Monetary Regimes, Price Dispersion and Optimal Inflation: Evidence from Spain

Mª Ángeles Caraballo. Universidad de Sevilla
Carlos Dabús. Universidad Nacional del Sur (Argentina) and CONICET

Abstract

This paper analyzes the relation between inflation and relative price variability for Spain across two different monetary regimes: before and after the entry into the European Monetary Union. We use disaggregated monthly price data of Consumer Price Index for the 1987.01-2009.08 period. Our findings indicate that the overall relation presents significant structural changes and a U-shape functional profile, which yields an annual optimal inflation rate around 4%. This is compatible with previous researches and higher than the inflation target proposed by the European Central Bank. In turn, the key link underlying the inflation-*RPV* relationship is unexpected inflation, and when it is zero *RPV* is minimal. Our results support the extended signal extraction model, and suggest that monetary policy matters: the welfare costs of inflation caused by the distorting impact of inflation on *RPV* can be removed with a credible and predictable inflation targeting policy.

Keywords: monetary regimes, optimal inflation, relative price variability.

JEL classification: E31

Authors

Mª Ángeles Caraballo. (Corresponding author)
Universidad de Sevilla. Departamento de Economía e Historia Económica. Avda. Ramón y Cajal nº1, 41018, Sevilla (Spain). E-mail: mcaraba@us.es. Phone: 0034-954557535. Fax: 0034-954556068

Carlos Dabús. CONICET and Universidad Nacional del Sur (Argentina), Departamento de Economía. 12 de octubre y San Juan, 7º. Bahía Blanca, 8000 (Argentina). E-mail: cdabus@criba.edu.ar. Phone: 0054-0291-4595138.

Acknowledgments
The authors acknowledge financial support from Junta de Andalucía (CICE, Proyecto de Excelencia SEJ-4546).

1. Introduction

There is a clear consensus in the literature on the welfare costs caused by the distorting impact of inflation on relative price variability (*RPV*). Several theoretical approaches try to explain the links underlying the relationship between inflation and *RPV*: search and menu cost models emphasize the role of expected inflation, while the Lucas-type incomplete information approach argues that non-neutrality is explained by uncertainty and unexpected inflation. Search models argue that buyers have incomplete information. Thus, higher expected inflation has two opposite effects: lowers the value of fiat money, which increases sellers' market power and *RPV*, but increases search benefits, which lowers sellers' market power, and then *RPV*. However, at higher inflation the *RPV* increasing effect should dominate, i.e. expected inflation may increase *RPV* only if it exceeds a critical value.

Menu cost models assume that nominal price changes are subject to price adjustment costs, which provoke that firms set prices discontinuously, according to an (S, s) pricing rule -see Sheshinski and Weiss (1977), Rotemberg (1983), Caplin and Spulber (1987) and Caplin and Leahy (1991)-. Hence, nominal prices are changed only when the real price hits a lower threshold, *s*, so that the new real price equals a higher return point *S*. In turn, as Caglayan et al. (2008) and Becker and Nautz (2009a) argue, the crucial point is that expected inflation induces to a higher width of the (S, s) band to conserve on menu costs. Thus, price dispersion increases and the frequency of price changes is reduced. During deflationary periods the model works in reverse, so *RPV* is increasing in the absolute value of expected inflation (i.e., the relationship is V-shaped). Furthermore, an increase in firms' menu cost or consumers' search cost also widens firms' band, increases dispersion, and lowers the frequency of price changes. And with different menu costs among firms or firms experiencing specific shocks, staggered price setting will arise exacerbating the effect of higher inflation on *RPV*.

Finally, the signal-extraction model proposed by Lucas (1973) and Barro (1976) states that *ex ante* inflation uncertainty generates "misperceptions" of absolute and relative prices, creating confusion between aggregate and relative shocks. However, in presence of firms with identical elasticity of supply, realized aggregate shocks have no effect on *RPV*, because all firms respond identically to any given aggregate shock, while *ex ante* inflation uncertainty has a positive effect on *RPV*. The extension of the signal extraction model developed by Hercowitz (1981) and Cukierman (1983) assumes firms with different price elasticity of supply, which implies different responses of prices to unexpected aggregate demand shock. Thus, the higher unexpected inflation the higher *RPV*, i.e. the key factor is the size of the shock, while the sign of unexpected inflation is irrelevant.

On the other hand, although empirical evidence does not support unambiguously any of the above approaches, the positive relationship between *RPV* and inflation has practically become a stylized fact in economics. Traditional works as Vining and Elwertowski (1976) and Parks (1978) conclude that such relation is linear but there is increasing evidence in favour of a non-linear relationship. In turn, more recent findings show that this is both non-linear and unstable among different monetary policy and inflation regimes -see Caglayan and Filiztekin (2003) for Turkey and Caraballo et al. (2009) for Argentina, Brazil and Peru-. Moreover, recent research presents three types of evidence. Firstly, Nautz and Scharff (2005) for Germany, and Nautz and Scharff (2006) for the Euro area find that *RPV* is increasing in inflation even in low inflation environments. Secondly, Bick and Nautz (2008), in a panel threshold model for several USA cities verify positive and negative effects of inflation on *RPV*, while the suggested annual inflation to minimize *RPV* is in the rank of 1,8-2,8%. In this branch, and more important, Fielding and Mizen (2008) for USA and Choi (2009) for USA and Japan do show evidence of a U-shape profile of the inflation-*RPV* relationship. These findings have relevant implications for monetary policy. If such relation is linear, the lower the inflation, the lower the *RPV* and therefore the optimal inflation rate that minimizes the welfare costs of price dispersion is zero. But this reasoning is no longer true if the inflation-*RPV* relationship shows a U-shape, because in this case the inflation rate that minimizes *RPV* is positive, so that reducing inflation beyond it could be harmful.

In short, empirical results suggest a changing inflation-*RPV* relationship, and support the idea of non-neutrality, independently of the inflation environment. Thus, inflation should provoke distorting effects, because it impedes the efficient allocation of resources, even in low inflation economies, with the consequent welfare costs.

From that evidence, the goal of this paper is to analyze the features of the inflation-*RPV* relationship in Spain for the 1987-2009 period. The main motivation is to find new insights across the two monetary policy regimes experienced by the Spanish economy before and after its entry into the European Monetary Union (EMU). More precisely, we focus on the functional form of such relationship among such regimes. This analysis allows us to obtain the optimal inflation rate that minimizes *RPV*, as well as to determine the role of mechanisms underlying the effects of inflation on *RPV*, i.e. the role of inflation expectations and uncertainty.

Similarly to Fielding and Mizen (2008) and Choi (2009), we find that the inflation-*RPV* relationship for Spain takes a U-shape profile, which allows us to obtain a rank for the optimal inflation rate. *RPV* falls with inflation but rises back when inflation increases over a certain positive rate, i.e. inflation affects positively *RPV* only beyond certain inflationary threshold.

Furthermore, unlikely Nautz and Scharff (2005) for Germany, who find that *RPV* is mainly affected by expected inflation, our results indicate that unexpected inflation is the key link between inflation and *RPV*. On the contrary, neither expected inflation nor uncertainty affect price dispersion. Finally, from a policy perspective, our results show that disinflation benefits due to a lower inflation may not necessarily bring about welfare improvement if they are outweighed by the cost of increased price dispersion.

The paper is organized as follows. Section 2 presents a brief description of the data and variables used in this study. Section 3 contains the basic of econometric analysis of the relation between inflation and *RPV* across the two monetary regimes: before and after the entry of Spain into the EMU. The results suggest that such relation can be changing across those periods. In fact, we find a non-stable relation: this is positive and significant before the entry of Spain into the EMU, but it vanishes in the post-EMU period. Thus, in section 4 we check the stability of coefficients and carry out semiparametric estimations, which in turn allow us to seek for the optimal inflation rate. In order to determine the links between inflation and *RPV*, in section 5 we introduce inflation expectations and uncertainty. Finally, section 6 concludes.

2. Price data and variables

This study covers the period between January 1987 and September 2009 and employs monthly data for 57 categories of the consumer price index (CPI), extracted from the Instituto Nacional de Estadística. This disaggregation offers an advantage in order to calculate *RPV*, because a higher aggregation hides price variability, and then underestimates the true *RPV*.

The inflation rate is the monthly log-difference of the CPI. *RPV* is a measure of the non-uniformity of the variations of individual prices, relative to the average inflation rate. It is obtained as a modified version of the coefficient of variation (CV), using the weighted sum of individual prices inflation rate. At time *t*, *RPV* can be defined as follows:[1]

$$RPV_t = \frac{\left(\sum_i w_{it} (IN_{it} - IN_t)^2 \right)^{1/2}}{/1 + IN_t/} \tag{1}$$

where w_{it} is the weight of price *i* in the price index, IN_{it} the inflation rate of group *i* and IN_t the overall inflation rate at time *t*. Both series, *IN* and *RPV*, were deseasonalised by tramo-seats method.

On the other hand, as Elliott and Timmermann (2008) point out, univariate time series models seem to be appropriate to forecast inflation. Therefore, a univariate autoregressive moving-average model was chosen for mean inflation and we have specified a GARCH equation for the variance of the inflation model error term, which allows us to estimate a proxy for inflation uncertainty (*UN*). As inflation in Spain has experienced relevant changes, the parameters of the ARMA-GARCH model are not stable along the period, so that rolling equations were used to

[1] We consider that this is the best option to estimate *RPV* because it avoids two important problems. In first place, instead of the simple variance or standard deviation, it is not is spuriously correlated with the mean of the distribution – the inflation rate. Secondly, and more important in cases of low inflation like Spain, unlikely the traditional formula of CV, this alternative can be defined when inflation is close to zero or in periods of deflation. In fact, previous studies based on CV (e.g. Reinsdorf (1994) and Silver and Ioannidis (2001)), find a negative relationship, and these could be due to an "artefact" of the formula: *RPV* tend to infinite when inflation is near zero.

obtain expected inflation (*EIN*) and *UN*.[2] *EIN* is derived as the one-period-ahead inflation forecast and unexpected inflation (*UIN*) is the resulting forecast error: *UIN=IN-EIN*.

Moreover, we take into account the updating information process for CPI inflation in Spain. Following the standard modelization of inflation forecast, it was assumed that disposable information in t-1 to forecast inflation in period t is the actual inflation until t-2 and the expected inflation for t-1, given that the actual inflation for t-1 is known about the middle of period t.

EIN was obtained from a two-step procedure. In a first step, inflation has been modelled as an ARMA process using the standard Box-Jenkins methodology.[3] As usual, the standard Akaike information criterion has been applied to determine the optimal lag structure, from which an ARMA (1,6,12)(12) was selected as the best fitting ARMA model. Nonetheless, the forecast errors of this model were heteroskedastic, so that the inflation model could signal uncertainty. To estimate a proxy for *UN*, we have specified a GARCH equation for the variance of the inflation model error term. A GARCH (1,1) minimizes the Akaike criterion and by the simultaneous estimate of the ARMA process for the mean inflation and the GARCH equation, the following new inflation model with homocedastic forecast errors is obtained:

$$IN_t = a_1 IN_{t-1} + a_2 IN_{t-6} + a_3 IN_{t-12} + a_4 \varepsilon_{t-12} + \varepsilon_t \qquad (2)$$

$$\sigma_{\varepsilon t}^2 = b_1 \varepsilon_{t-1}^2 + b_2 \sigma_{\varepsilon,t-1}^2 \qquad (3)$$

where $\sigma_{\varepsilon t}^2$ is the inflation uncertainty.

Equations (2) and (3) were estimated using the Marquardt algorithm. As the residuals were not conditionally normally distributed, the covariance matrix and standard errors were computed by using the methods proposed by Bollerslev and Wooldridge (1992). The p-value of the z-statistic is shown into brackets.

$$IN_t = \underset{(0,01)}{0{,}06}\, IN_{t-1} + \underset{(0,00)}{0{,}13}\, IN_{t-6} + \underset{(0,00)}{0{,}75}\, IN_{t-12} - \underset{(0,00)}{0{,}35}\, \varepsilon_{t-12} + \varepsilon_t \qquad (2')$$

---

[2] In order to apply rolling equations to obtain *EIN* and *UN*, we have used data for monthly inflation from December 1979 to August 2009.

[3] As it is well known, the first step to model uncertainty with the variance of the errors terms of the inflation model is to test if inflation is stationary –see section 3.1.-; if this is not the case, the variance of errors explodes and it makes no sense to use such variance as a proxy of uncertainty.

$$\underset{(0,04)}{\sigma^2_{\varepsilon t} = 0{,}082}\,\varepsilon^2_{t-1} + \underset{(0,00)}{0{,}90}\,\sigma^2_{\varepsilon,t-1} \tag{3'}$$

Adjusted R^2 0,58.

In a second step and given the relevant changes experienced by inflation along the period, the parameters stability was analyzed by the recursive coefficients technique. Results show instability (see appendix 1). Thus, *EIN* was derived by means of this technique. From the December 1979-November 1986 period we estimated *EIN* for January 1987,[4] and for the rest of the period we estimate the following model from December 1979 until t to derive EIN_{t+2}, as follows:

$$IN_t = a_{1,t} IN_{t-1} + a_{2,t} IN_{t-6} + a_{3,t} IN_{t-12} + a_{4,t}\varepsilon_{t-12} + \varepsilon_t \tag{4}$$

$$\sigma^2_{\varepsilon t} = b_{1,t}\varepsilon^2_{t-1} + b_{2,t}\sigma^2_{\varepsilon,t-1} \tag{5}$$

3. Monetary regimes, inflation and RPV: basic regression analysis

3.1 Stationarity

Previous to the analysis of the inflation-*RPV* relationship, stationarity of inflation and *RPV* is checked by applying ADF y Phillips-Perron unit root tests to the seasonally adjusted series for the total period. The results are presented in Appendix 2, Table 1. Unit root is rejected for inflation, even though for ADF test only at 10%. On the contrary, results for *RPV* are ambiguous: ADF test accepts unit root, while Phillips-Perron result rejects it. Nonetheless, since ADF test has low power under the presence of a structural break, it may falsely detect a unit root, and then both, ADF and Phillips-Perron tests, could cast different results. Hence, we apply unit root tests proposed by Perron (1994) and Volgelsan and Perron (1994) which allow for a break in the series at an unknown time. The results are presented in Appendix 2, table 2. Both inflation and *RPV* present possible breaks from 1997.04 to 1998.05, and once again unit root is rejected only for inflation. To deal with this result lags of *RPV* are included in the OLS regressions.

In turn, in order to check the robustness of our results, in this section we employ the seasonally adjusted monthly core inflation (*CIN*), i.e., inflation obtained by excluding food and energy prices, from which the corresponding *RPV* has been calculated (*CRPV*). ADF and PP show different results for *CIN* and *CRPV*. Once we apply Perron (1994) and Volgelsan and Perron (1994) tests, a unit root with a break is accepted for both variables. The possible break appears between 1997.08 and 1999.01. Clearly, in all cases the breaks are detected around the entry of Spain into the EMU. This was associated to a change of monetary policy regime, and then to a different inflation behavior: before the entry the inflation rate slumped from an annual rate of 6,9% to 1,4% and after the entry it has been fluctuating between 4,2-0,8%.

3.2. Basic regression analysis

A first approach to the relation between inflation and *RPV* is obtained from OLS regression analysis. Taking into account the results of previous section, the total period was divided into two sub-periods, leaving out the months in which the variables present possible breaks: for *IN* and *RPV*, the first period spans from January 1987-March 1997, and the second from June 1998 to September 2009. For *CIN* and *CRPV* the first period is 1987.01-1997.08 and the second is 1999.02-2009.08. In turn, to capture the impact of inflation and deflation on price dispersion, *RPV* is regressed on absolute value of inflation (*AIN*) and *CRPV* on the absolute value of core inflation (*ACIN*).

The estimations include the number of lags of *AIN*, *ACIN*, *RPV* and *CRPV* that minimize the Akaike criterion. Thus, the resulting regression equations are:

$$RPV_t = \alpha + \beta_1 AIN_t + \sum_{h=1}^{12} \delta_h RPV_{t-h} + \varepsilon_t \quad (6)$$

$$CRPV_t = \alpha + \beta_1 ACIN_t + \sum_{i=1}^{2} \phi_i ACIN_{t-i} + \sum_{h=1}^{10} \delta_h CRPV_{t-h} + \varepsilon_t \quad (7)$$

Table 1 presents the results. They show that *AIN* and *ACIN* are positive and significant for the first period, the pre-EMU stage, while they are negative and not significant for the second period, the post-EMU stage.[5] These results can hide structural changes in the inflation-*RPV*

[4] In this sense, to consider the delays in updating information, the expected value for January 1987 is calculated with the actual value until November 1986 and the expected value for December 1986.

[5] The same conclusions are achieved when inflation and core inflation, instead of their absolute values, are taken as explanatory variables.

relationship. Since the parametric model seems to be too restrictive to capture a changing relation, in the next section we undertake a test of stability and a semiparametric approach.

TABLE 1: BASIC REGRESSION ANALYSIS

DEPENDENT VARIABLE: RPV_t				DEPENDENT VARIABLE: $CRPV_t$			
PERIOD	1987.01-2009.08	1987.01-1997.03	1998.06-2009.08	PERIOD	1987.01-2009.08	1987.01-1997.08	1999.02-2009.08
α	0,0007 (0,18)	-0,0001 (0,88)	0,001 (0,08)	α	0,0004 (0,27)	0,003 (0,00)	0,0005 (0,44)
AIN_t	0,07 (0,29)	0,20 (0,08)	-0,03 (0,65)	$ACIN_t$	0,11 (0,43)	0,47 (0,00)	-0,15 (0,49)
RPV_{t-1}	0,18 (0,00)	0,16 (0,03)	0,29 (0,00)	$CRPV_{t-1}$	0,86 (0,00)	0,77 (0,00)	0,90 (0,00)
R^2adj.	0,81	0,78	0,85	R^2adj.	0,93	0,56	0,95

t-statistics are based on standard errors computed according to Newey-West procedure to allow for residuals that exhibit both autocorrelation and heteroskedasticity of unknown form. Terms in brackets are the p-values associated to t-statistics. To simplify the presentation only the first lag of *RPV* appears in table.

4. Coefficients stability and non-linearities

In order to determine the true shape of the relation between inflation and *RPV*, in this section the stability of coefficients is checked, and then we try to approximate the true shape of this relation by means of a semiparametric analysis.

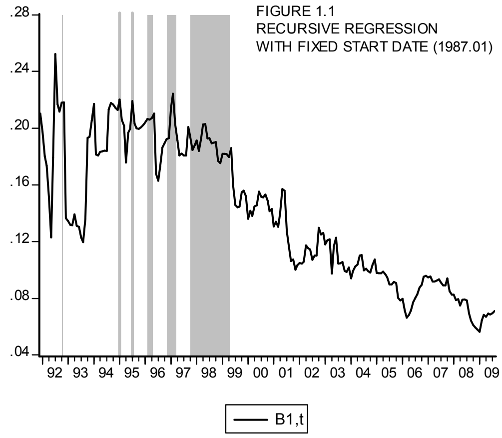
4.1 Coefficients stability

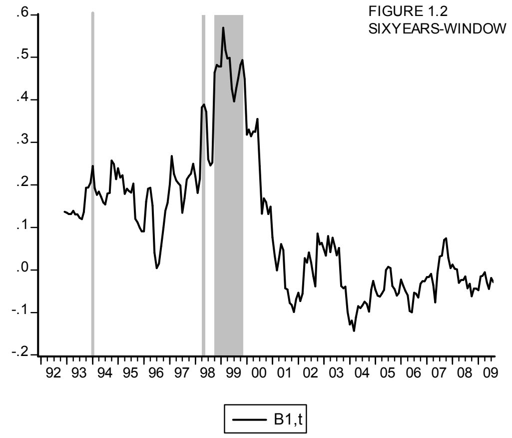
Additional precisions on previous evidence on a time-varying pattern of the inflation-*RPV* relationship were obtained by employing rolling regression equations, which allow to capture variations of the explanatory variables coefficients (in this case *AIN*) without imposing any prior on the timing of break points. Hence, it is flexible in detecting structural changes over time, by allowing for each rolling sample to have a completely different estimation. A parametric model is used where *RPV* is dependent variable, and the explanatory variables are inflation and the number of lags of *RPV* and inflation that minimize the Akaike criterion. Therefore, we estimate:

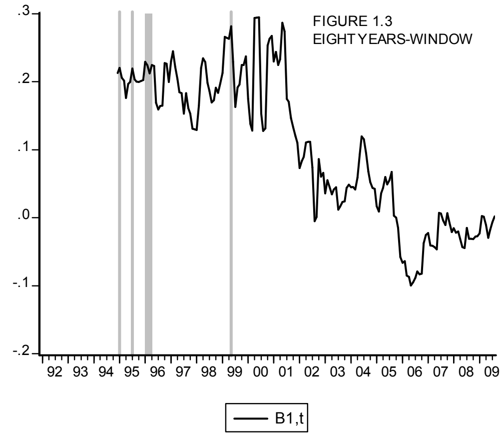
$$RPV_t = \alpha_t + \beta_{1,t} AIN_t + \sum_{h=1}^{12} \delta_{h,t} RPV_{t-h} + \varepsilon_t \quad (8)$$

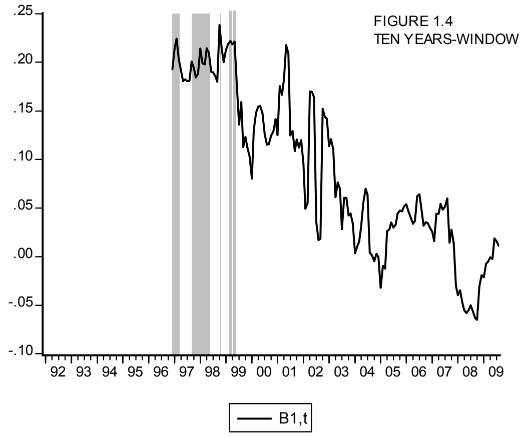
Thus, changes in the inflation-*RPV* relationship can be outlined by the parameters instability over rolling samples. Figures 1.1 to 1.5 present the results for $\beta_{1,t}$, our parameter of interest, obtained from different rolling regressions.

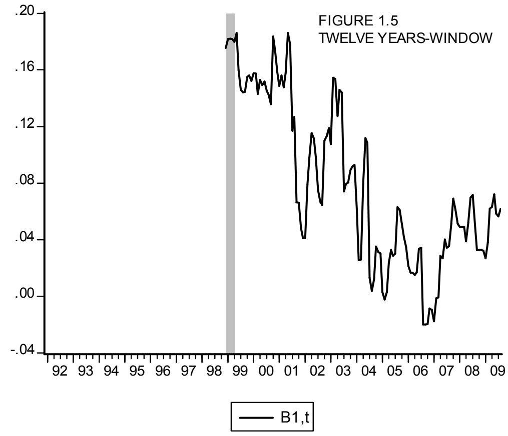
FIGURES 1.1 TO 1.5: ROLLING REGRESSIONS











Note: Figure 1.1 presents recursive coefficients, which were obtained by successive additions of one month to the 1987.01-1991.12 sub-sample. Figures 1.2 to 1.5 show the results for the 6, 8, 10 and 12-years windows, respectively (the results for windows of different extensions were very similar, so that these are omitted here). The significance of coefficients is for 10% of confidence intervals, and the months for which they are significant are marked in grey lines. The numbers on the horizontal axis represent the ending month of each window. For example, for a six-year window, the value of $\beta_{1,t}$ in 1992.12 captures the estimation of the parameter in (3) for 1987.01-1992.12, and so on.

As it can be seen from figures 1.1. to 1.5, in all cases $\beta_{1,t}$ is strongly unstable, and in special decreasing in time in the second half of the total period. In the case of recursive coefficients estimation, this result indicates a changing marginal impact of inflation on *RPV* when new months are incorporated in the estimation. In turn, the rolling regressions for fixed windows present a lower step of such coefficient for the post-EMU period, i.e. since 1998, approximately, and this result is robust for different size of the windows: 6, 8, 10 and 12 years.[6]

In short, previous results indicate an unstable relation between inflation and *RPV*. This varies significantly with the monetary policy regime. More precisely, coefficients are clearly sensitive to the addition of years from 1998, and they drop and lose significance in the post-EMU period. In this sense, the changing results across different samples suggest that parametric approach does not seem to be adequate. We try to overcome this problem in the next section.

4.2. Semiparametric approach and optimal inflation

In order to find out additional information on the shape of the inflation-*RPV* relation, as Fielding and Mizen (2008) and Choi (2009), we apply a partially linear model. To compare with our previous findings we have used the same number of lags for *RPV* and *IN* as in (6):

$$RPV_t = \sum_{h=1}^{12} \delta_h RPV_{t-12} + g(IN_t) + \varepsilon_t \quad (9)$$

where $g(IN_t)$ is an unknown smooth differential function that tries to capture the non-linear impact of inflation on *RPV* at time *t*. Therefore, our goal is the estimation of $g(IN_t)$ in (9).

The $g(IN_t)$ function has been estimated semi-parametrically in two stages. In the first one, the parameters λ_k are estimated from the regression equation:

$$RPV_t = \sum_{h=1}^{12} \lambda_h \overline{RPV}_{t-12} + \eta_t \quad (10)$$

where $\overline{RPV}_{t-h}$ are the residual series from a non-parametric regression of each lag of RPV_t on IN_t.

In the second stage, $g(IN_t)$ function is estimated non-parametrically from the regression:

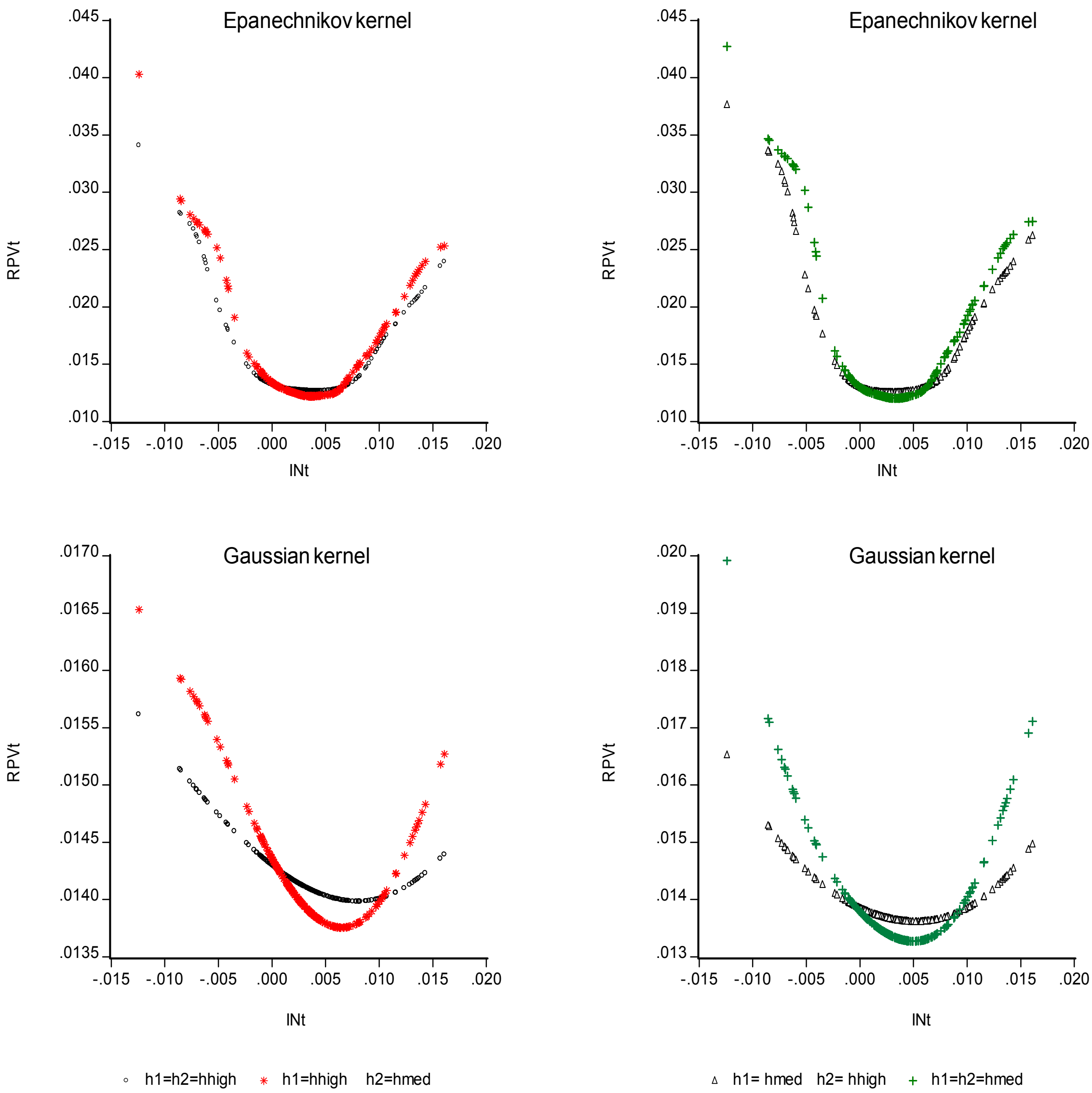
[6] Similar results for core inflation were obtained from rolling equations. Coefficient for core inflation starts to decline very sharply since 1998-1999, depending on the size of the window. In contrast to the inflation-*RPV* relationship, coefficient for *CIN* is strongly significant in the pre-EMU stage for all cases. These results are disposable from authors upon request.

$$\hat{\eta}_t = g(IN_t) + v_t \tag{11}$$

where $\hat{\eta}_t = RPV_t - \sum_{h=1}^{12} \lambda_h \overline{RPV}_{t-12}$

In both stages the non-parametrical regressions have been estimated using a Nadaraya-Watson kernel regression estimator. Given the data size, we have selected 150 points at which each local polynomial regression is evaluated. In order to know how the estimation of $g(IN_t)$ is affected by the treatment of extreme values of inflation, an unbounded Gaussian kernel and an outlier-robust Epanechnikov kernel were used. In turn, as the results of nonparametric regression can be sensitive to the bandwidth parameter (h), we have used three bandwidth based on automatic bandwidth selection method as follows: we have chosen h_{low}=0,5·h, h_{med}=h, h_{high}=1,5·h where h=0,15·(X_{max}-X_{min}) and (X_{max}-X_{min}) is the range of the explanatory variable in the non-parametric regression. Therefore, we have estimated eighteen $g(IN_t)$ functions in order to compare results. As neither the kernel function nor the bandwidth affect the shape of the function, only eight of them are plotted in figure 2.

FIGURE 2: $g(IN_t)$ FUNCTIONS



Note: *h1* and *h2* are the bandwith parameters in the first and second stage of the semiparametric estimation respectively. And in this case, the bandwiths used are h_{low}=0,00213, h_{med}=0,00427, h_{high}=0,006405

Clearly, this hints a non-linear inflation-*RPV* relationship, but a U-shape profile. *RPV* fall until certain threshold of low positive inflation, and rises back when inflation increases after passing this threshold. Thus, inflation affects positively *RPV* above a certain threshold inflation rate, but the relation seems to be negative below it, particularly for very low inflation rates and in deflation periods. This evidence is similar to the U-shape relation found by Choi (2009) for USA and Japan and by Fielding and Mizen (2008) for USA.

Table 2 shows the results for the parametric components of our semi-parametric model in (9). For purposes of shortness only λ_1 and λ_2 are presented. They are very similar to those obtained from equation (6) and from the quadratic model, i.e., when $g(IN_t)$ is defined as $g(IN_t)=\alpha+\beta_1 IN_t+ \beta_2 IN_t^2$ and equation (9) is estimated by OLS.

TABLE 2: PARAMETRIC COMPONENTS OF SEMIPARAMETRIC MODEL

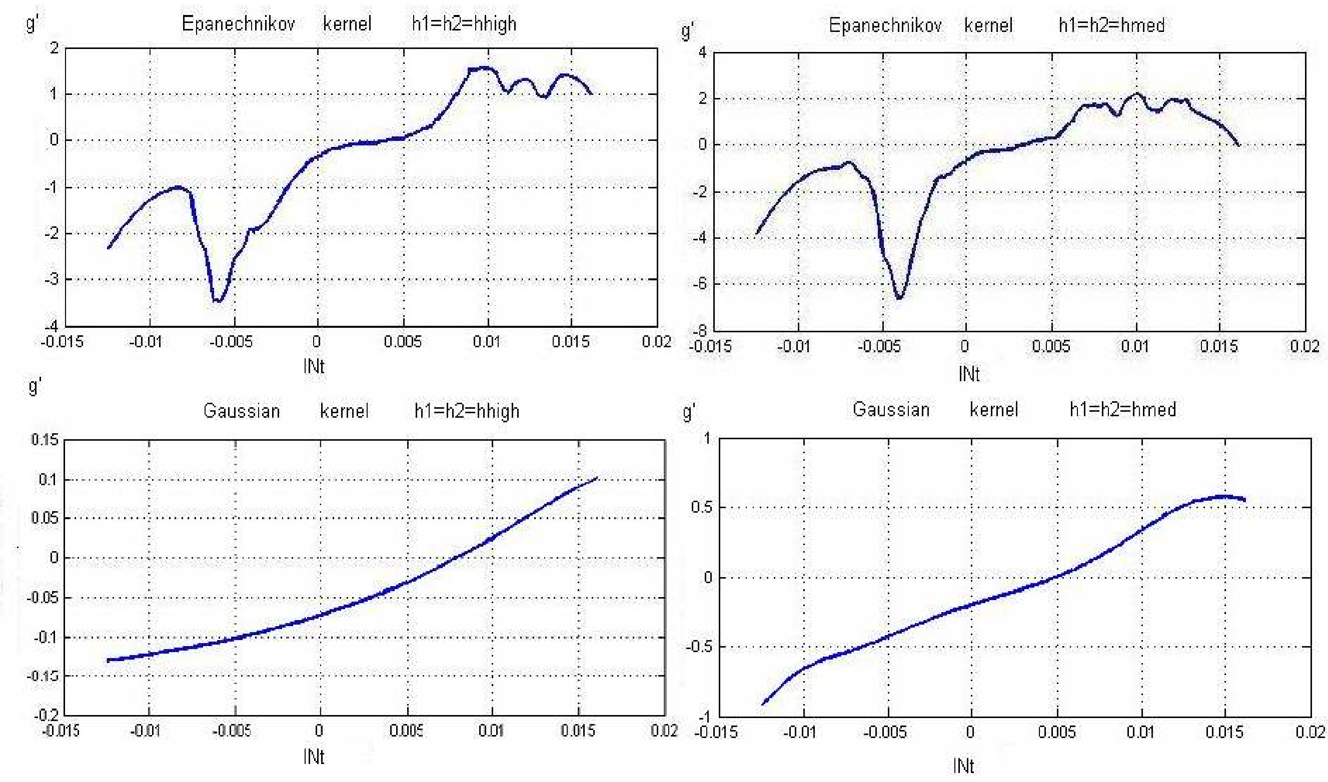
model	Kernel	Epanechnikov			Gaussian			equation (1)	quadratic model
	bandwidth	h_{low}	h_{med}	h_{high}	h_{low}	h_{med}	h_{high}		
RPV_{t-1}		0,19	0,19	0,18	0,18	0,17	0,17	0,18	0,18
RPV_{t-2}		0,14	0,17	0,17	0,17	0,16	0,16	0,16	0,16

On the other hand, the intuition behind the U-shape relationship between inflation and *RPV* found for Spain is the existence of an optimal inflation rate, i.e. one that minimizes *RPV*. Therefore, the next step is to achieve the derivative of the $g(IN_t)$ function, because it captures the sensitivity of *RPV* to marginal increase in inflation. If $g'(IN_t) >0$ ($g'(IN_t) < 0$), then *RPV* is increasing (decreasing) with inflation, while the inflation rate that minimizes *RPV* is given by $g'(IN_t) = 0$.

The derivative of $g(IN_t)$ was evaluated at different rates of inflation. Figure 3 presents the results for Epanechnikov and Gaussian kernel and different bandwidths.[7] They indicate a positive optimal inflation. The rank for the optimal annual inflation -obtained from the monthly inflation rate shown in figure 3- varies depending on the kernel function used: for the Epanechnikov kernel, the optimal inflation rate is in a rank of 2,8-4,8%, while for the Gaussian kernel this rank is 3,9-9,5%. Such difference can be due to the fact that the Epanechnikov kernel is outlier-robust, while the Gaussian kernel is not robust to outliers. In order to compare these results with additional information on optimal inflation values, two different outliers-robust kernel were applied: cosine and biweight. For the former the optimal inflation rank is 2,75-4,65% and for the latter is 3,21-4,44%. Hence, similarly to previous findings, like Fielding and Mizen (2008) for USA, the optimal annual inflation rate for Spain is around 4%. In turn, in terms of monetary policy, the U-shaped profile found for Spain show that in lower inflation periods with inflation rates under such threshold, disinflation efforts are not successful to reduce *RPV* and improve welfare.

[7] Even tough a total of eighteen graphs of derivatives have been obtained, for simplicity we include only four. However, all of them are disposable from authors upon request.

FIGURE 3: DERIVATIVES AND OPTIMAL INFLATION RATE



Note: *h1* and *h2* are the bandwith parameters in the first and second stage of the semiparametric estimation respectively. And in this case, the bandwiths used are h_{low}=0,00213, h_{med}=0,00427, h_{high}=0,006405

In sum, the evidence shown in sections 4.1 and 4.2 hints a changing and non-linear relation between inflation and *RPV*. Moreover, this is significant only for period previous to the entry of Spain into the euro zone. This result favours the hypothesis that monetary regimes matters. In particular, expected and unexpected inflation and uncertainty can be the underlying causes behind this kind of relation (Caraballo et al. (2006), Caraballo and Dabús (2008), Caglayan et al. (2008), Becker and Nautz (2009b), Choi (2009)). This issue is studied in the next section.

5. Inflation expectations and uncertainty

In order to find out the links between *IN* and *RPV*, this section introduces the components of inflation: expected and unexpected inflation and uncertainty. *EIN* series obtained in section 2 shows a seasonal component which has been removed using the tramo-seat method. *UIN* is the difference between seasonally adjusted *IN* and seasonally adjusted *EIN*. *UN* does not present a seasonal component.

To capture the V-shaped relationship between inflation and *EIN* predicted by menu cost we should take the absolute value of *EIN*, but for our data *EIN* is always positive. In turn, we distinguish between positive unexpected inflation (UIN^+) and the absolute value of negative unexpected inflation ($AUIN^-$) to test the implications of the extended signal extraction model. The lags of *RPV* that minimizes Akaike criterion are included, but lags for *EIN*, UIN^+, $AUIN^-$ and *UN* were not considered because of the speed of information publications: CPI is published with monthly periodicity. Finally, the following equation is estimated for the total period and the pre-EMU and post-EMU periods shown in section 3 for the inflation-*RPV* relationship:

$$RPV_t = \alpha + \beta_0 EIN_t + \beta_2 UIN_t^+ + \beta_3 AUIN_t^- + \beta_4 UN_t + \sum_{k=1}^{12} \lambda_k RPV_{t-k} + \varepsilon_t \qquad (12)$$

Table 3 presents the results. Recall that menu costs model predicts $\beta_0 > 0$, $\beta_2 = \beta_3 > 0$ can be considered as evidence in favour of the extended signal extraction model, and $\beta_4 > 0$ is supported by signal extraction model. Wald test is used to check if $\beta_2 = \beta_3 > 0$. χ^2 statistic is reported given that the variances have been estimated using the Newey-West method, and therefore the F-statistics does not possess the desired finite-sample properties

TABLE 3: *RPV*, EXPECTED AND UNEXPECTED INFLATION AND UNCERTAINTY

	1987.01-2009.08	1987.01-1997.03	1998.06-2009.08
α	0,00 (0,29)	-0,00 (0,44)	0,00 (0,60)
EIN_t	-0,002 (0,99)	1,06 (0,15)	0,32 (0,50)
UIN_t^+	0,31 (0,00)	0,74 (0,00)	0,06 (0,61)
$AUIN_t^-$	0,24 (0,02)	0,53 (0,08)	0,17 (0,05)
UN	0,05 (0,81)	0,54 (0,44)	0,03 (0,85)
RPV_{t-1}	0,18 (0,00)	0,20 (0,00)	0,27 (0,00)
R^2adj.	0,81	0,78	0,85
Wald test: H_o: $\beta_2 = \beta_3$ $\chi^2(1)$ statistics, p-values into brackets	0,36 (0,54)	0,61 (0,43)	0,87 (0,35)

Note: t-statistics are based on standard errors computed according to Newey-West procedure to allow for residuals that exhibit both autocorrelation and heteroskedasticity of unknown form. The terms in brackets are the p-values associated to t-statistics. To simplify the presentation only the first lag of *RPV* appears in the table.

As it can be concluded from table 3, *EIN* and *UN* are not significant. This result holds when rolling equations techniques are applied: none of them are significant independently of the size of the window or the sample size. Hence, there is no evidence in favour of menu cost or signal extraction models. As far as unexpected inflation is concerned, both UIN^+ and $AUIN^-$ are significant for the whole period. However, for the pre-EMU period $AUIN^-$ is significant just at 10%, and for the post-EMU period UIN^+ is not significant, but as Wald test fails to reject the null of $\beta_2 = \beta_3$ there is evidence in favour of extended signal extraction model. Rolling equations show that β_2 and β_3 decline along the period and they show weak sensitivity to sample or window size.[8] The latter result is in line with those presented in section 3, given that if unexpected inflation explains the relation between inflation and *RPV*, and this relation becomes weaker along the period, coefficients of unexpected inflation are supposed to decline as well.[9]

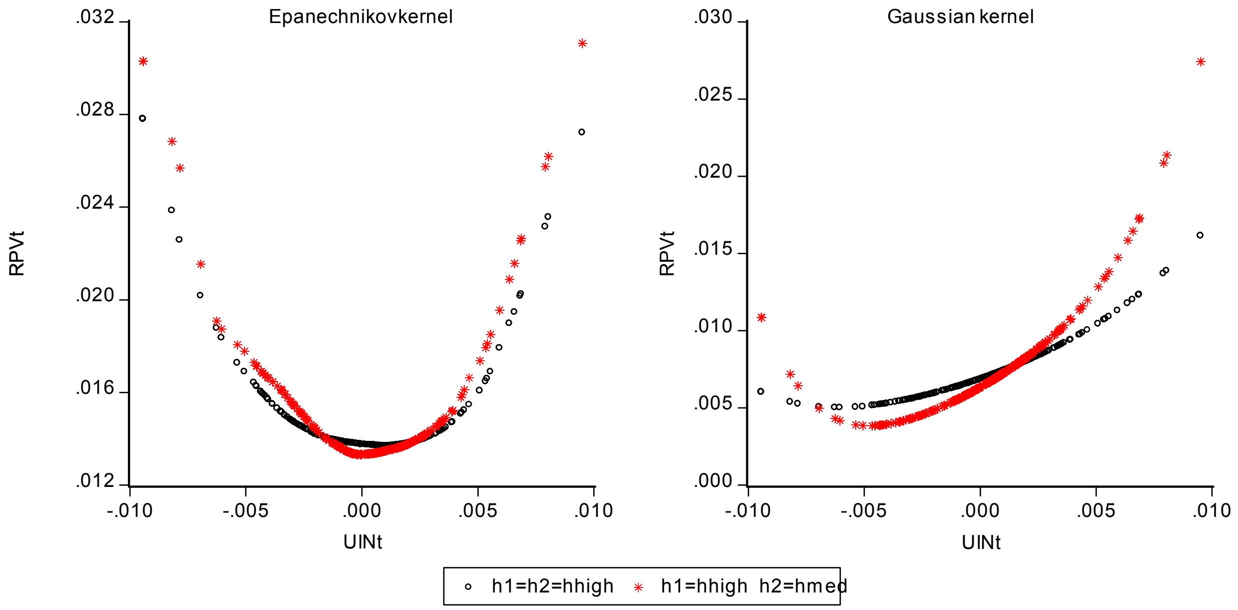
Finally, we try to find the true shape of the relationship between inflation and unexpected inflation by means of semiparametric approach as in section 4. The variables included in the parametric part are the lags of *RPV*, *EIN* and *UN*, while $g(UIN_t)$ tries to capture a non-linear relation between *RPV* and unexpected inflation. Therefore the following equation is estimated:

$$RPV_t = \alpha + \beta_0 EIN_t + \beta_4 UN_t + g(UIN_t) + \sum_{k=1}^{12} \lambda_k RPV_{t-k} + \varepsilon_t \qquad (13)$$

Results for $g(UIN_t)$ appear in figure 4. Once again the kernel function affects the results. The Epanechnikov kernel shows a clear U-shape for the relation between *UIN* and *RPV*, but it is not so evident for the Gaussian kernel. According to the former method, the unexpected inflation that minimizes *RPV* is around zero. Nonetheless, for the Gaussian kernel *RPV* is minimized at negative values of *UIN*. As it was mentioned above, it can be due to the different treatment of the outliers in each kernel. In fact, results for cosine and biweight kernels are very similar to those obtained with the Epanechnikov kernel: UIN_t around zero minimizes RPV_t.

[8] Results of rolling equations are available from authors upon request.
[9] In this sense, recent works also present evidence of a changing role for inflation expectations. Nautz and Scharff (2005, 2006) and Becker and Nautz (2009b) find that the impact of expected inflation on *RPV* is strongly declining in lower inflation periods, because inflation expectations had been stabilized on a low level.

FIGURE 4: *g(UIN$_t$)* FUNCTIONS



Note: h1 and h2 are the bandwidth parameters in the first and second stage of the estimation respectively. And in this case the bandwidths used are h_{low}=0,00142, h_{med}=0,00284, h_{high}=0,00426

In sum, if unexpected inflation is near zero, i.e. if there is no difference between actual and expected inflation, then welfare costs derived from price dispersion are minimal. From a monetary policy perspective, this means that credibility and fulfillment of announcements regarding inflation matter. Only a credible and predictable monetary policy could reach the goal of minimizing the welfare costs caused by the distorting impact of inflation on *RPV*.

6. Conclusions

This paper analyzes the relation between inflation and *RPV* for Spain and the implications for monetary policy that can be derived from it. Our results suggest that monetary regimes matter. More precisely, the inflation-*RPV* relationship changes before and after the entry of Spain into the EMU. Similarly to previous papers for USA and Japan, we find a changing and U-shape inflation-*RPV* relationship, which in turn allowed us to determine an optimal inflation rate in this framework, i.e. the inflation that minimizes *RPV*. Even though results differ depending on the kernel functions used, outlier robust kernels yield a rank of 3-4% for optimal annual inflation rate, higher than the inflation target proposed by the European Central Bank.

Moreover, the key link underlying the relation between inflation and *RPV* appears to be unexpected inflation, which is significant for the total period and for the first period before the entry of Spain into the EMU while only negative unexpected inflation is significant in the second period. These results are compatible with extended signal extraction model predictions. Besides, in order to minimize *RPV*, the evidence indicates an optimal value of unexpected inflation near zero. These results have clear implications for the monetary policy: the welfare costs of inflation caused by the distorting impact of inflation on *RPV* can be avoided with a credible and predictable inflation targeting policy.

Finally, further steps in this line of research might include a comparative analysis among economies with different inflation regimes. Natural candidates are economies with experience of high inflation. A wider sample of countries should help us to determine if the U-shape functional relation between inflation and RPV remain among different inflationary environments.

REFERENCES

Barro, R. J. (1976): Rational expectations and the role of monetary policy. *Journal of Monetary Economics*, 2, 1–32.

Bollerslev, T. and Wooldridge, J. (1992): Quasi-maximum likelihood estimation and inference in dynamic models with time-varying covariances. *Econometric Reviews*, 11, 143-172.

Becker, S. and Nautz, D. (2009a): Inflation, Price Dispersion, and Monetary Search: Evidence from the European Union. *Department of Money and Macroeconomics Discussion Paper*. Goethe University Frankfurt.

Becker, S. and Nautz, D. (2009b): Inflation and Relative Price Variability: New Evidence for the United States. *Southern Economic Journal*, 76(1), 146–164.

Bick, A. and Nautz, D. (2008): Inflation threshold and relative price variability: Evidence from U.S. cities. *International Journal of Central Banking*, 4, 61–76.

Caglayan, M. and Filiztekin, A. (2003): Nonlinear Impact of Inflation on Relative Price Variability. *Economics Letters*, 79 (2), 213-218.

Caglayan, M.; Filiztekin, A. and Rauh, M.T. (2008): Inflation, price dispersion, and market structure. *European Economic Review*, 52, 1187–1208.

Caplin, A. and Leahy,J. (1991): State Dependent Pricing and the Dynamics of Money and Output. *Quarterly Journal of Economics*, 106, 683-708.

Caplin, A. and Spulber, D. (1987): Menu Costs and the Neutrality of Money. *Quarterly Journal of Economics*, 102, 703-725.

Cukierman, A. (1983): Relative Price Variability and Inflation: a Survey and Further Results, in K. Brunner and A.H. Meltzer (eds.) Variability in Employment, Prices and Money, Elsevier, Amsterdam.

Caraballo, M.A. and Dabús, C. (2008): The Determinants of Relative Price Variability: further Evidence From Argentina. *Cuadernos de Economía - Latin American Journal of Economics*, 45, 235-255.

Caraballo, M.A.; Dabús, C. and Usabiaga, C. (2006): Relative Prices and Inflation: New Evidence from Different Inflationary Contexts. *Applied Economics*, 38, 1931-44.

Caraballo, M. A.; Dabús, C. and Caramuta, D. (2009): Price Behaviour at High Inflation, in *Inflation: Causes and Effects, Control and Prediction*. Nova Science Publishers, Leon V. Schwartz (editor) (2009), Nueva York.

Choi, C. (2009): Relationship between Inflation and Relative Price Variability: Revisited. *University of Texas at Arlington Working Paper*. Available at SSRN: http://ssrn.com/abstract=1434147

Elliott, G. and Timmermann, A. (2008): Economic forecasting. *Journal of Economic Literature*, 46, 3–56.

Fielding, D. and Mizen, P. (2008): Evidence on the functional relationship between relative price variability and inflation with implications for monetary policy, *Economica*, 75, 683-699.

Hercowitz, Z. (1981): Money and the dispersion of relative prices. *Journal of Political Economy*, 89, 328–56.

Lucas, R.E. (1973): Some international evidence on output-inflation tradeoffs. *American Economic Review*, 63, 326–35.

MacKinnon, J. G. (1996): Numerical Distribution Functions for Unit Root and Cointegration Tests. *Journal of Applied Econometrics*, 11, 601–618.

Nautz, D. and Scharff, J. (2005): Inflation and relative price variability in a low inflation country: Empirical evidence from Germany. *German Economic Review*, 6, 507–23.

Nautz, D. and Scharff, J. (2006): Inflation and relative price variability in the Euro-area: Evidence from a panel threshold model. *Deutsche Bundesbank Discussion Paper Series* 1. No. 14/2006.

Parks, R. W. (1978): Inflation and relative price variability. *Journal of Political Economy*, 86, 79–95.

Perron, P. (1994): Further Evidence on Breaking Trend Functions in Macroeconomics Variables, *Working Paper 2594,* CRDE, Université de Montréal (Canadá).

Reinsdorf, M. (1994): New evidence on the relation between inflation and price dispersion. *American Economic Review*, 84, 720–31.

Rotemberg, J. (1983): Aggregate consequences of fixed costs of price adjustment. *American Economic Review*, 73, 433–63.

Sheshinski, E. and Weiss, Y. (1977): Inflation and costs of price adjustment. *Review of Economic Studies*, 44, 287–303.

Silver, M. and Ioannidis, C. (2001): Intercountry differences in the relationship between relative price variability and average prices. *Journal of Political Economy*, 109, 355–74.

Vining, D.R. and Elwertowski, T.C. (1976): The Relationship between Relative Prices and General Price Level. *American Economic Review*, 66 (4), 699-708.

Vogelsang, T. J. y Perron, P. (1994): «Additional Tests for a Unit Root Allowing for a Break in the Trend Function at an Unknown Time», *Working Paper 9422* CRDE, Université de Montréal (Canadá).

Appendix 1: STABILITY OF COEFFICIENTS- ARMA MODEL

This appendix shows the results of the recursive coefficients estimation of the ARMA-GARCH model presented in equations (4) and (5) – see section 2-:

$$IN_t = a_{1,t} IN_{t-1} + a_{2,t} IN_{t-6} + a_{3,t} IN_{t-12} + a_{4,t} \varepsilon_{t-12} + \varepsilon_t$$

$$\sigma^2_{\varepsilon t} = b_{1,t} \varepsilon^2_{t-1} + b_{2,t} \sigma^2_{\varepsilon,t-1}$$

FIGURE I

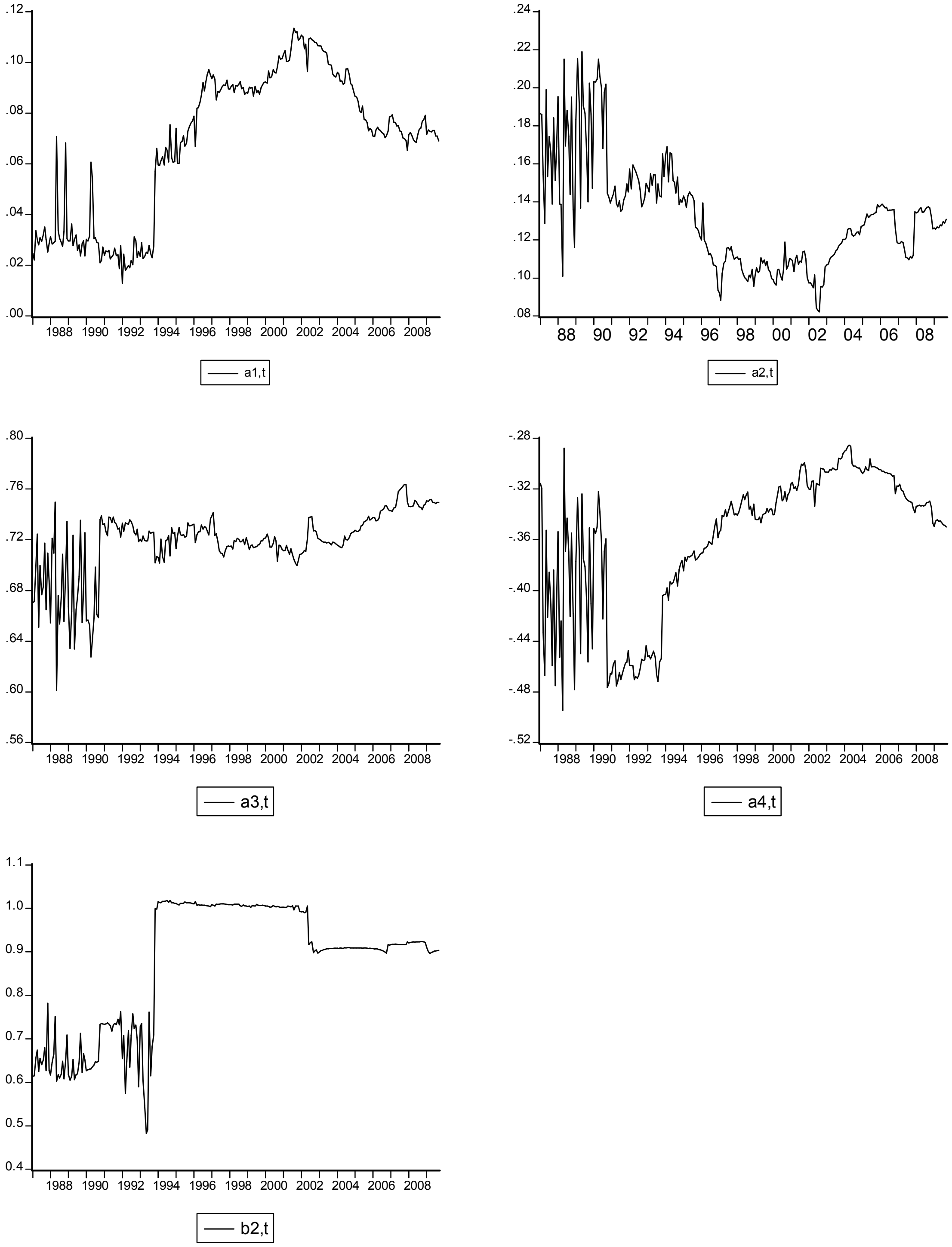


Figure I presents recursive coefficients estimated using the Marquardt algorithm and computing the covariance matrix and standard errors with Bollerslev and Wooldridge (1992) method. Coefficients have been obtained by successive additions of one month to the 1979.12-1986.11 sub-sample. The numbers on the horizontal axis represent the ending month of each estimation.

Appendix 2: UNIT ROOTS TEST

Table I presents the results of ADF and Phillips-Perron unit root tests for seasonally adjusted variables

TABLE I: ADF AND PHILLIPS-PERRON UNIT ROOT TESTS (1987.01-2009.08)

	ADF test					**PP test**			
Variable	Number of lags-criterion		Constant	Trend	t- statistic*	Bandwidth	Constant	Trend	Adj. t-statistics*
IN_t	Akaike	14	Yes	Yes	-3,27 (0,07)	6	Yes	Yes	-11,039 (0,00)
	Schwarz	0	Yes	No	-6,45 (0,00)				
CIN_t	Akaike	13	Yes	Yes	-3,09 (0,10)	3	Yes	Yes	-6,68 (0,00)
	Schwarz	13	Yes	Yes	-3,09 (0,10)				
RPV_t	Akaike	12	No	No	-0,54 (0,48)	6	Yes	No	-4,60 (0,00)
	Schwarz	4	No	No	1,23 (0,94)				
$CRPV_t$	Akaike	15	No	No	1,10 (0,92)	13	Yes	Yes	-3,45 (0,04)
	Schwarz	3	No	Yes	-2,31 (0,42)				

Note: A Bartlett kernel-based estimator of the frequency zero spectrum is used for the Phillips Perron test
*MacKinnon (1996) one-sided p-values into brackets.

On the other hand, we check unit root test with structural breaks by applying the tests proposed by Vogelsang and Perron (VP test from now on). These allow us to distinguish two key properties: 1) if the break affects the constant, the trend or both of them in the series, and 2) if the rupture impact on the variable immediately (additional outlier) or gradually (innovational outlier). Taking into account the evolution of inflation and *RPV* series, we consider that additional outlier model must fit better to check structural breaks and unit root, because of the entry into the Euro affects inflation and *RPV* once and for all. In second place, we select two models, one includes breaks in constant and trend, and the other one considers changes only in trend.

Following VP (1994), testing for a unit root test in the additional outlier framework includes two steps. In the first one, the following equation is estimated:

$$y_t = u + \beta t + v^i DU_t + g^i DT_t + \varepsilon_t \qquad (1)$$

where y_t is the variable under study (in our case inflation and *RPV*), u is a constant, t is the trend, and DU_t *and* DT_t are dummies for the constant and the trend respectively. Three models can be distinguished: 1) if i=A the break only affects the constant, and $g=0$, 2) i=C indicates rupture in trend, and then $v=0$, and 3) i=B corresponds to the case that the rupture is in both constant and trend. In turn, calling T_B the breakpoint, $DU_t=1$ and $DT_t=t-T_B$ if $t>T_B$, and zero otherwise.

In a second stage, and from the residuals of the regression of equation (1), we estimate by OLS (2) if i=A, B, and (3) if i=C.

$$\varepsilon_t = \alpha \ \varepsilon^i{}_{t-1} + \sum_{j=0}^{k} DTB_{t-j} + \sum_{j=1}^{k} \Delta\varepsilon^i_{t-j} + u_t \quad (2)$$

$$\varepsilon_t = \alpha\varepsilon^i_{t-1} + \sum_{j=1}^{k} \Delta\varepsilon^i_{t-j} + u_t \quad (3)$$

where $DTB=1$ for $t=T_{B+1}$ and 0 otherwise.

Following VP (1994), two data dependent methods can be applied to detect the breakpoints. The first one selects T_B that minimizes t_α (t-statistics corresponding to the estimated α in equations (2) and (3)). In this case, the choice of T_B corresponds to the break date which is most likely to reject the unit root hypothesis. The second method can be used for model A and B. In this case we pay attention to t_v and t_g (t-statistics associated to v (model A) or g (model B) in equation (1)). We choose the breakpoint that maximizes (minimizes) the t-statistics when the direction of the break is known a priori to be positive (negative) or the absolute value of the t-statistics when the direction of the break is unknown. Once T_B is determined, the corresponding t_α in equations (2) allows us to accept or reject unit root.

On the other hand, to choose k in (2) and (3) we apply two criteria. The first one consists in choosing a fix value for k, we have considered k=5 (as in VP (1994)).The second one is based on selecting a value of k ($k=k^*$) in such a way that in regressions (2) and (3) coefficient corresponding to k^* is significant, while it is not significant for $k>k^*$.

Results of applying the above methodology to *IN*, *RPV*, *CIN* and *CRPV* are presented in tables II-V. These series show a change in trend along the period, therefore in the paper we have taken into account results obtained by model C. Nevertheless, we have also included results for model B. Trimming is slightly different in each case but in all of them the first twelve months and the last twenty four months have been removed.

BREAK IN TREND-MODEL C

TABLE II

		IN		*CIN*		*RPV*		*CRPV*	
		k=5	k(t-sig)	k=5	k(t-sig)	k=5	k(t-sig)	k=5	k(t-sig)
Min t_α	T_B	1997:07	1997:07	1997:08	1998:02	1997:08	1997:06	1999:01	1998:09
	t_α	-4,748**	-4,07*	-3,052	-4,269	-3,708	-3,630	-3,643	-2,895

Note: t_α-critical values values in Perron(1994)
*,**,*** indicate significance at the 10%, 5% and 1% levels respectively

As it can be seen from table II, only for *IN* unit root is rejected

TABLE III

		IN	*CIN*	*RPV*	*CRPV*
Max t_g	T_B	1997:05	1998:05	1998:05	1998:02
	t_g	1,84	3,270	18.949	10.451

Note: obviously for model C, max t_g give some information about T_B but it can't be used to test unit root

BREAK IN CONSTANT AND TREND-MODEL B

TABLE IV

		IN		*CIN*		*RPV*		*CRPV*	
		k=5	k(t-sig)	k=5	k(t-sig)	k=5	k(t-sig)	k=5	k(t-sig)
Min t_α	T_B	1999:03	1997:08	1994:09	1995:03	1996:07	2000.12	2002.01	2002.01
	t_α	-5.084**	-4,82*	-3.493	-4,82**	-3,908	-4,005	-6.396***	-5,008*

Note: t_α-critical values values in Vogelsang and Perron (1994)
*,**,*** indicate signicance at the 10%, 5% and 1% levels respectively

TABLE V

			IN	*CIN*	*RPV*	*CRPV*
Max t_g	T_B		1997:07	1998:06	1998:02	1997:06
	t_g		1.859	3.283	18,783	11.742
	t_α	k=5	-4,771	-3,009	-3,678	-3,643
		k(t-sig)	-4.698	-4.269**	-3,556	-2,699

Note: t_α-critical values values in Perron(1994)
*,**,*** indicate signicance at the 10%, 5% and 1% levels respectively

With model B, results are not conclusive with respect to *IN*, *CIN* and *CRPV*. Only for *RPV*, unit root can't be rejected in all cases