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

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The impact of inflation on relative price variability (RPV) is an important channel for real effects of inflation. With a view to the recent debate on the Federal Reserve's implicit lower and upper bounds of its inflation objective, we introduce a modified version of Hansen's panel threshold model to explore the inflation-RPV linkage in U.S. cities. We find two significant inflation thresholds and 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 percent and 2.8 percent.

JEL Codes: E31, C23.

1. Introduction

There is a growing consensus that inflation affects the economy through its impact on relative price variability (RPV). In theoretical models, the link between inflation and RPV is typically generated by menu costs or imperfect information about the price level.[1] In both cases, inflation increases RPV and, thus, distorts the information content of nominal prices. The resulting real effects of inflation

*We thank Günter Beck, Ben Craig, Juliane Scharff, Sven Schreiber, and the editor, Hyun Shin, for helpful comments. Financial support by the Monetary Stability Foundation is gratefully acknowledged. E-mails: bick@wiwi.uni-frankfurt.de, nautz@wiwi.uni-frankfurt.de.

[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.

play a predominant role for recent macroeconomics. In particular, in standard New Keynesian dynamic stochastic general equilibrium (DSGE) models, increased relative price variability is “the root of all evil” caused by inflation (see Green 2005, 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), Aarstol (1999), Jaramillo (1999), Chang and Cheng (2000), Konieczny and Skrzypacz (2005), and Nautz and Scharff (2005). Yet, there are notable exceptions: Following Lastrapes (2006), the established relationship between U.S. inflation and RPV breaks down in the mid-eighties, while Reinsdorf (1994) found that the relation is even negative during the disinflationary early 1980s. In the same vein, Fielding and Mizen (2000) and Silver and Ioannidis (2001) show that RPV decreases in inflation for several European countries.

A common feature of empirical contributions on the inflation-RPV linkage is that they restrict the attention to *linear* relationships. However, the mixed evidence provided by the empirical literature suggests that the relationship between inflation and RPV is more complex. In particular, the marginal impact of inflation on RPV may differ for high- and low-inflation regimes. For example, Jaramillo (1999) finds that U.S. inflation’s impact on RPV is stronger when it is below zero. Further evidence in favor of threshold effects of inflation on RPV is provided by Caglayan and Filiztekin (2003) for Turkey and Caraballo, Dabús, and Usabiaga (2006) for Spain and Argentina. In all these contributions, however, both the number and the location of inflation thresholds are not estimated but are imposed exogenously.

This paper sheds more light on the empirical relevance of inflation thresholds for RPV by applying a modified version of the panel threshold model introduced by Hansen (1999) to recent price data from U.S. cities. This enables us to estimate the number of inflation thresholds, the threshold levels, and the marginal impact of inflation on RPV in the various regimes. Although our sample focuses on the recent low-inflation period, the panel data provides us with a sufficient variation of inflation rates in a range that should be of particular interest for assessing the current low-inflation environment.

Threshold models nest the linear case, such that they can be viewed as a first, natural step to generalize the standard inflation-RPV equations. Of course, one may think of alternative nonlinear specifications—see, e.g., Fielding and Mizen (2008), who investigate the inflation-RPV linkage with nonparametric methods. Our particular interest in the empirical relevance of inflation thresholds is stirred by 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 Federal Reserve has *implicit* upper and lower limits of its inflation objective; see, e.g., Thornton (2006). Threshold effects of inflation are not required for explaining the increasing role of inflation targets.[2] However, 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.

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 fourteen U.S. cities from January 1998 through August 2005.[3] As a consequence, our sample of yearly inflation rates starts in January 1999.

[2]Inflation targets may help anchor 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. 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.

[3]We 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

The frequency and timing of the CPI publication differ across cities: for eleven cities data are released every second month. Only for three cities (Chicago, Los Angeles, and New York) are price data 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.[4] 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.

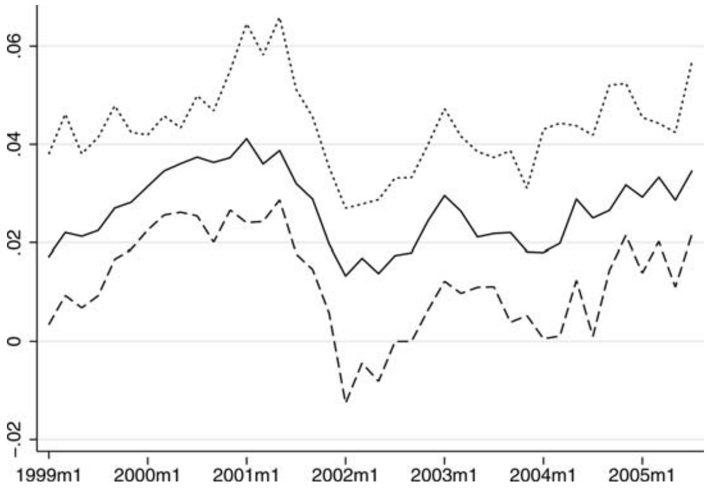
U.S. inflation has been low and stable over the last years. Since 1999 the average inflation rate across U.S. cities has fluctuated around 2.7 percent. Figure 1 further displays the minimum and the maximum of the city-specific inflation rates, indicating that inflation in U.S. cities exceeded 6 percent and even went 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.[5] Figure 2 reveals more information about the distribution of city inflation rates from a timeless perspective. Note that our sample provides us with a sufficient variation of inflation rates. In particular, 25 percent of the observed inflation rates were below 1.88 percent or above 3.50 percent.

eight subcategories are food and beverages, housing, apparel, transportation, medical care, recreation, education and communication, and other goods and services. We selected January 1998 as a starting point for two reasons: (i) before then, data for Atlanta, Seattle, and Washington were only published twice a year and (ii) two new major groups were introduced in the CPI-U in January 1998.

[4]Note that the lack of synchronicity in the data is not a problem, because both the traditional linear-equation model and the threshold model will contain no lagged variables. Data for Atlanta, Detroit, Houston, Miami, Philadelphia, San Francisco, and Seattle are released only in even months; data for Boston, Cleveland, Dallas, and Washington are released only in odd months.

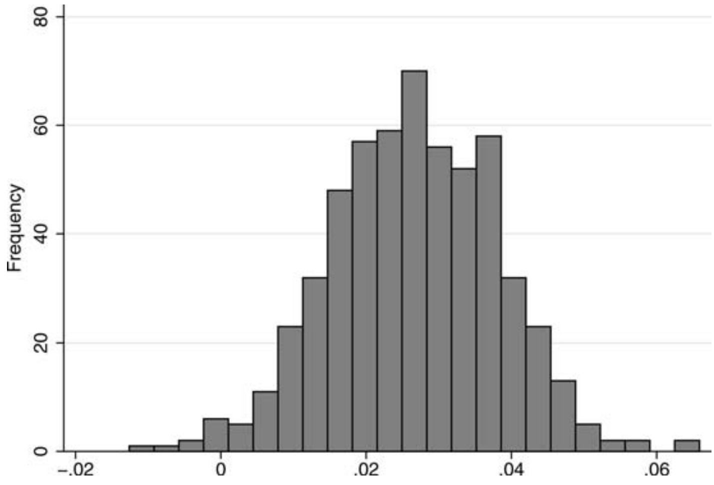
[5]The persistence of inflation differentials between U.S. cities has been relatively low. Therefore, inflation differentials between cities did not lead to significant price-level divergence.

Figure 1. Inflation Rates across U.S. Cities



Note: Minimum (dashed line), mean (solid line), and maximum (dotted line) of yearly CPI-U inflation rates of fourteen U.S. cities from January 1999 through August 2005. Source: Bureau of Labor Statistics.

Figure 2. Distribution of City Inflation Rates in the United States



Note: Yearly CPI-U inflation rates of fourteen U.S. cities from January 1999 through August 2005. Source: Bureau of Labor Statistics.

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 $\sum_{j=1}^{8} w_j = 1$.[6] 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.

2.3 The Linear Relation between Inflation and RPV

Following the empirical literature on inflation and RPV, we begin our analysis with a linear panel regression of RPV on aggregate inflation with city-specific fixed effects α_i:

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

In line with Lastrapes (2006), the coefficient of inflation in the linear RPV equation (2) is clearly insignificant for recent U.S. data.[7] In a widely used alternative specification, RPV is regressed on $|\pi_{it}|$, the absolute value of inflation; see, e.g., Parks (1978) and Jaramillo (1999). In a low-inflation environment, where some inflation rates are actually below zero (compare figures 1 and 2), this could make a difference. For our data, however, the results presented in table 1

[6] 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.

[7] As a consequence, Lastrapes (2006) suggests including all individual prices in a linear VAR to estimate the cross-sectional distribution of impulse responses of these prices to, e.g., monetary shocks.

Table 1. The Linear Relation between Inflation and RPV

	$\hat{\beta}$	R^2
$RPV_{it} = \alpha_i + \beta\pi_{it} + \varepsilon_{it}$ (2)	−0.025 (0.03)	0.00
$RPV_{it} = \alpha_i + \beta\|\pi_{it}\| + \varepsilon_{it}$ (2a)	−0.016 (0.02)	0.00

Note: Standard errors are given in parentheses. Inflation and relative price variability (RPV) for fourteen U.S. cities; sample period is January 1999 through August 2005. Data source: Bureau of Labor Statistics.

(equation (2a)) reveal that this plausible nonlinearity in inflation's impact on RPV is not supported by the data.

3. Inflation Thresholds and the Inflation-RPV Linkage

3.1 The Threshold Model

In this section, we investigate whether the linear model (2) is misspecified because the marginal impact of inflation on RPV depends on the inflation level. In contrast to recent work by, e.g., Caglayan and Filiztekin (2003) for Turkish provinces and Caraballo, Dabús, and Usabiaga (2006) for Spain and Argentina, we do not impose the number and the locations of the different inflation regimes a priori. Rather, we employ a modified version of Hansen's (1999) panel threshold model that enables us to test for the number of thresholds and to estimate the threshold values—i.e., the critical inflation levels where the impact of inflation on RPV changes.

Specifically, we consider the following threshold model for the inflation-RPV linkage:

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

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. Given the observed inflation differentials across cities, it is an additional feature of the panel threshold model that different cities are allowed to be in different inflation regimes.

It is worth noting that the panel threshold model (3) generalizes the original setup in 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.

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 1 of table 2 shows the results obtained for the 5 percent rule predominantly applied in empirical applications of the threshold model. For our data set, the 5 percent rule implies that inflation thresholds may range from 0.81 percent to 4.44 percent. 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 percent significance level, while the hypothesis of a double threshold ($K = 2$) cannot be rejected at the 10 percent significance level.

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 percent constraint is not binding, implying that adopting the less restrictive 1 percent rule (where feasible inflation thresholds range from 0 percent to 5.24 percent) leads to 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 percent 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 2 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.

Table 2. Test Procedure Establishing the Number of Thresholds

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

	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)

Note: $\gamma_0 = -\infty$, $\gamma_{K+1} = \infty$. The sequential test procedure indicates that the number of thresholds is $K = 2$. One thousand 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.

3.3 A Double-Threshold Model for the Relation between Inflation and RPV

In view of the evidence in favor of two inflation thresholds, 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. A Double-Threshold Model for the Inflation-RPV Linkage

$RPV_{it} = \alpha_i + (\delta_1 + \beta_1\pi_{it})I(\pi_{it} \leq \gamma_1) + (\delta_2 + \beta_2\pi_{it}) \times I(\gamma_1 < \pi_{it} \leq \gamma_2) + \beta_3\pi_{it}I(\gamma_2 < \pi_{it}) + \varepsilon_{it}$		
	5% Rule	**10% Rule**
Threshold Estimates		
$\hat{\gamma}_1$	1.672	1.672
95% confidence interval	[1.586, 1.803]	[1.586, 1.820]
$\hat{\gamma}_2$	4.274	3.648
95% confidence interval	[2.852, 4.385]	[2.824, 4.102]
Regime-Dependent Inflation Coefficients:		
$\hat{\beta}_1$	−0.595**	−0.593**
	(0.13)	(0.13)
$\hat{\beta}_2$	−0.198**	−0.189**
	(0.05)	(0.07)
$\hat{\beta}_3$	0.548*	0.712**
	(0.23)	(0.14)
Regime-Dependent Intercepts:		
$\hat{\delta}_1$	0.028**	0.036**
	(0.01)	(0.01)
$\hat{\delta}_2$	0.026**	0.035**
	(0.02)	(0.01)
R^2	0.095	0.091
Observations in Regime 1	105	105
Observations in Regime 2	415	344
Observations in Regime 3	40	111

Note: ** and * indicate significance at the 1% and 5% level, respectively. Standard errors are in parentheses. Each regime consists of at least 5% and 10% of all observations.

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

The upper panel 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 percent) and the corresponding 95 percent confidence intervals are very similar. Note that the confidence interval does not contain 2 percent, probably the most popular number for inflation targets. The second inflation threshold estimated for the 5 percent rule (4.274 percent) exceeds the upper limit for feasible thresholds under the 10 percent rule (4.12 percent). As a consequence, the point estimate for the second threshold decreases under the 10 percent rule. Yet, the main conclusions about the threshold's location are very robust: according to the 95 percent confidence intervals, the upper threshold is clearly above 2.8 percent and certainly below 4.4 percent. Note that observations of all three regimes do not belong exclusively to a small subset of cities; see table 4. Therefore, the established nonlinearity of the inflation-RPV linkage does indeed result from a *regime*-dependent marginal impact of inflation and cannot be captured by city-specific inflation coefficients.

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 middle panel 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. However, both magnitude and sign of the inflation coefficient depend on the level of inflation. In the low-inflation regime—i.e., when inflation is below 1.672 percent—the marginal impact of inflation on RPV is significantly negative (–0.595). Thus, a further decline of inflation would increase RPV significantly. According to, e.g., Akerlof, Dickens, and Perry (1996) or Jaramillo (1999), this effect of inflation rates close to zero may point to the presence of nominal downward wage and price rigidities. In fact, in the intermediate-inflation regime, when inflation is low but well above zero, the impact of inflation on RPV is significantly weaker. In the high-inflation regime, the marginal impact of inflation is positive under both the 5 percent ($\widehat{\beta}_3 = 0.548$) and the 10 percent ($\widehat{\beta}_3 = 0.712$) specification. When inflation exceeds an upper threshold, it seems that RPV-increasing aspects of inflation (including, e.g., menu costs and imperfect information about the price level) become eventually dominant, while RPV-decreasing aspects of inflation have faded out.

Table 4. U.S. Cities and Inflation Regimes

	Low-Inflation Regime	Medium-Inflation Regime	High-Inflation 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)

Note: The table shows how often a city appears in the various inflation regimes estimated for the inflation-RPV linkage under the 5 percent rule; compare with table 3. The respective numbers under the 10 percent rule are given in parentheses.

4. Concluding Remarks

The impact of inflation on relative price variability is a major channel for real effects of inflation. This paper focused on the recent low-inflation period using price data from a panel of U.S. cities from 1998 through 2005. 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 percent 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 percent to 4.4 percent. By contrast, inflation decreases RPV for inflation rates close to zero or, more precisely, below 1.67 percent. The weakest impact of inflation on RPV is found for the intermediate regime, when inflation is still low but well above zero.

Even central banks with a strong commitment to price stability are not really interested in zero inflation rates.[8] Typically, central banks prefer a more sophisticated notion of price stability. According to, e.g., Blinder et al. (1998, 98), "one prominent definition of 'price stability' is inflation so low that it ceases to be a factor in influencing people's decisions." Therefore, given the crucial importance of relative prices for economic decisions, an acceptable band of inflation rates should ensure the smallest impact of inflation on the variability of relative prices.

The recent literature on the importance of price rigidities revealed that there are notable differences in the frequency of price adjustments and implied durations between sectors. Golosov and Lucas (2007), and Klenow and Kryvtsov (2007) demonstrated that idiosyncratic shocks are an important factor for the price setting of firms. In accordance with Woodford (2003), the efficient level of RPV in an economy with multiple sectors is typically not zero, since it reflects relative price changes driven by fundamentals. Therefore, the optimal rate of inflation need not drive the level of RPV to zero. It is the marginal effect of inflation on RPV that has to be minimized. From this perspective, our empirical results may shed light on the location and width of an appropriate inflation-targeting band. In particular, the inflation thresholds in the inflation-RPV nexus suggest that U.S. inflation should range between 1.8 percent and 2.8 percent.

The repercussions of the introduction of an explicit targeting band on inflation's impact on relative prices are not obvious. The analysis of highly disaggregated price data indicates that the relation between inflation and the price setting of firms has been underresearched; see, e.g., Golosov and Lucas (2007). In particular, the Calvo and Taylor sticky-price models predominantly used in current New Keynesian DSGE models generate only poor predictions for the persistence and volatility of inflation; see Bils and Klenow

[8]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).

(2004). Recent evidence on the price setting of firms seems to support the relevance of inflation thresholds. According to Nakamura and Steinsson (2007), aggregate inflation plays no role in the frequency and the size of price changes during the low-inflation period 1998 through 2005. However, from 1988 to 1997, when average inflation exceeded 2.8 percent, the impact of aggregate inflation is significant and plausibly signed.

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.

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