Inflation and Relative Price Variability in the Euro Area: Evidence from a Panel Threshold Model

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Abstract

The impact of inflation on relative price variability (RPV) generates an important channel for real effects of inflation. This paper provides first evidence on the empirical relation between inflation and RPV in the Euro area. Stirred by the widespread use of inflation caps or target bands in monetary policy practice, we are particularly interested in threshold effects of inflation. In line with the ECB's definition of price stability, we find that expected inflation significantly increases RPV only if inflation is either very low (below 0.95% p.a.) or very high (above 4.96% p.a.).

Keywords: Inflation Thresholds, Relative Price Variability, Panel Threshold Models

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1 Introduction

A widely accepted definition of price stability is inflation so low as it ceases to be a factor in influencing economic decisions. In accordance with recent macroeconomic theory, an acceptable band of inflation rates should therefore ensure the smallest impact of inflation on relative prices.[1] This paper provides first evidence on the relation between inflation and relative price variability in the European Monetary Union (EMU).

Several studies found a significant impact of inflation on RPV for the US (see e.g. Parsley (1996), Debelle and Lamont (1997), Jaramillo (1999)), as well as for various European countries for the pre-EMU period (Fielding and Mizen (2000), Silver and Ioannidis (2001), Konieczny and Skrzypacz (2005), Nautz and Scharff (2005)). Typically, these contributions restrict the attention to *linear* relationships assuming that the marginal impact of inflation on RPV does not depend on the inflation level. However, a linear impact of inflation on the economy seems to be at odds with the widespread use of inflation caps or target bands in monetary policy practice. The ECB, for example, defines price stability as an inflation rate "below but close to 2%". As a consequence, the central bank's reactions to increases in the inflation rate from, say, 1% to 1.5% and 2% to 2.5% may be very different. Similar non-linear policy responses may be observed for inflation reductions if inflation is already close to zero. We are therefore particularly interested in threshold effects of inflation.[2]

A first attempt to model a non-linear relation between inflation and RPV can be found in Caglayan and Filiztekin (2003) who consider the inflation-RPV nexus for Turkish provinces. In Turkey there has been an obvious break in the inflation process around

[1] New Keynesian stochastic dynamic general equilibrium models support price stability as an outcome of optimal monetary policy only because inflation increases relative price variability (RPV) beyond its efficient level, see e.g. Woodford (2003). As Green (2005, p.132) put it, price dispersion is "the root of all evil" caused by inflation.

[2] Of course, inflation targets do not require threshold effects of inflation. Inflation targets may help anchoring inflation expectations and can increase the transparency and accountability of the central bank.

1976. Therefore, Caglayan and Filiztekin (2003) simply divide the sample in a high and a low inflation period and estimate the RPV equation for the two periods separately. Obviously, this approach should not be applied to recent Euro area data where neither the number of inflation thresholds nor the threshold levels are clear. In this case, the panel threshold model introduced by Hansen (1999, 2000) is a natural candidate. This model enables us to test for the number of inflation regimes and to estimate both the threshold levels as well as the regime-dependent marginal impact of inflation on RPV. Inflation differentials have been considerable in the Euro area. It is therefore an additional feature of Hansen's model that it allows different Euro area countries to be in different inflation regimes.

Our empirical results indicate two inflation thresholds in the European inflation-RPV nexus. Specifically, expected core inflation significantly increases RPV only if it is either very low (below 0.95% p.a.) or very high (above 4.96% p.a.). In the intermediate regime, however, when expected inflation is low but well above zero, its impact on relative prices remains only small and insignificant.

The paper is organized as follows: Section 2 introduces the data and the RPV measure. Starting with the linear relation between inflation and RPV, we discuss the role of expected versus unexpected inflation and determine core inflation as a feasible inflation measure. Section 3 briefly reviews the econometrics of the panel threshold model by Hansen (1999, 2000) which is applied to Euro area data in Section 4. Section 5 summarizes our main results and offers some conclusions.

2 The linear relation between inflation and RPV in the Euro area

2.1 Data

Our empirical analysis of the link between inflation and RPV in the Euro area employs monthly data for a complete set of subcategories of the harmonized index of consumer prices (HICP) provided by the Eurostat database. The data set contains seasonally adjusted data of twelve HICP subcategories for all 13 EMU member countries.[3] The data are available since 1996 and our sample ends in December 2006.

Following the empirical literature, the variability of relative price changes for country i in period t (RPV_{it}) is defined as the square root of the weighted sum of squared deviations of subcategory-inflation π_{ijt} around the average inflation for country i (π_{it}), i.e.

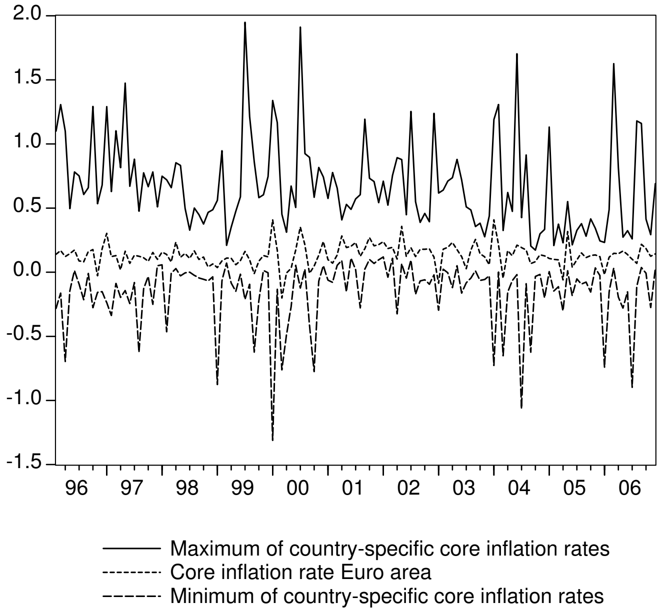
$$RPV_{it} = \sqrt{\sum_{j=1}^{12} w_{ijt}(\pi_{ijt} - \pi_{it})^2}$$

where $\pi_{ijt} = 100\Delta lnP_{ijt}$ and P_{ijt} is the price index of the jth subcategory in country i in period t. This definition of RPV is also compatible with the price dispersion measure used in theoretical models, see e.g. Woodford (2003, p.399). w_{ijt} denotes the country-specific weight of the jth subcategory in the aggregate index so that $P_{it} = \sum_{j=1}^{12} w_{ijt}P_{ijt}$ gives the aggregate price level in country i and the inflation rate π_{it} is $100\Delta lnP_{it}$. The country-specific weights are adjusted on a yearly basis by Eurostat.

The inflation-RPV relation can be distorted by supply shocks which jointly affect inflation and relative price variability. In this case, a regression of RPV on inflation will lead to correlation between inflation and the error term which would imply that inflation

[3] In 2007, these are Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Slovenia, and Spain. The HICP subcategories are food and non-alcoholic beverages; alcoholic beverages, tobacco and narcotics; clothing and footwear; housing, water, electricity, gas and other fuels; furnishings, household equipment and routine maintenance of the house; health; transport; communication; recreation and culture; education; restaurants and hotels; miscellaneous goods and services. The data are seasonally adjusted by the Census X11-method.

Figure 1: Minimum and maximum of core inflation in the Euro area



Notes: Sample 1996–2006. Minimum (maximum) of monthly, seasonally adjusted core inflation rates is defined as the Euro area wide minimum (maximum) of national inflation rates. Source: Eurostat.

may not be longer regarded as exogenous. Following e.g. Jaramillo (1999) the resulting endogeneity bias can be avoided by the application of core inflation as explanatory variable. Eurostat publishes core inflation defined as HICP inflation without food and energy prices, i.e. prices that are particularly driven by supply side shocks, compare ECB (2005).

Since the mid-nineties, Euro area inflation has been at a very moderate level but there were also periods with negative or relatively high inflation rates for some countries, see Appendix. Illustrating the significance of inflation differentials in the Euro area, Figure 1 displays the minimum and the maximum of country-specific core inflation rates.

2.2 Inflation and RPV: the basic relationship

Following Parsley's (1996) analysis of the inflation-RPV link for US cities, let us include country-specific fixed effects α_i in a least squares panel regression of RPV on inflation:

$$RPV_{it} = \alpha_i + \beta|\pi_{it}| + \varepsilon_{it}. \tag{1}$$

The results for the fixed-effects estimation (1) for the two alternative inflation measures are reported in Table 1. In both cases, inflation has a significant and positive impact on RPV. However, the Davidson-MacKinnon test indicates that the estimates based on headline inflation have to be interpreted with caution.[4] As a consequence, the following empirical analysis on the RPV-inflation linkage will focus on the core inflation rate. Euro area's headline and core inflation show a concurrent pattern and their difference has been typically small for most periods, compare ECB (2005). In particular, our main results will not depend on the particular choice of the inflation measure.

2.3 Expected inflation versus unexpected inflation

According to the simple linear specification (1), the impact of inflation on RPV does not depend on inflation expectations. However, a different role of expected and unexpected inflation is not only found empirically but is also suggested by various theories explaining the inflation-RPV link. While menu cost models by e.g. Rotemberg (1983) emphasize a positive relationship between RPV and *expected* inflation, only *unexpected* inflation affects RPV in Lucas-type misperception models, see e.g. Hercowitz (1981).

In the following, we therefore advance on the basic inflation-RPV relationship and allow for different coefficients on expected (π^e_{it}) and unexpected ($\pi_{it} - \pi^e_{it}$) inflation:

$$RPV_{it} = \alpha_i + \beta_1|\pi^e_{it}| + \beta_2|\pi_{it} - \pi^e_{it}| + \varepsilon_{it}, \tag{2}$$

[4] The Davidson-MacKinnon test computes a test of exogeneity for a fixed-effect regression estimated via instrumental variables, see Davidson and MacKinnon (1993). A rejection of the null hypothesis indicates that endogenous regressors' effects on the estimates are meaningful, and instrumental variables techniques are required.

Table 1: The linear relation between inflation and RPV in the Euro area

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

	headline inflation	core inflation
$\widehat{\beta}$	0.61^{**} (0.04)	0.67^{**} (0.04)
Exogeneity test (F-statistic)	3.21 [0.07]	0.07 [0.79]
Serial corr. test (F-statistic)	0.60 [0.45]	0.05 [0.82]
$\bar{R}^2$	0.22	0.24
Observations	1703	1703
Countries	13	13

Notes: ** indicate significance at the 1% level. Standard errors are given in parentheses, p-values in brackets. The exogeneity test tests the null hypothesis that a least squares estimator yields consistent estimates, see Davidson-MacKinnon (1993). Panel serial correlation test according to Wooldridge (2002). Sample: 1996–2006.

where α_i are again the fixed effects for each country. In practice many measures of inflation expectations exist, including the forecasts of professional economists, results from surveys of consumers or information extracted from financial markets. In spite of the increasing importance and quality of this kind of data, survey data is not available for all EMU member countries over the whole sample period and on a monthly basis. In particular, there are no surveys on expectations about core inflation. Therefore, we follow e.g. Silver and Ioannidis (2001) and Konieczny and Skrzypacz (2005) and use country-specific AR(12) time series representations to forecast inflation. In fact, beating the forecasting performance of univariate time series models of inflation is typically not an easy task, particulary over a monthly forecast horizon, see e.g. Stock and Watson (2007).[5]

[5] The country-specific inflation forecasts are shown in the Appendix. Since we found no evidence on time varying inflation uncertainty in our sample, inflation uncertainty does not appear in the RPV equation as an additional regressor, compare Neumann and von Hagen (1991).

Table 2: The effects of expected and unexpected inflation on RPV

	$RPV_{it} = \alpha_i + \beta_1 \lvert\pi^e_{it}\rvert + \beta_2 \lvert\pi_{it} - \pi^e_{it}\rvert + \varepsilon_{it}$	$RPV_{it} = \alpha_i + \beta_1 \lvert\pi^e_{it}\rvert + \beta_2 (\pi_{it} - \pi^e_{it})^+ + \beta_3 \lvert(\pi_{it} - \pi^e_{it})^-\rvert + \varepsilon_{it}$
$\widehat{\beta}_1$	0.35^{**} (0.08)	0.35^{**} (0.08)
$\widehat{\beta}_2$	1.02^{**} (0.06)	1.02^{**} (0.06)
$\widehat{\beta}_3$		1.03^{**} (0.08)
$F(\beta_1 = \beta_2)$	48.22 [0.00]	
$F(\beta_2 = \beta_3)$		0.04 [0.85]
serial corr. (F-stat.)	0.28 [0.61]	0.28 [0.61]
$\bar{R}^2$	0.28	0.28
Obs.	1547	1547
Countries	13	13

Notes: Expected and unexpected inflation are based on an AR(12) forecast of core inflation. ** indicate significance at the 1% level. Standard errors are given in parentheses, p-values in brackets. $F(\beta_i = \beta_j)$ indicates the F-statistic testing $H_0 : \beta_i = \beta_j$. Panel serial correlation test according to Wooldridge (2002). Sample: 1996-2006.

The first column of Table 2 shows the results for the fixed effects estimation of Equation (2). In line with the findings for Germany obtained by Nautz and Scharff (2005), the impact of unexpected inflation is significantly stronger in the Euro area. Nautz and Scharff (2005) argue that the influence of expected inflation in Germany disappeared because a credible monetary policy had stabilized inflationary expectations on a low level. In fact, Konieczny and Skrzypacz (2005) establish a more pronounced effect of expected inflation on RPV during the transition of Poland from a planned to a market economy when inflation expectations were relatively high. This suggests that the impact of expected inflation on RPV depends on the level of inflation. In the following, our main interest is therefore in the analysis of potential threshold effects of expected inflation.

For the US, Aarstol (1999) finds that the effect of inflation on RPV is more pronounced when inflation is unexpectedly high, i.e. when unexpected inflation is positive. Following this approach, we regress RPV on expected as well as positive and negative unexpected inflation:

$$RPV_{it} = \alpha_i + \beta_1|\pi_{it}^e| + \beta_2(\pi_{it} - \pi_{it}^e)^+ + \beta_3|(\pi_{it} - \pi_{it}^e)^-| + \varepsilon_{it}, \tag{3}$$

where $(\pi_{it} - \pi_{it}^e)^+ = (\pi_{it} - \pi_{it}^e)$ if $(\pi_{it} - \pi_{it}^e) \geq 0$ and $(\pi_{it} - \pi_{it}^e)^- = (\pi_{it} - \pi_{it}^e)$ if $(\pi_{it} - \pi_{it}^e) \leq 0$ (zero otherwise). The fixed effects estimation of Equation (3) (reported in the second column of Table 2) does not indicate an asymmetric impact of unexpected inflation on RPV. Thus, Equation (2) shall serve as the starting point for the analysis of inflation thresholds in the relation between inflation and RPV in the Euro area.

3 The Panel Threshold Model

Hansen (1999, 2000) 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. In this section, we briefly recall how to estimate and evaluate single and multiple panel threshold models.

3.1 The single threshold model

Consider the following single threshold model

$$y_{it} = \alpha_i + \beta_1' x_{it} I(q_{it} \leq \gamma) + \beta_2' x_{it} I(q_{it} > \gamma) + \varepsilon_{it}, \tag{4}$$

for a balanced panel where the subscript i stands for the cross-sections with $1 \leq i \leq N$ and t indexes time $(1 \leq t \leq T)$. $I(\cdot)$ is an indicator function and the error term ε_{it} is independent and identically distributed with zero mean and finite variance σ^2. The dependent variable y_{it} and the threshold variable q_{it} are scalar, the regressor x_{it} is a k-dimensional vector of exogenous variables. x_{it} and y_{it} are assumed to be stationary variables. x_{it} may contain variables with slope coefficients constrained to be the same in the two regimes which have no effect on the following distribution theory. If the threshold variable q_{it} is below or above a certain value of q_{it}, namely γ, then the regressor x_{it} has a different impact on y_{it} represented by coefficients $\beta_1 \neq \beta_2$. The threshold variable q_{it} may be an element of x_{it} but this is not necessarily the case. In our application y_{it} is RPV and a natural choice of q_{it} is a measure of inflation. x_{it} contains expected and unexpected inflation.

Hansen (1999, 2000) chooses a fixed effects approach to estimate Equation (4). For given γ, the slope coefficient β can be estimated by ordinary least squares in a first step. In a second step, the estimator for the threshold $\hat{\gamma}$ is achieved by minimizing the sum of squared errors, i.e. $\hat{\gamma} = \underset{\gamma}{\operatorname{argmin}}\ S(\gamma)$, and the estimate for the slope coefficient is obtained by $\hat{\beta} = \hat{\beta}(\hat{\gamma})$.

Having estimated the threshold $\hat{\gamma}$, it is important to check whether it is in fact statistically significant. Obviously, the null hypothesis "no threshold effect in Equation (4)" is equivalent to $H_0: \ \beta_1 = \beta_2$. However, standard tests have non-standard distributions since the threshold is not identified under H_0. Therefore, Hansen (1996) suggests a bootstrap method to simulate the asymptotic distribution of the likelihood ratio test. The bootstrap procedure is applied by Hansen (1999) to a large number of cross sections (N) but only a few time periods. In our application, the bootstrap procedure has to be modified because the number of countries is only ten while T is large.[6] Hansen (2000) uses the likelihood ratio statistic (denoted by F_1) for tests on γ to form valid

[6] Hansen (1999) groups the regression residuals by individual $\hat{\varepsilon}_i^* = \{\hat{\varepsilon}_{i1}^*, \hat{\varepsilon}_{i2}^*, \ldots, \hat{\varepsilon}_{iT}^*\}$ and takes the sample $\{\hat{\varepsilon}_1^*, \hat{\varepsilon}_2^*, \ldots, \hat{\varepsilon}_N^*\}$ with size N as the empirical distribution. Since N is limited but T is large in our empirical analysis, we treat the sample $\{\hat{\varepsilon}_{11}^*, \ldots, \hat{\varepsilon}_{1T}^*, \ldots, \hat{\varepsilon}_{i1}^*, \ldots, \hat{\varepsilon}_{iT}^*, \ldots, \hat{\varepsilon}_{N1}^*, \ldots, \hat{\varepsilon}_{NT}^*\}$ as the empirical distribution to be used for bootstrapping. For the bootstrap procedure, the variable x_{it} and the threshold variable q_{it} are given, i.e. their values are fixed in repeated bootstrap samples. We take with replacement a sample of size NT from the empirical distribution and create a bootstrap sample under the null hypothesis of no threshold. This bootstrap sample is used to estimate the model under H_0 and H_1 and to calculate the bootstrap value of the likelihood ratio statistic. This procedure is frequently repeated – 1000 bootstrap replications in our application – and the bootstrap estimate of the asymptotic p-value under H_0 is the percentage of draws for which the simulated likelihood ratio statistic exceeds the actual statistic. The null hypothesis of no threshold effect is rejected if the p-value is smaller than the desired significance level.

asymptotic confidence intervals for γ ($T \to \infty$ or $N \to \infty$). Note that these confidence intervals for γ can be highly asymmetric.

3.2 Multiple thresholds

To recall the testing and estimation procedure in case of *multiple* thresholds, consider the double threshold model

$$y_{it} = \alpha_i + \beta_1' x_{it} I(q_{it} \leq \gamma_1) + \beta_2' x_{it} I(\gamma_1 < q_{it} \leq \gamma_2) + \beta_3' x_{it} I(\gamma_2 < q_{it}) + \varepsilon_{it} \quad (5)$$

with $\gamma_1 < \gamma_2$. The sum of squared residuals $S(\gamma_1, \gamma_2)$ can be calculated as in the single threshold model and the joint least squares estimates of (γ_1, γ_2) are the values which jointly minimize $S(\gamma_1, \gamma_2)$. Since a grid search over (γ_1, γ_2) requires approximately $(NT)^2$ regressions, it is important to note that sequential estimation is consistent, see e.g. Hansen (1999). These results can be generalized to higher order threshold models.

Bai (1997) develops a sequential testing procedure to determine the number of significant thresholds in a multiple threshold model. If F_1 implies that the null of no threshold has to be rejected, the likelihood ratio statistic F_2 discriminates between one and two thresholds using the difference between the minimized sum of squares obtained from the two competing threshold models. The bootstrap procedure to approximate the asymptotic p-value for this likelihood ratio test works as for the single-threshold case. The sequential testing sequence stops if – according to the likelihood ratio statistic F_K – the null of a maximum number of $(K-1)$ thresholds is rejected but the null of at most K thresholds is not.[7]

According to Bai (1997), the threshold estimators in the multiple threshold model have the same asymptotic distributions as the threshold estimate in the single threshold model. Therefore, the confidence intervals for multiple threshold parameters are constructed in the same way as in the single threshold case.

[7] The GAUSS program underlying this analysis is based on the GAUSS code by Bruce Hansen, see http://www.ssc.wisc.edu/∼bhansen/.

4 Inflation thresholds and RPV: Empirical results for the Euro area

4.1 Model specification

Let us now apply the panel-threshold model to the analysis of the relationship between RPV and inflation in the Euro area. According to Section 2, we use the linear specification (2) that allows for a different impact of expected (π^e_{it}) and unexpected ($\pi_{it} - \pi^e_{it}$) core inflation on RPV as the starting point of our analysis. Using the notation introduced in Section 3, we have $y_{it} = RPV_{it}$ and $x_{it} = (|\pi^e_{it}|, |\pi_{it} - \pi^e_{it}|)$. In order to shed more light on the findings of Konieczny and Skrzypacz (2005) and Nautz and Scharff (2005), we focus on potential threshold effects of expected inflation (i.e. $q_{it} = \pi^e_{it}$). Specifically, we consider the following class of threshold equations:

$$RPV_{it} = \alpha_i + \sum_{k=0}^{K} \beta_{k+1} |\pi^e_{it}| I(\gamma_k < \pi^e_{it} \leq \gamma_{k+1}) + \delta |\pi_{it} - \pi^e_{it}| + \varepsilon_{it}, \qquad (6)$$

where $\gamma_0 = -\infty$, $\gamma_{K+1} = \infty$, K is the number of thresholds and, thus, $(K+1)$ the number of inflation regimes.[8]

4.2 The number of inflation regimes

Table 3 presents the results of the test procedure determining the number of inflation thresholds for the whole sample (1996-2006) and the sub-period associated with the actual EMU-period (1999-2006). Note that the results are virtually identical. For both samples, the p-values associated to F_1 and F_2 provide strong evidence in favor of two thresholds and, thus, three inflation regimes in the European inflation-RPV relation.

Following Hansen (1999), the tests assume that at least 5% of all observations lie in each regime. Note that this restriction is not binding (compare Table 4 and the number

[8] Inflation thresholds can be confirmed in a comprehensive sensitivity analysis for alternative specifications, including subsets of countries, alternative threshold variables and inflation measures. For brevity, these results are not presented but are available on request.

Table 3: Test procedure establishing the number of inflation regimes

$RPV_{it} = \alpha_i + \sum_{k=0}^{K} \beta_{k+1} \lvert\pi_{it}^e\rvert I(\gamma_k < \pi_{it}^e \leq \gamma_{k+1}) + \delta \lvert\pi_{it} - \pi_{it}^e\rvert + \varepsilon_{it}$		
	1996–2006	1999–2006
H_0: no threshold (K=0)		
F_1	59.60	49.97
p-value	0.00	0.00
(10%, 5%, 1% critical values)	(8.47, 9.91, 13.42)	(8.38, 9.90, 13.79)
H_0: at most one threshold (K=1)		
F_2	17.72	25.54
p-value	0.01	0.00
(10%, 5%, 1% critical values)	(9.02, 11.39, 14.89)	(8.94, 10.28, 14.28)
H_0: at most two thresholds (K=2)		
F_3	4.33	2.97
p-value	0.68	0.88
(10%, 5%, 1% critical values)	(10.30, 12.17, 16.19)	(9.54, 11.15, 14.59)

Notes: The threshold variable π_{it}^e is expected core inflation. $\gamma_0 = -\infty$, $\gamma_{K+1} = \infty$. The sequential test procedure indicates that the number of thresholds is $K = 2$. 1000 bootstrap replications were used to obtain the p-values. The required minimum number of observations in each regime is 77 (53), i.e. 5% of all observations.

of observations in each regime) which already indicates that the rejection of the linear model cannot be explained by the effects of some outliers.

4.3 A double threshold model of the inflation-RPV nexus in the Euro area

Having established two inflation thresholds in the inflation-RPV nexus, Table 4 presents the results obtained for the corresponding double threshold model. The estimated thresholds and the 95% confidence intervals are reported in the upper part of Table 4. As in the previous section, we also present the estimates for the subsample starting in 1999. Since the results for both samples are very similar, let us concentrate on the estimates obtained for the whole sample period. The point estimates of the two thresholds for expected core inflation are 0.0421 and 0.4743. According to the 95% confidence intervals, the lower inflation threshold does not exceed 0.0793 while the upper threshold value is at least 0.4133. These numbers refer to *monthly* rates of inflation implying annualized threshold rates of about 0.95% and 4.96%, respectively.

Table 4: A double threshold model of the inflation-RPV nexus in Euro area

$$RPV_{it} = \alpha_i + \beta_1|\pi^e_{it}|I(\pi^e_{it} \leq \gamma_1) + \beta_2|\pi^e_{it}|I(\gamma_1 < \pi^e_{it} \leq \gamma_2) + \beta_3|\pi^e_{it}|I(\gamma_2 < \pi^e_{it}) + \delta|\pi_{it} - \pi^e_{it}| + \varepsilon_{it}$$

	1996–2006	1999–2006
Threshold estimates		
$\widehat{\gamma}_1$	0.0421	0.06
95% confidence interval	[0.0184, 0.0793]	[0.0178, 0.0793]
$\widehat{\gamma}_2$	0.4743	0.4715
95% confidence interval	[0.4133, 0.4946]	[0.4421, 0.4784]
Regression estimates		
$\widehat{\beta}_1$	1.73^{**} (0.22)	1.72^{**} (0.23)
$\widehat{\beta}_2$	-0.03 (0.10)	0.02 (0.11)
$\widehat{\beta}_3$	0.40^{**} (0.08)	0.56^{**} (0.09)
$\widehat{\delta}$	0.99^{**} (0.05)	1.00^{**} (0.06)
Serial correlation test (F-stat.)	0.01 [0.91]	0.00 [0.98]
Observations in regime 1	120	153
Observations in regime 2	1337	1030
Observations in regime 3	90	65

Notes: The threshold variable π^e_{it} is monthly expected core inflation. Standard errors are given in parentheses, p-values in brackets. ** indicate significance at the 1% level. Panel serial correlation test according to Wooldridge (2002).

The estimates ($\hat{\beta}_1$, $\hat{\beta}_2$, $\hat{\beta}_3$) for the marginal impact of expected inflation in the three inflation regimes can be found in the lower part of Table 4. Reconciling the seemingly conflicting evidence obtained for high and low inflation countries, the coefficients of expected inflation are highly significant in both outer regimes. However, in accordance with the evidence obtained by Nautz and Scharff (2005) for Germany, the coefficient in the intermediate regime, where inflation is low but well above zero, is both economically and statistically insignificant. Irrespective of the sample period, there is a significant impact of expected inflation on RPV in the Euro area if inflation is either very low or very high.

5 Concluding Remarks

The impact of inflation on relative price variability (RPV) generates an important channel for real effects of inflation. If inflation drives RPV above its efficient level, the information content of nominal prices is reduced impeding an efficient allocation of resources. This paper reexamined the empirical relationship between inflation and RPV in the Euro area focusing on threshold effects of inflation.

Employing the panel threshold model proposed by Hansen (2000), we found that RPV increases in expected inflation only if expected inflation gets either too high ($\geq 4.96\%$ p.a.) or too low (≤ 0.95 % p.a.). In the intermediate inflation regime, when inflation is low but well above zero, there is no significant impact of expected inflation. In line with the ECB's notion of price stability, this suggests that in this regime inflation ceases to be a factor in influencing relative prices and, thereby, in distorting economic decisions.

Many central banks, including the ECB and the Bank of England, communicate their monetary policy strategy by use of inflation caps or target bands. Although, the usefulness of inflation targets does not require inflation thresholds, various approaches have been proposed to explain the implied non-linearity in the behavior of central banks. For example, Athey, Atkeson and Kehoe (2005) and Mishkin and Westelius (2006) show that inflation target bands can solve the time-consistency problem of optimal monetary policy. Orphanides and Wieland (2000) derive inflation band targeting as optimal policy when the structure of the economy exhibits zone-linearity. The current paper presented new evidence pointing to non-linear effects of inflation on the economy via its influence on the variability of relative prices. Further research is needed to assess whether the estimated inflation thresholds could even serve as guidelines for the determination of the width and the location of the optimal inflation band.

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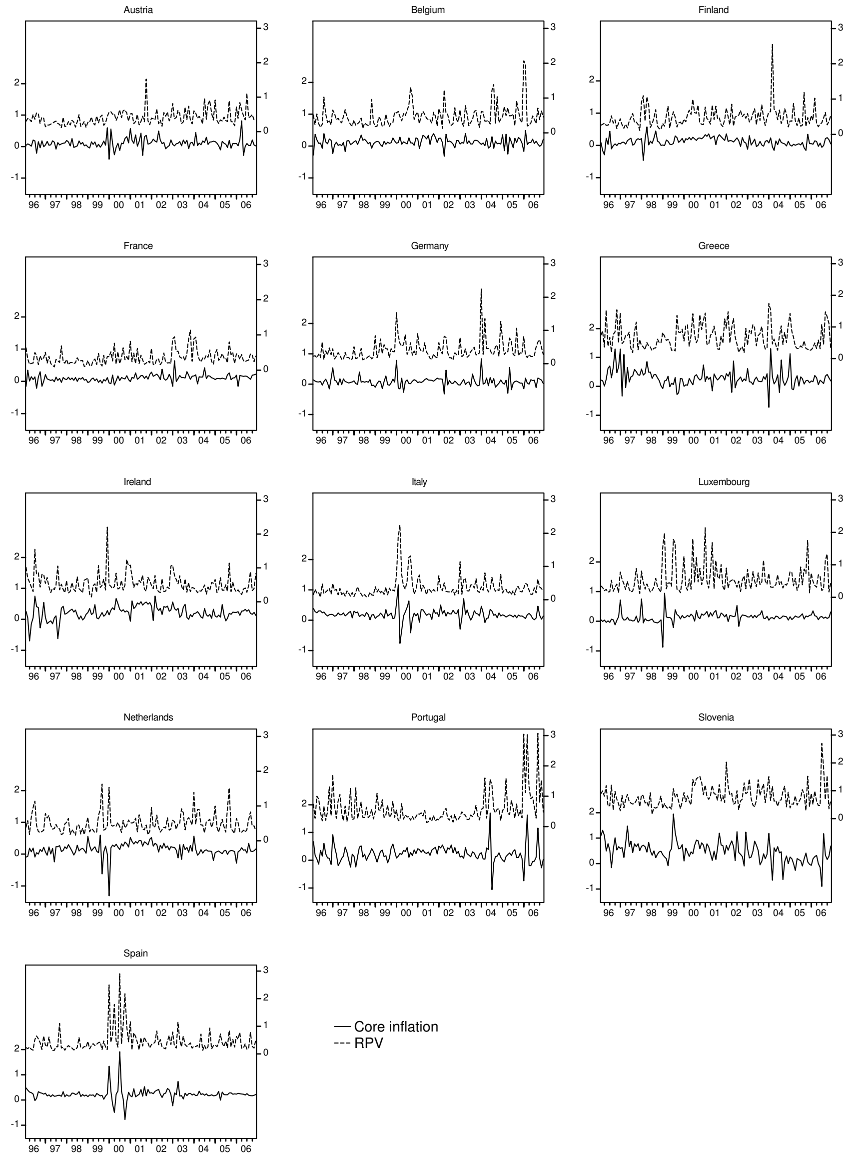
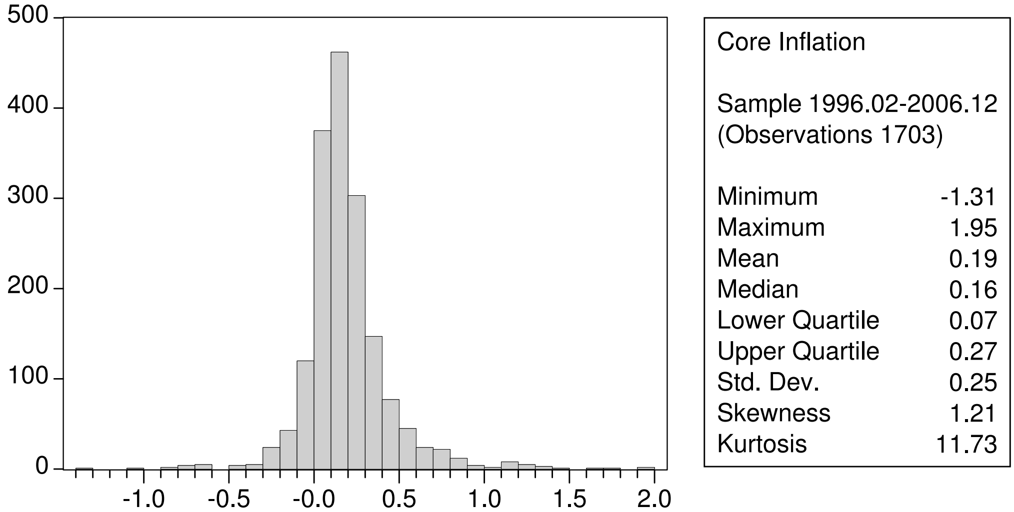
A Appendix

Figure 2: Core inflation and RPV



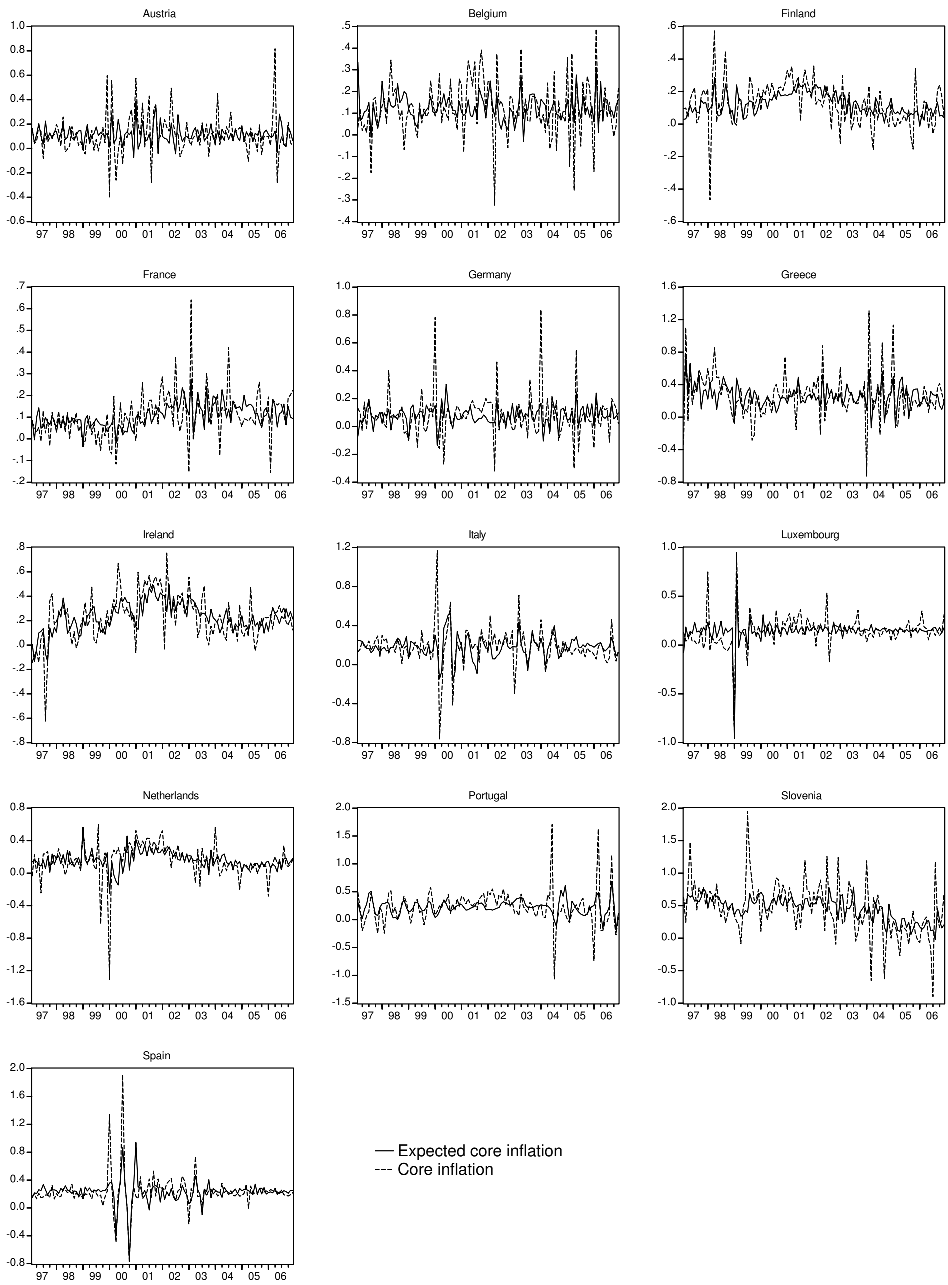
Notes: Monthly core inflation rates and monthly RPV (percentage points, seasonally adjusted) confer to the HICP, 1996–2006. Source: Eurostat.

Figure 3: Distribution of national core inflation rates in the Euro area



Notes: Monthly national core inflation rates (percentage points, seasonally adjusted) confer to the HICP, 1996–2006. Source: Eurostat.

Figure 4: Core inflation and expected core inflation



Notes: Monthly core inflation rates (percentage points, seasonally adjusted) confer to the HICP. Expected inflation is based on an AR(12) inflation forecast of core inflation. For brevity, results from country-specific forecast equations are not presented but are available on request.