

Table 2: Test Procedure Establishing the Number of Thresholds

$$RPV_{it} = \alpha_i + \sum_{k=0}^{K-1} (\delta_{k+1} + \beta_{k+1}\pi_{it})\, I(\gamma_k < \pi_{it} \leq \gamma_{k+1}) + \beta_{K+1}\pi_{it} I(\gamma_K < \pi_{it} \leq \gamma_{K+1}) + \varepsilon_{it}$$

	5 % Rule	10 % Rule
No threshold (H_0: K=0)		
F_1	37.75	36.48
p-value	0.00	0.00
(10%, 5%, 1% critical values)	(11.48,13.13, 16.17)	(11.16, 12.78, 16.30)
One threshold (H_0: K=1)		
F_2	18.61	18.31
p-value	0.00	0.01
(10%, 5%, 1% critical values)	(11.24, 12.53, 16.46)	(10.61, 12.48, 16.72)
Two thresholds (H_0: K=2)		
F_3	10.64	11.68
p-value	0.15	0.07
(10%, 5%, 1% critical values)	(11.48, 12.99, 15.94)	(10.82, 12.16, 14.78)

Notes: $\gamma_0 = -\infty$, $\gamma_{K+1} = \infty$. The sequential test procedure indicates that the number of thresholds is $K = 2$. 1000 bootstrap replications were used to obtain the p-values. Following Hansen (1999), each regime is required to contain at least 5% or 10% of all observations, respectively.

3.2 The Number of Inflation Thresholds

In a first step, we applied Hansen's (1999) sequential testing procedure for determining the number of inflation thresholds. Following Hansen (1999), we require that each regime contains a minimum number of observations. Column 2 of Table 2 shows the results obtained for the 5% rule predominantly applied in empirical applications of the threshold model. For our data set, the 5% rule implies that inflation thresholds may range from 0.81% to 4.44%. The results indicate a clear rejection of a linear relation ($K = 0$) between RPV and inflation in favor of a double threshold model. Specifically, the null hypothesis of a single inflation threshold ($K = 1$) in the inflation-RPV equation can be rejected at the 1% significance level, while the hypothesis of a double threshold ($K = 2$) cannot be rejected at the 10% significance level.

It is worth noting that this conclusion appears very robust with respect to different assumptions concerning the minimum number of observations in each regime. In particular, we found that the 5% constraint is not binding which implies that e.g. adopting the less restrictive 1% rule where feasible inflation thresholds range from 0.00% to 5.24% leads to exactly identical results. This already indicates that the evidence in favor of a regime-dependent influence of inflation on RPV is not driven by a few outliers. We also performed the test adopting the unusually restrictive 10% rule. In this case, the range of feasible inflation thresholds shrinks to $[1.22\%, 4.12\%]$ which leads to slightly different values of the test statistics, see Column 3 of Table 2. Yet, the main result of the test remains unaffected. In particular, regardless of the minimum number of observations contained in each regime, Table 2 strongly suggests that the inflation-RPV linkage is characterized by two inflation thresholds and, thus, three different inflation regimes.

3.3 A Double Threshold Model for the Relation Between Inflation and RPV

In line with the results shown in Table 2, we estimated the following double threshold model:

$$RPV_{it} = \alpha_i + (\delta_1 + \beta_1 \pi_{it})\, I(\pi_{it} \leq \gamma_1) + (\delta_2 + \beta_2 \pi_{it})\, I(\gamma_1 < \pi_{it} \leq \gamma_2) + \beta_3 \pi_{it} I(\gamma_2 < \pi_{it}) + \varepsilon_{it} \tag{4}$$

Table 3 reports both the estimates obtained under the 5% and the 10% rule. Results for the 1% rule are not presented since they are exactly identical to those received for the 5% rule.

The upper part of the table shows the results for the two inflation thresholds. The results for the lower inflation threshold are virtually unaffected by the applied rule. Both, the point estimate for the threshold (1.672%) and the corresponding 95% confidence intervals are very similar. Note that the confidence interval neither contains 0.00% nor 2.00%, probably the most popular number for inflation targets.[12] The second inflation threshold estimated for the 5% rule (4.274%) exceeds the upper limit for feasible thresholds under the 10% rule (4.12%). As a consequence, the point estimate for the second threshold decreases under the 10% rule. Yet, the main conclusions about the threshold's location are very robust: according to the 95% confidence intervals the upper threshold is clearly above 2.8% and certainly below 4.4%.

[12]Note that zero would have been a feasible inflation threshold under the 1% rule.

Table 3: A Double Threshold Model for the Inflation-RPV Linkage

$$RPV_{it} = \alpha_i + (\delta_1 + \beta_1 \pi_{it}) I(\pi_{it} \leq \gamma_1) + (\delta_2 + \beta_2 \pi_{it}) I(\gamma_1 < \pi_{it} \leq \gamma_2) + \beta_3 \pi_{it} I(\gamma_2 < \pi_{it}) + \varepsilon_{it}$$

	5 % Rule	10 % Rule
Threshold estimates		
$\widehat{\gamma}_1$	1.672	1.672
95% confidence interval	[1.586, 1.803]	[1.586, 1.820]
$\widehat{\gamma}_2$	4.274	3.648
95% confidence interval	[2.852, 4.385]	[2.824, 4.102]
Regime dependent inflation coefficients:		
$\widehat{\beta}_1$	-0.595^{**} (0.13)	-0.593^{**} (0.13)
$\widehat{\beta}_2$	-0.198^{**} (0.05)	-0.189^{**} (0.07)
$\widehat{\beta}_3$	0.548^{*} (0.23)	0.712^{**} (0.14)
Regime dependent intercepts:		
$\widehat{\delta}_1$	0.028^{**} (0.01)	0.036^{**} (0.01)
$\widehat{\delta}_2$	0.026^{**} (0.02)	0.035^{**} (0.01)
R^2	0.095	0.091
Exogeneity test (p-value)	0.66	0.68
Observations in regime 1	105	105
Observations in regime 2	415	344
Observations in regime 3	40	111

Notes: **,* indicate significance at the 1%, 5% level, standard errors in parentheses. Each regime consists of at least 5% and 10% of all observations, respectively. Exogeneity test according to Davidson-McKinnon (1993).

The estimates ($\hat{\beta}_1$, $\hat{\beta}_2$, $\hat{\beta}_3$) for the marginal impact of inflation in the three inflation regimes are shown in the lower part of Table 3. In contrast to the results obtained for the linear specification, the threshold model reveals that inflation has a significant impact on RPV. Regardless of the assumption concerning the minimum number of observations, both magnitude and sign of the inflation coefficient depend on the level of inflation.

In the regimes of low and moderate inflation, the estimated coefficients do not depend on the minimum observation rule. In the low inflation regime, i.e. when inflation is below 1.672%, the marginal impact of inflation on RPV is significantly negative (-0.59) implying that a marginal increase in inflation decreases RPV. Following Akerlof et al. (1996), this response of RPV to inflation can be explained by downward nominal wage rigidities: if inflation rates are already close or even below zero, real wages that cannot be adjusted act like a real cost shock. In this regime, inflation ameliorates the distorting effect of downward nominal wage rigidities and, thus, *decreases* relative price variability.

In the intermediate inflation regime, the estimated marginal impact of inflation on RPV is still significantly negative but remains only small. When inflation is low but well above zero, the estimated marginal impact of inflation can be expected to be a combination of partly offsetting effects. While the role of e.g. downward nominal wage rigidities declines when inflation increases, other aspects of inflation which increase RPV become more important. In fact, in the high inflation regime, i.e. when inflation exceeds the upper threshold, the marginal impact of inflation is significantly positive under both the 5% ($\widehat{\beta}_3 = 0.548$) and the 10% ($\widehat{\beta}_3 = 0.712$) specification. Therefore, if inflation is sufficiently high, menu costs, imperfect information about the price level or other effects implying a positive relation between inflation and RPV become eventually dominant while other offsetting effects of inflation have faded out.

Let us close this section with some further remarks on the specification of the estimated threshold model for the inflation-RPV linkage. First, a critical assumption in Hansen's (1999) threshold model is the homoskedasticity of the errors (ε_{it}). Hansen (1999) calculated heteroskedasticity consistent standard errors which in fact revealed some heteroskedasticity in his original application. By contrast, for the relation between inflation and RPV, heteroskedasticity consistent[13] and OLS standard errors are virtually identical for all estimates. Second, following Bick (2007), we augmented the standard threshold model by regime dependent intercepts which proved to be highly significant. As as consequence, Hansen's (1999) threshold model without regime-dependent intercepts would be misspecified. Finally, note that observations of all three regimes do not belong exclusively to a small subset of cities, see Table 4 in the Appendix. Therefore, the established non-linearity of the inflation-RPV linkage does indeed result from a *regime-*dependent marginal impact of inflation and should not be captured by a city-specific inflation coefficient.

4 Concluding Remarks

The impact of inflation on relative price variability (RPV) is a major channel for real effects of inflation. Typically, the empirical evidence on the inflation-RPV linkage confirms the implications of menu cost or signal extraction models where inflation increases RPV and, thereby, decreases the information content of nominal prices. Yet, in a low inflation environment, these distorting effects of inflation might not be the whole story.

Advancing on Parsley (1996) and Debelle and Lamont (1997) we use price data from a panel of U.S. cities focussing on the recent low-inflation period. For this sample, we found that the common linear inflation-RPV equation has to be rejected in favor of a double threshold model with surprisingly small 95% confidence intervals for both inflation thresholds. Partly reconciling the mixed evidence provided by the empirical literature, the estimated inflation coefficients reveal that there are both, positive and negative effects of inflation on RPV. Inflation increases RPV only if it exceeds a critical value which is estimated to range from about 2.8% to 4.4%. By contrast inflation decreases RPV for inflation rates close to zero, or more precisely below 1.67%. In the intermediate regime, when inflation is still low but well above zero, both effects seem to be weaker or partly offset each other such that the resulting combined effect of inflation on RPV is significantly smaller.

[13]For brevity, heteroskedasticity consistent standard errors are not presented.

Our empirical results on inflation thresholds in the inflation-RPV linkage may be related to the discussion about upper and lower limits of the Fed's inflation objective. In recent years, it has become increasingly obvious that even central banks with a strong commitment to price stability are not really interested in zero inflation rates.[14] Typically, central banks prefer a more sophisticated notion of price stability. According to e.g. Blinder, Canetti, Lebow and Rudd (1998, p.98), "one prominent definition of 'price stability' is inflation so low that it ceases to be a factor in influencing people's decisions." Since relative prices are crucial for economic decisions, an acceptable band of inflation rates may be defined as that range which ensures the smallest impact of inflation on RPV. Therefore, referring to the upper and lower bounds of the 95% confidence intervals of the estimated inflation thresholds, the empirical relation between U.S. inflation and RPV suggests that U.S. inflation should range between 1.8% and 2.8%.

Generalizing the traditional linear inflation-RPV regressions, the current paper employed Hansen's (1999) panel threshold model to allow for a more complex relation between inflation and RPV. In the current model, the marginal impact of inflation jumps to a new value whenever inflation exceeds a threshold. Of course, allowing for a more gradual change of the inflation coefficient might lead to a more realistic view on the inflation-RPV linkage. Therefore, following Strickholm and Teräsvirta (2006), incorporating elements of smooth transition into a threshold model could be a natural extension and is left for future research.

[14]Several arguments point to the difficulties implied by inflation rates too close to zero. For example, positive inflation rates may ameliorate problems caused by the zero-bound for nominal interest rates, see e.g. Adam and Billi (2006).

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Appendix

Table 4: U.S. Cities and Inflation Regimes

	Low Regime	Medium Regime	High Regime
Atlanta	14 (14)	25 (20)	1 (6)
Boston	2 (2)	28 (21)	10 (17)
Chicago	11 (11)	28 (27)	1 (2)
Cleveland	14 (14)	26 (23)	0 (3)
Dallas	9 (9)	27 (23)	4 (8)
Detroit	7 (7)	32 (29)	1 (4)
Houston	10 (10)	27 (19)	3 (11)
Los Angeles	0 (0)	39 (28)	1 (12)
Miami	8 (8)	30 (25)	2 (7)
New York	1 (1)	38 (34)	1 (5)
Philadelphia	5 (5)	30 (26)	5 (9)
San Francisco	14 (14)	16 (9)	10 (17)
Seattle	9 (9)	30 (25)	1 (6)
Washington	1 (1)	39 (35)	0 (4)
Total	105 (105)	415 (344)	40 (111)

Notes: The Table shows how often a city appears in the various inflation regimes estimated for the inflation-RPV linkage under the 5% rule, compare Table 3. The respective numbers under the 10% rule are given in parantheses.