OPTIMAL INFLATION, PRICE DISPERSION AND INFLATION EXPECTATIONS: THE SPANISH CASE

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This paper studies the relation between inflation and relative price variability (RPV) in Spain during the 1987-2009 period. We find that this relation presents a U-shape profile, and that the optimal annual inflation rate (defined as the one that minimizes RPV) is around 4%, higher than the 2% inflation target proposed by the European Monetary Union. More importantly, this result does not depend on whether the monetary regime is before or after the euro. Hence, the main policy implication is that disinflation efforts to achieve the 2% inflation target result in welfare losses. The key link between inflation and RPV is unexpected inflation, whose optimal level is around zero. This suggests that monetary policy matters: the welfare costs associated with higher RPV can be minimized with a credible and predictable inflation targeting policy set at the appropriate level.

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I. Introduction

This paper provides new evidence on the key features of the relationship between inflation and relative price variability (RPV), focusing on Spain during the 1987-2009 period. We additionally try to determine if the inflation target proposed by the European Monetary Union (EMU) is a good guideline for monetary policy. Unlike previous studies, we consider all the features of the inflation-RPV relationship for two different monetary regimes, the first of higher inflation from the beginning of the period to the entry of Spain in the EMU in 1999, and the second sub-period of lower inflation thereafter. This allows us to investigate not only the functional form but also the stability of the relationship. Secondly, we estimate the optimal inflation rate, which depends heavily on the shape of the inflation-RPV relationship: if such a relation is linear, then the lower the inflation, the lower the RPV, so 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

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rate that minimizes RPV could be positive.[1] In the third place, we study the role of inflation expectations and uncertainty as the linkages between inflation and RPV.

Our central findings are that the relation between inflation and RPV presents a U-shape profile and, more importantly, the annual optimal inflation rate is around 4%, which is higher than the EMU's 2% inflation target. This result is robust: it holds for different time periods. The main policy implication is that disinflation efforts to achieve the 2% target level result in welfare losses. The key link between inflation and price dispersion is unexpected inflation, whose its optimal level is around zero. Hence, our results suggest that monetary policy matters: the welfare costs associated with higher price dispersion can be minimized with a credible and predictable inflation targeting policy set at the appropriate level.

A huge body of empirical evidence has found a positive inflation-RPV relationship. Whilst traditional works like Glejser (1965) and Parks (1978) show that such a relationship is linear, more recently Caglayan and Filiztekin (2003) for Turkey and Caraballo et al. (2009) for Argentina, Brazil and Peru show that it is both non-linear and unstable across different inflationary regimes. Similarly, Fielding and Mizen (2008) for the USA and Choi (2010) for the USA and Japan find a U-shape functional profile, as well as a positive optimal inflation rate. Meanwhile, Bick and Nautz (2008), in a panel threshold model for several US cities, point out that the annual optimal inflation, i.e., the inflation that minimizes RPV, is in the range of 1.8%-2.8%. This result has important monetary policy implications: if the optimal inflation rate is positive, reducing inflation below it should increase RPV, and the welfare costs associated with higher price dispersion. Finally, Nautz and Scharff (2005) for Germany, and Nautz and Scharff (2006) for the Euro-area find that RPV is increasing with inflation, even in a low inflation environment. Several theoretical approaches try to disclose the links underlying such a relation: 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.[2]

In sum, the empirical results suggest a changing inflation-RPV relationship and support the idea of non-neutrality, regardless of the inflation environment. Thus, inflation increases price dispersion, and thus welfare costs. Nevertheless, the literature has focused only on some of the key features of that relationship. For instance, Choi (2010) studies the shape and the stability across different inflation regimes for the USA, while Choi et al. (2011) extend the study to a wider sample with inflation targeting countries. However, these papers do not

[1] In the literature focused on the inflation-RPV relationship, where this paper is included, the optimal inflation rate is defined as the one that minimizes RPV. Obviously, the term "optimal inflation" could refer to many other different meanings –such as the inflation rate related to the optimal monetary rule proposed by Friedman (1969), or the optimal output-inflation trade-off assessed in Tobin (1972) and Lucas (1973), among others– but they are not going to be taken into account in this paper.

[2] For additional explanations see Sheshinski and Weiss (1977), Rotemberg (1983), Caplin and Spulber (1987) and Caplin and Leahy (1991). In turn, early developments of the signal-extraction model can be found in Lucas (1973) and Barro (1976), Hercowitz (1981) and Cukierman (1983).

consider the crucial role played by the channels that connect inflation and RPV, in particular uncertainty and inflation expectations in different monetary and inflationary regimes.

In order to explore the key features of the inflation-RPV relationship for the case of Spain, the paper is organized as follows. In Section II we describe the price data and variables. Section III contains an Ordinary Least Square (OLS) estimation of the inflation-RPV relationship for the full sample and for the two sub-samples, before and after the entry of Spain into the EMU. This is not stable across both periods, which suggests a non-linear relation between inflation and RPV. Thus, in Section IV we check the stability of coefficients and carry out semiparametric estimations, which allow us to obtain the optimal inflation rate. Section V studies the role of inflation expectations and uncertainty. Finally, Section VI concludes.

II. Price Data and Variables

In this study we employ monthly price data of the Spanish Consumer Price Index (CPI), disaggregated into 57 categories, over the 1987.01-2009.09 period.[3] Data were extracted from the Instituto Nacional de Estadística (National Institute of Statistics). For each category and for the average, the inflation rate is the monthly log-difference of the CPI.

RPV is a measure of the non-uniformity: it captures the inflation rate of individual prices in relation to the average inflation rate. In order to avoid spurious correlation between the mean (the average inflation) and the variance (in this case RPV) we use a modified version of the coefficient of variation (CV), as follows:

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

where w_{it} is the weight of price i in the price index, IN_{it} the inflation rate of the price i, and IN_t the average inflation rate at time t.

We consider that expression RPV in equation (1) is the best option to define RPV because it avoids two important problems. On the one hand, instead of the simple variance or standard deviation, it is not spuriously correlated with the mean of the distribution, i.e., the inflation rate. On the other hand, and more important in low inflation economies, for inflation rates near zero, the traditional formula of CV generates values of RPV tending

[3] The 57 categories appear in Appendix A.

to infinite, which implies an “artificial” negative inflation-RPV relationship. Both series, *IN* and *RPV*, were deseasonalized by the TRAMO-SEATS method.

Since we study in Section V the role of inflation expectations and uncertainty in explaining the relation between inflation and RPV, we need to decompose inflation (*IN*) into its components: expected inflation (*EIN*), unexpected inflation (*UIN*) and uncertainty (*UN*). In order to do this, as Elliott and Timmermann (2008) point out, univariate time series models seem to be appropriate to forecast inflation, especially when data are monthly. Thus, we have chosen a univariate autoregressive moving-average model for the 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 *UN*.[4] Since the dynamics of inflation has changed during the period, to obtain the estimations of *EIN* and *UN* the parameters of the ARMA-GARCH model have been estimated by means of recursive regression in which we used monthly inflation data from December 1979 to August 2009. *EIN* is derived as the one-period-ahead inflation forecast and *UIN* is the resulting forecast error: *UIN=IN-EIN*.

We take into account the updating information process for CPI inflation in Spain. Following the standard model of inflation forecasting, it was assumed that the available 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*.

Therefore, *EIN* was obtained from a two-step procedure. In the first stage, in order to select the appropriate number of lags, inflation for the total period has been modelled as an ARMA process using the standard Box-Jenkins methodology. As is well known, the first step to model uncertainty with the variance of the error terms of the inflation model is to test if inflation is stationary (see Section III). If this is not the case, the variance of errors explodes and it makes no sense to use such a variance as a proxy of uncertainty.

As usual, the 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 indicate 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)

[4] The GARCH model implies that uncertainty changes slowly over time. Given the role played by the monetary policy in Spain during the period analyzed in this paper, we can assume such behaviour for uncertainty. The Spanish economy reached the highest inflation rates in 1977 and 1978, due to the political transition process and the oil crises. Since then, the monetary policy aimed at reducing inflation through different restrictive measures. In fact, inflation dropped from 26% in 1977 to rates below 8% in 1987, the first year of the period under study. Monetary policy has maintained its goal of stabilizing inflation, especially with the independence law of the Bank of Spain in 1994, the government’s commitment to meet the inflation targets for the incorporation into the euro, and after the EMU, with the 2% target set by the European Central Bank. Thus, inflation uncertainty should change slowly due to the credibility of the commitments of monetary policy to reduce and stabilize inflation.

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 homoscedastic forecast errors is obtained:

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

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

where $\sigma_{\varepsilon t}{}^2$ is inflation uncertainty (*UN*).

In the second step, once we have selected the optimal lag structure for the ARMA-GARCH model, we have calculated *EIN* and *UN* using the recursive coefficients technique. As expectations are based on past information, *EIN* was derived as follows: the expected value for January 1987 is calculated with the actual value from December 1979 to November 1986 and with the expected value for December 1986, and for the rest of the period we estimate the following model from December 1979 until *t* to derive EIN_{t+2}:

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

$$\sigma_{\varepsilon t}^2 = b_{1,t} \varepsilon_{t-1}^2 + b_{2,t} \sigma_{\varepsilon,t-1}^2 \,. \quad (5)$$

III. Monetary Regimes, Inflation and RPV: Basic Regression

Prior to the regression analysis, the stationarity of *IN* and *RPV* is checked by applying the Augmented Dickey-Fuller (ADF), Dickey-Fuller with GLS Detrending (DF-GLS) and Phillips-Perron (PP) unit root null tests to the seasonally adjusted series for the total period. The results are presented in Appendix B, Table A2. A unit root is rejected for inflation, even though only at 10% for the ADF and DF-GLS tests when we apply the Akaike criteria to select the optimal lag length. On the contrary, the results for *RPV* are ambiguous: both the ADF and DF-GLS tests fail to reject a unit root, while PP rejects it. Such differences can be due to the presence of structural breaks in the series. Hence, we apply the unit root tests proposed by Perron (1997) and Vogelsang and Perron (1998), which allow for a break in the series at an unknown time. The results are presented in Appendix B, Tables A3 and A4. Both *IN* and *RPV* present possible breaks from 1997.04 to 1998.05, and a unit root is rejected only for inflation.

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 unprocessed food and energy prices, from which the corresponding RPV has been calculated (*CRPV*). The ADF, DF-GLS and PP tests show different results for *CIN* and *CRPV*. Once we apply the tests proposed by Perron (1997) and Vogelsang and Perron (1998), a unit root with a break cannot be rejected for both variables. A possible break appears between 1997.08 and 1999.01 (see Appendix B for details). In all cases the breaks are associated with a change in the monetary policy regime, and with different inflation behaviour: the annualized monthly inflation rate slumped from 7.4% to 1.4% before the entry of Spain into the EMU, while it has been fluctuating between 5.3% and -0.8% thereafter.

A. Basic Regression Analysis

A first approach to the relation between inflation and RPV is obtained from OLS regression analysis. Taking into account the breaks mentioned above, we have run the OLS regression for the full period and for two sub-periods: before and after the EMU. In order to avoid distortions, the months in which the variables may present breaks were left out. Therefore, for *IN* and *RPV* we drop the period from 1997.04 to 1998.05. Hence, the first sub-period spans from 1987.01 to 1997.03, and the second from 1998.06 to 2009.09. For *CIN* and *CRPV*, we leave out the period from 1997.09 to 1999.01, so that we have two sub-periods: 1987.01-1997.08 and 1999.02-2009.08. Moreover, to capture the impact of inflation and deflation on price dispersion, *RPV* is regressed on the 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:

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

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

Table 1 presents the results of estimations of (6) and (7). They show that the coefficients of *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 relationship. Since the parametric model seems to be too restrictive to capture a changing relation, in the next section we undertake a stability test and a semiparametric approach.

[TABLE 1 HERE]

IV. Coefficient Stability and Non-Linearities

A. Coefficient Stability

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

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

Thus, changes in the inflation-RPV relationship can be outlined by the instability of the parameters over rolling samples. The results for $\beta_{1,t}$, our parameter of interest, which is obtained from different rolling and recursive regression windows, are reported in Figure 1. Recursive coefficients have been estimated by successive additions of one month to the 1987.01-1991.12 sub-sample and rolling equations have been estimated for fixed windows of six, eight and ten years. Figure 1 shows that in all cases $\beta_{1,t}$ is strongly unstable and decreasing in the second half of the total period. In the case of recursive coefficients estimation, this result indicates a changing marginal impact of *AIN* on *RPV* when new months are incorporated into the estimation. In turn, the rolling

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

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 sizes of the windows.[6]

[FIGURE 1 HERE]

In short, the empirical evidence shows an unstable relation between inflation and RPV. This varies significantly with the monetary policy regime. More precisely, coefficients are sensitive to the addition of years from 1998, and they drop and lose significance in the post-EMU period. Even more, $\beta_{1,t}$ reaches negatives values, which implies a negative relation between inflation and RPV. This negative value and, in general, the less influence of inflation on RPV, is associated with the fact that, after the entry of Spain into the EMU, the inflation rate was stabilised but RPV increased sharply. This was due to the fact that the existence of a single currency lowered the inflation of tradable goods and services compared to inflation of non-tradable ones. Hence, the lower the mean inflation the higher the difference between the inflation of tradable and non-tradable goods and services, and, therefore, RPV is high for both high and low inflation, and low for medium inflation. This explains that a linear specification for the inflation-RPV relationship does not seem to be adequate. We mean to tackle this issue in the next section.

B. Semiparametric Approach and Optimal Inflation

In order to find out additional information about the relation between inflation and RPV, we apply a partially linear model. As a preliminary step, we try to approximate the shape of such a relation by including a squared inflation term, as follows:

$$RPV_t = \alpha + \beta_1 IN_t + \beta_2 IN_t^2 + \sum_{h=1}^{12} \delta_h RPV_{t-h} + \varepsilon_t \cdot \qquad (9)$$

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

The results displayed in Table 2 show evidence of a U-shape for such a relationship for the total period and the second sub-period (the squared inflation is significant) but not for the first one (the squared inflation is not significant).

[TABLE 2 HERE]

The next step is to carry out a semiparametric analysis, which in turn allows us to obtain the optimal inflation. To compare this with our previous findings, the same number of lags for *RPV* and *IN* of equation (6) are included, so that the following equation is estimated:

$$RPV_t = \sum_{h=1}^{12} \delta_h RPV_{t-12} + g(IN_t) + \varepsilon_t \,, \tag{10}$$

where $g(IN_t)$ is an unknown non-linear smooth differential function, which relates inflation and price dispersion at time *t*. Therefore, our goal is the estimation of $g(IN_t)$ in (10).

The $g(IN_t)$ function has been estimated semiparametrically in two stages. In the first one we estimate λ_k from the following regression equation:

$$RPV_t = \sum_{h=1}^{12} \lambda_h \overline{RPV}_{t-12} + \eta_t \,, \tag{11}$$

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 we estimate $g(IN_t)$ non-parametrically from the regression:

$$\hat{\eta}_t = g(IN_t) + v_t \,, \tag{12}$$

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

In both stages we estimate the non-parametric regressions by applying a Nadaraya-Watson kernel regression estimator. Since the results of nonparametric regression are very sensitive to the bandwidth parameter (*h*), *h* has been selected using a mean squared forecast error (MSFE) criterion. Moreover, as the treatment of extreme values of inflation can affect the results, we use an unbounded Gaussian kernel and an outlier-robust Epanechnikov kernel.

[TABLE 3 HERE]

The results of MSFE for different values of *h* in the semiparametric estimation are presented in Table 3. This indicates that the optimal bandwidth parameter is higher for the outlier-robust kernel. In turn, to capture the sensitivity of *RPV* to marginal increases in inflation, after the estimation of $g(IN_t)$, we obtain the derivative of this function. If $g'(IN_t) > 0$ ($g'(IN_t) < 0$), *RPV* is increasing (decreasing) in inflation, and the optimal inflation rate is given by $g'(IN_t) = 0$. The derivative of $g(IN_t)$ was evaluated at 55 monthly inflation rates from -0.00097 to 0.01. The results for Epanechnikov and Gaussian kernels and different bandwidths are depicted in Figure 2. This shows that $g'(IN_t)$ are upward sloping and approximately linear for higher values of the bandwidth. Thus, $g(IN_t)$ is non linear and in fact can be represented by a U-shape functional profile, while $g'(IN_t) = 0$ for a positive inflation rate, which is specified below.

[FIGURE 2 HERE]

Table 4 presents the annual optimal inflation rate obtained for each bandwidth considered in Figure 2. The results show that the optimal inflation is 4.17% for both the Epanechnikov and Gaussian kernels and higher at higher bandwidths.[7] Such a value is similar to evidence found by Fielding and Mizen (2008) for the USA. In turn, our findings suggest an interesting issue in terms of monetary policy. As the inflation-RPV relationship presents a U-shape profile, and the optimal inflation rate is around 4%, additional efforts to reduce inflation under this value do not bring benefits of a lower RPV. Hence, the current goal of an annual inflation rate of 2% should increase price dispersion, which distorts the informational content of nominal prices, and thus impedes an efficient resource allocation.

[TABLE 4 HERE]

[7] We have checked our results for two additional outlier robust kernels: Cosine and Biweight. For both of them, the optimal bandwidth is 0.0013, the corresponding MSFE are $3.939*10^{-5}$ and $3.641*10^{-5}$ and the optimal annual inflation rates are 4.17% and 3.93% respectively.

Following the same methodology, we estimate the optimal inflation rate for different time periods. In order to compare these results with those of the total period, we keep the optimal bandwidth of each kernel (i.e., 0.0007 for the Gaussian kernel and 0.0013 for the Epanechnikov kernel).[8] The time periods were selected as follows: we consider the last month (2009.08) as fixed and change the initial date. Therefore, initially we carry out the estimations for the period 1988.01-2009.08, then for the period 1989.01-2009.08, and so on. Finally, the process stops when we consider only the post-EMU period. Table 5 presents the results, which show that the optimal inflation is around 4%-5%. This result differs from other findings, in particular from Choi et al. (2011). They find an optimal inflation within the announced target that, jointly with the inflation trend, declines over time. On the contrary, our results point out that the optimal inflation rate is higher than the inflation target of the euro zone, while it is not declining over time.

[TABLE 5 HERE]

To sum up, the evidence in Section IV hints at a non-linear relation between inflation and RPV, from which an optimal inflation rate around 4-5% is obtained. Furthermore, this result is robust to different time periods as Table 5 shows.[9] Moreover, inflation expectations and uncertainty can be the underlying causes behind this kind of relation (see, for example, Caraballo et al. 2006, Caraballo and Dabús 2008, Caglayan et al. 2008, Becker and Nautz 2009b and Choi 2010). This issue is studied in the next section.

V. Inflation Expectations and Uncertainty

There are several theoretical approaches trying to explain the links underlying the inflation-RPV relationship. Firstly, menu cost models assume that nominal price changes are costly, which imply that firms set prices discontinuously, according to an (S, s) pricing rule. 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*. The crucial point is that an

[8] Results presented in Table 5 do not change if we use different kernels, like Biweight and Cosine, or different values of the bandwidth. For the Biweight kernel, the optimal inflation is in the range of 3.64%-4.75%, and for the Cosine kernel in the range of 4.13%-4.91%. As far as the bandwidth is concerned, we have calculated the optimal inflation using the Epanechnikov kernel and values for the bandwidth from 0.0005 to 0.0031 and results do not differ from those in Table 5. These results are disposable from the authors upon request.

[9] These results reinforce those obtained in Section IV.A: monetary regime matters in order to reject a linear relation between inflation and RPV. In the period before the euro, high inflation was associated with a high RPV. The single currency implied an increase in RPV for low inflation, given the different behaviour between tradable and non-tradable inflation. This explains that a medium inflation rate of 4-5% minimizes RPV.

increase in the expected inflation induces a higher width of the (S, s) band to conserve on menu costs. This should reduce the price-change frequency, and then, assuming staggering price setting, increase price dispersion. Thus, this approach predicts that during deflationary periods the model works in reverse, so that RPV is increasing in the absolute value of the expected inflation (i.e., the relationship is V-shaped). Secondly, 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 the presence of firms with identical elasticity of supply, realized aggregate shocks do not have effects on RPV, because all firms respond identically to any given aggregate shock, while *ex ante* inflation uncertainty has a positive effect on RPV. In an extension of the signal extraction model, Hercowitz (1981) and Cukierman (1983) assume firms with different supply price elasticity, which implies different responses of prices to unexpected aggregate demand shocks. Thus, the higher the unexpected inflation the higher the RPV, i.e., the key factor is the size of the shock, while the sign of unexpected inflation is irrelevant.

Therefore, in order to find the links between inflation and RPV, this section introduces the components of inflation: expected and unexpected inflation, and uncertainty. All of them were estimated as explained in Section II. The *EIN* series shows a seasonal component which has been removed using the TRAMO-SEATS method. *UIN* is the difference between seasonally adjusted *IN* and seasonally adjusted *EIN*, whereas *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. Meanwhile, we test the implications of the extension of the signal extraction model mentioned above by means of the positive unexpected inflation (UIN^+) and the absolute value of negative unexpected inflation ($AUIN^-$). The lags of *RPV* that minimize the Akaike criterion are included. Besides, lags for *EIN*, UIN^+, $AUIN^-$ and *UN* were not included as regressors because lags of inflation were previously included in the calculus or inflation expectations, so that the persistence of inflation is taken into account in equation (13).

Finally, the following equation is estimated for the total period and the pre-EMU and post-EMU periods shown in Section III for the inflation-RPV relationship:

$$RPV_t = \alpha + \beta_0 EIN_t + \beta_2 UIN_t^+ + \beta_3 AUIN_t^- + \beta_4 UN_t + \sum_{k=1}^{12} \lambda_k RPV_{t-k} + \varepsilon_t \cdot \quad (13)$$

[TABLE 6 HERE]

Table 6 shows that *EIN* and *UN* are not significant. This result holds when recursive and rolling equations techniques are applied: none of them is significant regardless of the size of the window or the sample. Meanwhile, as far as unexpected inflation is concerned, both UIN^{+} and $AUIN^{-}$ are significant for the whole period, but for the pre-EMU period $AUIN^{-}$ is significant only at 10%, and for the post-EMU period UIN^{+} is not significant. These results are consistent with the U-shape inflation-RPV relationship found in Section IV: since both high and low inflation increase RPV, positive unexpected inflation implies a higher inflation than expected and thus a higher RPV, while negative unexpected inflation implies a lower inflation than expected, and then a higher RPV.

However, the Wald test fails to reject the null of $\beta_2 = \beta_3$, which indicates the relevance of unexpected inflation in explaining RPV. Rolling and recursive equations show that β_2 and β_3 decline during the period and they show weak sensitivity to sample or window size (see Appendix C). The latter result is in line with those presented in Section IV, given that if unexpected inflation explains the relation between inflation and RPV, and this relation becomes weaker during the period, the coefficients of unexpected inflation are supposed to decline as well.[10] Therefore, the role of unexpected inflation seems to be compatible with the signal extraction model's predictions. Nonetheless, these models refer to zero inflation, while our evidence is framed in a case of positive inflation, so that our results are only partially compatible with their predictions.

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

$$RPV_t = \alpha + \beta_0 EIN_t + \beta_4 UN_t + g(UIN_t) + \sum_{k=1}^{12} \lambda_k RPV_{t-k} + \varepsilon_t \cdot \tag{14}$$

The optimal bandwidth parameter (h) has been selected as in Section IV, and again unbounded Gaussian and outlier-robust Epanechnikov kernels have been used. Table 7 shows the results concerning the MSFE and, as we obtained in Section IV, the optimal h is higher for the outlier-robust kernel.

[10] Recent work presents 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 at a low level.

[TABLE 7 HERE]

The derivative of $g(UIN_t)$ was evaluated at 85 monthly unexpected inflation rates from -0.0095 to 0.0095. Figure 3 presents the results for the Epanechnikov and Gaussian kernels and different bandwidths. All $g'(UIN_t)$ are upward sloping and very close to a linear function, therefore $g(UIN_t)$ seems to be nonlinear and hints at a U-shape profile.

[FIGURE 3 HERE]

Table 8 summarizes the results. As the bandwidth decreases, the optimal value of *UIN* is nearer zero. According to the Epanechnikov kernel, the *UIN* that minimizes *RPV* is negative but near zero. Nonetheless, for the Gaussian kernel *RPV* is minimized at negative values of *UIN*, which can be due to the different treatment of the outliers in each kernel. In fact, the results for Cosine and Biweight kernels are very similar to those obtained with the Epanechnikov kernel: *UIN* around zero minimizes *RPV*.

[TABLE 8 HERE]

In short, 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 predictable monetary policy could minimize welfare costs caused by the impact of inflation on RPV.

VI. Summary and Conclusions

Evidence of a positive impact of inflation on RPV is widely documented. However, the features of this relation among different monetary and inflation regimes appear mixed. Similarly to previous papers for the USA and Japan, we find a U-shape inflation-RPV relationship. In turn, the optimal annual inflation rate is around 4%, and it remains despite the change in monetary regimes in 1999. Interestingly, this is higher than the inflation target proposed by the European Central Bank. Thus, the goal of an inflation rate lower than the optimal inflation should increase RPV, which is harmful to resource allocation and economic performance.

Moreover, our results suggest that the change of monetary regime, which has implied a change of the inflation regime and an in increase in RPV in the Spanish economy, helps to explain the U-shape inflation-RPV relationship: before the euro, high inflation was associated with a high RPV, but the lower inflation required by the single currency resulted in an increase in RPV, largely due to the differential inflation between tradable and non-tradable goods and services.

Furthermore, the mechanisms underlying the inflation-RPV relation is unexpected inflation. This 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. Besides, in order to minimize RPV, the evidence indicates an optimal value of unexpected inflation near zero, which has clear implications for monetary policy: the welfare costs of inflation can be lessened with a credible and predictable inflation targeting policy, even though the optimal target we find is larger than the EMU target. In this sense, this evidence requires further research. More precisely, it is necessary to investigate which factors could explain an optimal inflation around 4% for Spain, which is clearly above the target set by the EMU.

Appendix

A. Items of the Spanish CPI

[TABLE A1 HERE]

B. Unit roots tests

Table A2 presents the results of the ADF, DF-GLS and PP unit root tests for seasonally adjusted variables.

[TABLE A2 HERE]

On the other hand, we check for the existence of a unit root with structural breaks by applying the tests proposed by Vogelsang and Perron (1998). These tests allow us to distinguish two key properties: i) if the break affects the constant, the trend or both of them in the series, and ii) if the rupture impact on the variable is immediate (additional outlier) or gradual (innovational outlier). Taking into account the evolution of *IN* and *RPV*, we consider that additional outlier model must fit better to check structural breaks and unit root, because the entry into the euro affects *IN* and *RPV* once and for all. In turn, we select two models, one includes breaks in the constant and the trend, and the other one considers changes only in the trend.

Following Vogelsang and Perron (1998), testing for a unit root test in the additional outlier framework includes two steps. In the first one, the following equation is estimated:

$$y_t = u + \beta t + v^i DU_t + g^i DT_t + \varepsilon_t , \tag{A1}$$

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: i) if i=A the break only affects the constant, and $g=0$, ii) i=C indicates a rupture in the trend, and then $v=0$, and iii) i=B corresponds to the case that the rupture is in both the constant and the 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 (A1), we estimate by OLS (A2) if i=A, B, and (A3) if i=C.

$$\varepsilon_t = \alpha \ \varepsilon^i{}_{t-1} + \sum_{j=0}^{k} DTB_{t-j} + \sum_{j=1}^{k} \Delta\varepsilon^i_{t-j} + u_t \ , \tag{A2}$$

$$\varepsilon_t = \alpha\varepsilon^i_{t-1} + \sum_{j=1}^{k} \Delta\varepsilon^i_{t-j} + u_t \ , \tag{A3}$$

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

According to Vogelsang and Perron (1998), two data dependent methods can be applied to detect the breakpoints. The first one (method I) selects T_B that minimizes t_α (t-statistics corresponding to the estimated α in equations (A2) and (A3)). 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 (method II) 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 (A1)). 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 equation (A2) allows us to accept or reject unit root.

On the other hand, to choose the lag length *k* of ε^i_{t-j} in (A2) and (A3) we apply two criteria. The first one consists of choosing a fixed value for *k*, we have considered *k* =5 (as in Vogelsang and Perron (1998)).The second one is based on selecting a value of *k* (*k* =*k**) in such a way that in regressions (A2) and (A3) the 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 A3 y A4. These series show a change in the trend during the period, therefore in the paper we have taken into account the 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. As can be seen from Table A3 (model C), the unit root is rejected only for *IN*. With model B results are not conclusive with respect to *IN*, *CIN* and *CRPV*. The unit root cannot be rejected in all cases only for *RPV*.

[TABLES A3 AND A4 HERE]

C. Rolling and recursive equations for unexpected inflation

We estimate the following equation for different window sizes:

$$RPV_t = \alpha_t + \beta_{0,t} EIN_t + \beta_{2,t} UIN_t^+ + \beta_{3,t} AUIN_t^- + \beta_{4,t} UN_t + \sum_{k=1}^{12} \lambda_{k,t} RPV_{t-k} + \varepsilon_t \cdot$$

Figure A1 presents the results for $\beta_{2,t}$ and $\beta_{3,t}$. As usual, recursive coefficients have been obtained by successive additions of one month to the 1987.01-1991.12 sub-sample and rolling regressions have been estimated for a window of 8 years. As can be seen from Figure A1, both coefficients decline during the period: $\beta_{2,t}$ seems to be significant in the pre-EMU stage and $\beta_{3,t}$ is more sensitive to the sample considered. Similar results are obtained for windows of 6, 10 and 12 years (they are available from authors upon request).

[FIGURE A1 HERE]

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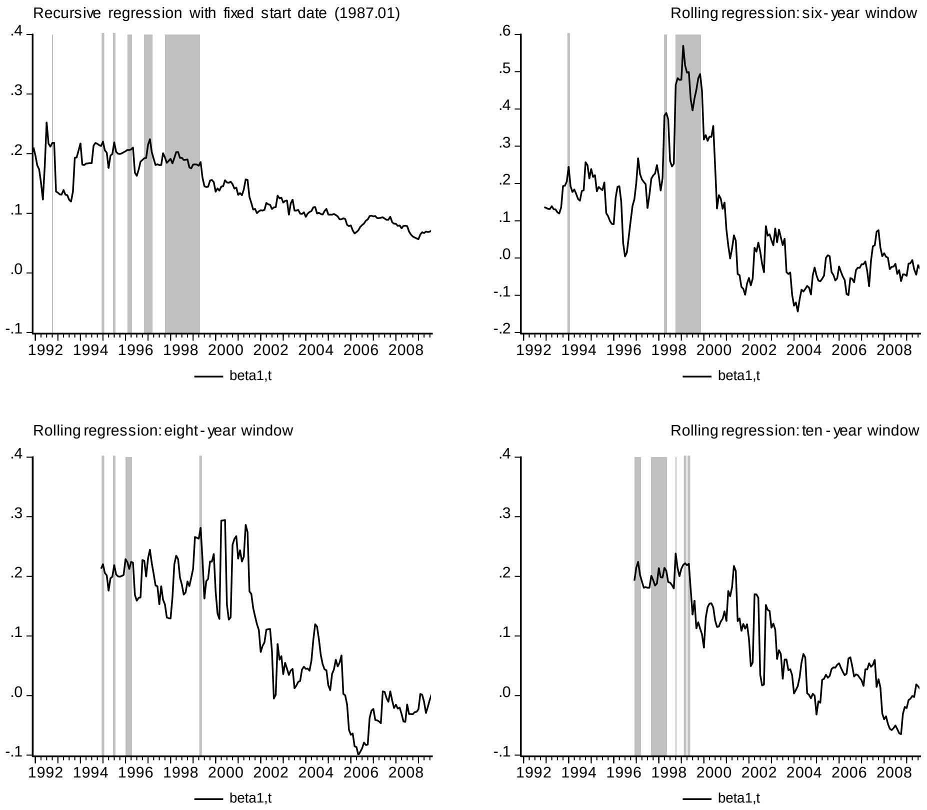
Figure 1. Recursive and rolling regressions



Notes: 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 beta1,t in 1992.12 captures the estimation of the parameter in (8) for 1987.01-1992.12, and so on.

Figure 2. Derivatives of *g(IN$_t$)*

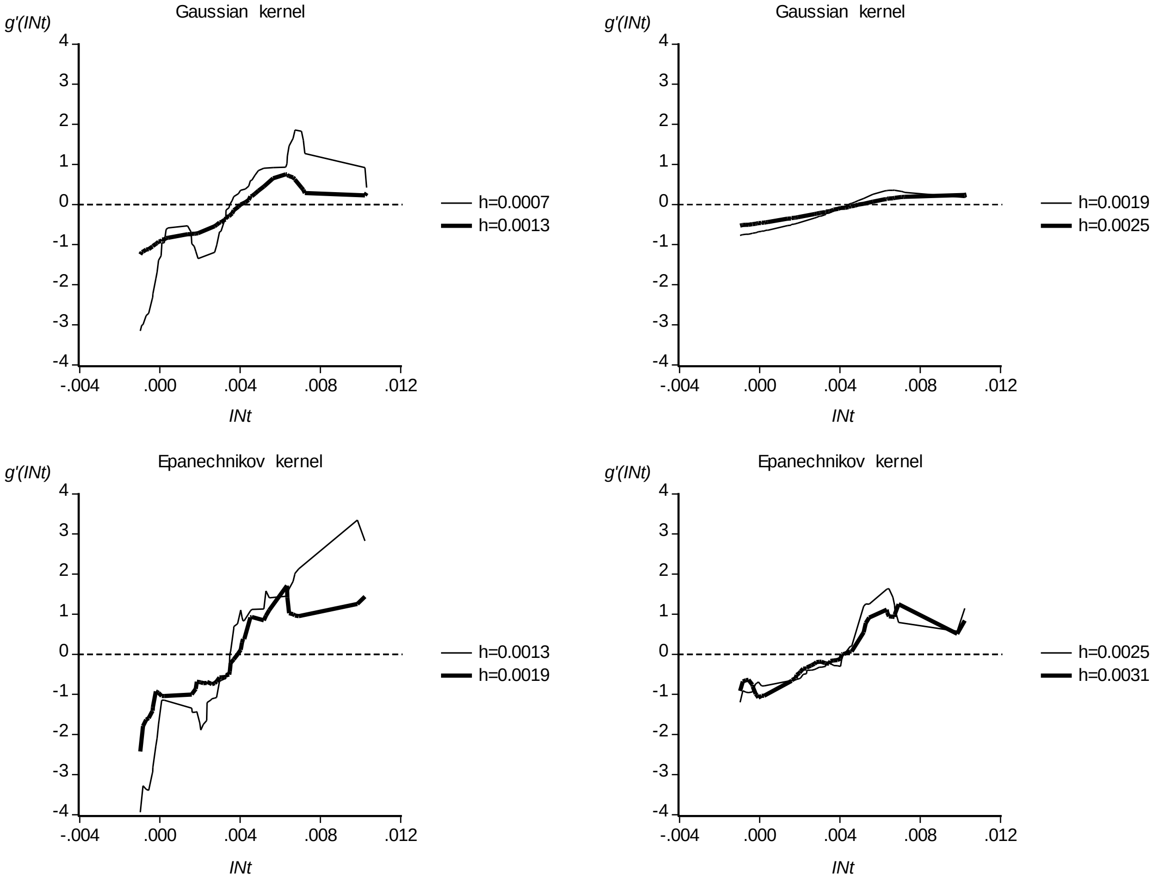


Figure 3. Derivatives of *g(UIN$_t$)*

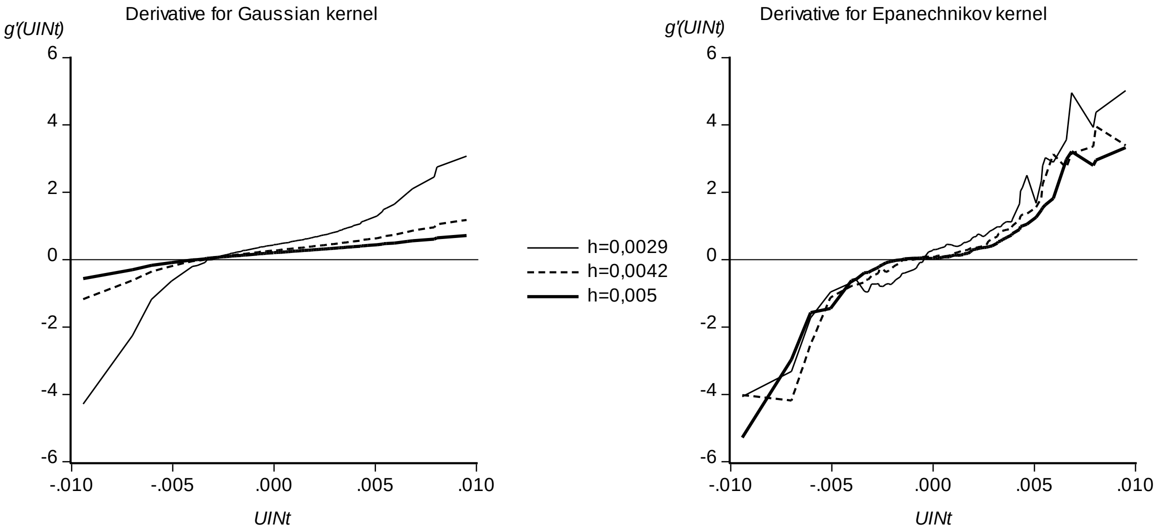


Table 1. Basic regression analysis

	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	-0.0001	0.001*	α	0.0004	0.003***	0.0005
	(0.18)	(0.88)	(0.08)		(0.27)	(0.00)	(0.44)
AIN_t	0.07	0.20*	-0.03	$ACIN_t$	0.11	0.47***	-0.15
	(0.29)	(0.08)	(0.65)		(0.43)	(0.00)	(0.49)
RPV_{t-1}	0.18***	0.16**	0.29***	$CRPV_{t-1}$	0.86***	0.77***	0.90***
	(0.00)	(0.03)	(0.00)		(0.00)	(0.00)	(0.00)
R^2adj.	0.81	0.78	0.85	R^2adj.	0.93	0.56	0.95

Notes: *, **, *** denote that the coefficients are significant at 10%, 5% and 1% levels respectively. The t-statistics are based on standard errors computed according to the Newey-West (1987) procedure to allow for residuals that exhibit both autocorrelation and heteroskedasticity of unknown form. Terms in parentheses are the p-values associated with t-statistics. To simplify the presentation only the first lag of *RPV* appears in the table.

Table 2. Inflation, squared inflation and *RPV*

DEPENDENT VARIABLE: RPV_t			
PERIOD	1987.01-2009.08	1987.01- 1997.03	1998.06-2009.08
α	0.001**	0.0008	0.001***
	(0.03)	(0.53)	(0.01)
IN_t	-0.21**	-0.30	-0.25***
	(0.03)	(0.45)	(0.00)
IN_t^2	41.88***	58.32	40.08**
	(0.00)	(0.15)	(0.03)
RPV_{t-1}	0.18***	0.16**	0.28***
	(0.00)	(0.03)	(0.00)
R^2adj.	0.81	0.78	0.86

Notes: *, **, *** denote that the coefficients are significant at 10%, 5% and 1% levels respectively. The t-statistics are based on standard errors computed according to the Newey-West (1987) procedure to allow for residuals that exhibit both autocorrelation and heteroskedasticity of unknown form. Terms in parentheses are the p-values associated with t-statistics. To simplify the presentation only the first lag of *RPV* appears in the table.

Table 3. MSFE for different values of the bandwidth parameter

Bandwidth parameter	Gaussian kernel	Epanechnikov kernel
0.0005	$7.149*10^{-5}$	$4.11*10^{-5}$
0.0007†	**7.134**$*10^{-5}$	$4.097*10^{-5}$
0.0013‡	$7.182*10^{-5}$	**$4.018*10^{-5}$**
0.0019	$7.252*10^{-5}$	$4.065*10^{-5}$
0.0025	$7.348*10^{-5}$	$4.140*10^{-5}$
0.0031	$7.457*10^{-5}$	$4.190*10^{-5}$

Notes: †Optimal bandwidth for Gaussian kernel. ‡Optimal bandwidth for Epanechnikov kernel.

Table 4. Optimal annual inflation rate

Bandwidth	Gaussian kernel	Epanechnikov kernel
0.0007†	**4.17%**	
0.0013‡	4.83%	**4.17%**
0.0019	5.31%	4.64%
0.0025	5.88%	5.04%
0.0031		4.91%

Notes: †Optimal bandwidth for Gaussian kernel. ‡Optimal bandwidth for Epanechnikov kernel.

Table 5. Optimal annual inflation rate at different time periods

PERIOD	Gaussian kernel	Epanechnikov kernel	PERIOD	Gaussian kernel	Epanechnikov kernel
1988.01-2009.08	4.42%	4.39%	1994.01-2009.08	4.70%	4.47%
1989.01-2009.08	4.83%	4.39%	1995.01-2009.08	4.91%	4.41%
1990.01-2009.08	4.38%	4.38%	1996.01-2009.08	4.91%	4.47%
1991.01-2009.08	3.93%	4.13%	1997.01-2009.08	4.91%	4.75%
1992.01-2009.08	4.35%	4.17%	1998.01-2009.08	5.10%	4.91%
1993.01-2009.08	5.10%	4.47%	1999.01-2009.08	5.49%	4.75%

Table 6. *RPV*, expected and unexpected inflation and uncertainty

	1987.01-2009.08	1987.01- 1997.03	1998.06-2009.08
α	0.00	-0.00	0.00
	(0.29)	(0.44)	(0.60)
EIN_t	-0.002	1.06	0.32
	(0.99)	(0.15)	(0.50)
UIN_t^+	0.31***	0.74***	0.06
	(0.00)	(0.00)	(0.61)
$AUIN_t^-$	0.24**	0.53*	0.17**
	(0.02)	(0.08)	(0.05)
UN	0.05	0.54	0.03
	(0.81)	(0.44)	(0.85)
RPV_{t-1}	0.18***	0.20***	0.27***
	(0.00)	(0.00)	(0.00)
R^2adj.	0.81	0.78	0.85
Wald test: H_o: $\beta_2 = \beta_3$	0.36	0.61	0.87
$\chi^2(1)$ statistics (p-values)	(0.54)	(0.43)	(0.35)

Notes: *, **, *** denote that the coefficients are significant at 10%, 5% and 1% levels respectively. The t-statistics are based on standard errors computed according to the Newey-West (1987) procedure to allow for residuals that exhibit both autocorrelation and heteroskedasticity of unknown form. Terms in parentheses are the p-values associated with t-statistics. To simplify the presentation only the first lag of *RPV* appears in the table.

Table 7. MSFE for different values of the bandwidth parameter

Bandwidth parameter	Gaussian kernel	Epanechnikov kernel
0.0023	$9.384*10^{-5}$	$7.912*10^{-5}$
0.0029†	**9.333**$*10^{-5}$	$7.869*10^{-5}$
0.0035	$9.349*10^{-5}$	$7.846*10^{-5}$
0.0042‡	$9.350*10^{-5}$	**$7.838*10^{-5}$**
0.005	$9.386*10^{-5}$	$7.873*10^{-5}$
0.0055	$9.396*10^{-5}$	$7.909*10^{-5}$

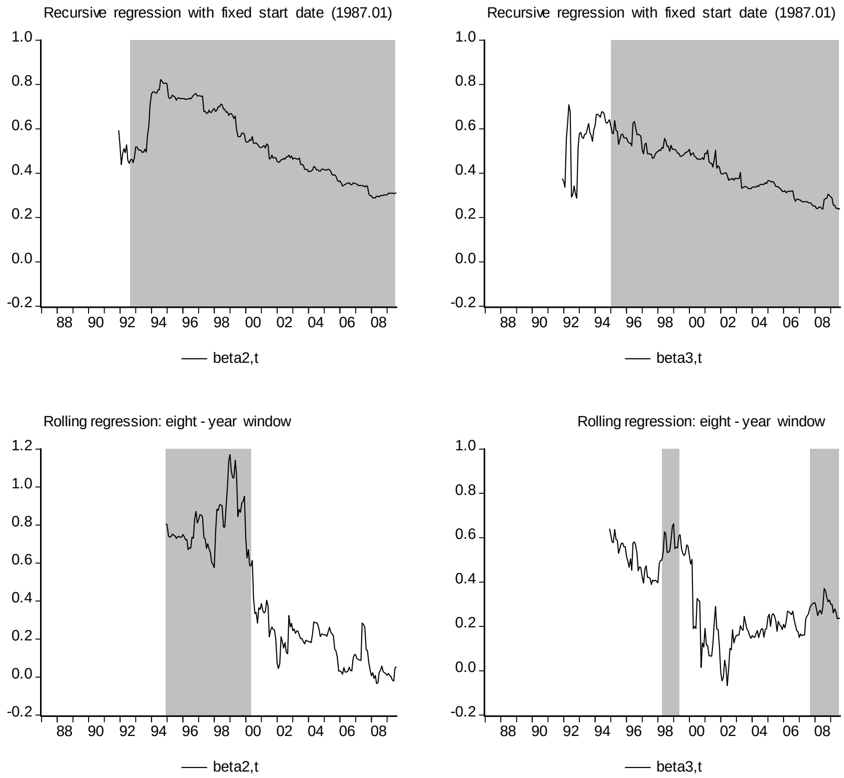
Notes: †Optimal bandwidth for Gaussian kernel. ‡Optimal bandwidth for Epanechnikov kernel.

Table 8. Optimal monthly unexpected inflation

Bandwidth	Gaussian kernel	Epanechnikov kernel	Cosine kernel	Biweight kernel
0.0023	-0.29%	-0.02%	-0.02%	-0.04%
0.0029†	**-0.30%**	-0.04%	-0.04%	-0.04%
0.0035	-0.33%	-0.08%	-0.06%	-0.06%
0.0042‡	-0.34%	**-0.09%**	**-0.07%**	-0.07%
0.0049*	-0.38%	-0.12%	-0.12%	**-0.09%**
0.0055	-0.41%	-0.18%	-0.12%	-0.12%

Notes: †Optimal bandwidth for Gaussian kernel. ‡Optimal bandwidth for Epanechnikov and Cosine kernels. *Optimal bandwidth for biweight kernel. The MSFE corresponding to the optimal bandwidth for cosine and biweight kernels are $7.789*10^{-5}$ and $7.908*10^{-5}$ respectively.

Figure A1. Recursive and rolling regressions for unexpected inflation



Note: the months for which the coefficients are significant are marked in grey lines (10% significance level). The number of the horizontal axis represent the ending month of each window.

Table A1. Items of the Spanish CPI

Data description	Average weights[†]	Data description	Average weights[†]
Cereals and by-products	13.837	Footwear for women	7.543
Bread	18.171	Footwear for children and infants	4.252
Bovine meat*	13.360	Repair of footwear	0.515
Sheep meat*	4.845	Rentals for housing	22.127
Swine meat*	7.199	Heating. electricity and water supply*	48.316
Poultry meat*	8.745	Maintenance and repair of the dwelling	41.330
Other meats*	22.167	Furniture and floor coverings	16.097
Fresh and frozen fish*	16.228	Household textiles and decorations	6.295
Seafood and processed fish	12.445	Household appliances including repair	11.146
Eggs*	3.108	Household utensils and tools	4.930
Milk	13.189	Non-durable household goods	17.205
Milk-based products	15.490	Household services	13.100
Oils and fats	8.049	Medical. dental and paramedical services	20.232
Fresh fruit*	15.988	Medical products. appliances and equipment	18.140
Canned and dried fruit	3.182	Personal transport*	164.378
Fresh vegetables*	9.686	Local transport*	5.621
Processed vegetables	5.143	Long-distance transport*	5.352
Fresh potatoes and potatoes preparations	3.494	Communications	27.660
Coffee, cocoa and infusions	3.281	Recreational items	23.874
Sugar	1.352	Printed matter	11.819
Other food products	9.175	Recreational services	14.580
Mineral waters, soft drinks and juices	6.761	Pre-primary and primary education	5.997
Alcoholic beverages	8.729	Secondary education	4.485
Tobacco	22.626	Tertiary education	6.605
Garments for men	27.945	Other educational goods and services	4.328
Garments for women	32.754	Personal effects	21.660
Garments for children and babyclothes	12.954	Tourism. catering and accommodation services	124.893
Clothing accesories and repair of clothing	5.309	Other goods and services	15.420
Footwear for men	6.886	Total	1000

Notes: [†]From 1987 to 2001 the weights are kept constant. Since then, they have changed each year. The table shows the average weights of each product over time. *These items were excluded from the core inflation calculus because they correspond to unprocessed food and energy-related items.

Table A2. ADF, DF-GLS and PP unit root tests (1987.01-2009.08)

Variable	Criteria to select lags	Constant and trend			Constant			No constant, no trend	
		ADF	DF-GLS	PP	ADF	DF-GLS	PP	ADF	PP
IN_t	Akaike	-3.27*	-2.84*	-11.03***	-1.93	-1.56	-9.78***	-1.24	-3.88***
	Schwarz	-10.69***	-10.32***		-9.11***	-5.08***		-1.24	
CIN_t	Akaike	-3.09	-2.50	-6.68***	-1.06	-0.54	-4.43***	-1.18	-2.41**
	Schwarz	-3.09	-2.50		-1.19	-0.70		-1.13	
RPV_t	Akaike	-1.36	-1.09	-4.61***	-1.57	-0.93	-4.60***	-0.54	-1.01
	Schwarz	-1.36	-1.09		-1.73	-0.93		-0.54	
$CRPV_t$	Akaike	-1.33	-1.09	-3.45**	-0.35	-0.02	-1.90	1.10	0.10
	Schwarz	-2.31	-1.78		-0.95	-0.74		0.22	

Notes:*,**,*** denote rejection of the null at 10%, 5% and 1% level of significance. A Bartlett kernel-based estimator of the frequency zero spectrum is used for the Phillips Perron test.

Table A3. Unit root tests with structural breaks. Method I

Model B: break in trend and constant									
Statistics		*IN* *k*=5	*IN* *k(t-sig)*	*CIN* *k*=5	*CIN* *k(t-sig)*	*RPV* *k*=5	*RPV* *k(t-sig)*	*CRPV* *k*=5	*CRPV* *k(t-sig)*
Min	T_B	1999.03	1997.08	1994.09	1995.03	1996.07	2000.12	2002.01	2002.01
t_α	t_α	-5.084**	-4.820*	-3.493	-4.82**	-3.908	-4.005	-6.396***	-5.008*
Model C: break in trend									
Statistics		*IN* *k*=5	*IN* *k(t-sig)*	*CIN* *k*=5	*CIN* *k(t-sig)*	*RPV* *k*=5	*RPV* *k(t-sig)*	*CRPV* *k*=5	*CRPV* *k(t-sig)*
Min	T_B	1997.07	1997.07	1997.08	1998.02	1997.08	1997.06	1999.01	1998.09
t_α	t_α	-4.748**	-4.070*	-3.052	-4.269	-3.708	-3.630	-3.643	-2.895

Notes: *,**,*** indicate significance at the 10%, 5% and 1% levels respectively. t_α: critical values in Vogelsang and Perron (1994). *k(t-sig):* the lag length *k* of ε^i_{t-j} has been selected (*k*=*k(t-sig)*) in such a way that in regressions (B2) and (B3) (for models B and C respectively) the coefficient corresponding to *k(t-sig)* is significant, while it is not significant for *k* >*k(t-sig)*.

Table A4. Unit root tests with structural breaks. Method II

		Statistics	*IN*	*CIN*	*RPV*	*CRPV*
Model B. Break in trend and constant	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
		t_α (*k(t-sig)*)	-4.698	-4.269**	-3.556	-2.699
Model C: Break in trend	Max t_g	T_B	1997.05	1998.05	1998.05	1998.02
		t_g	1.840	3.270	18.949	10.451

Notes: *,**,*** indicate significance at the 10%, 5% and 1% levels respectively. t_α: critical values in Perron (1994). *k(t-sig):* the lag length *k* of ε^{i}_{t-j} has been selected (*k*=*k(t-sig)*) in such a way that, in regression (B2), the coefficient corresponding to *k(t-sig)* is significant, while it is not significant for *k* >*k(t-sig)*. Obviously for model C, max t_g gives some information about T_B but it cannot be used to test unit root.