NOMINAL RIGIDITIES, SKEWNESS AND INFLATION REGIMES

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ABSTRACT

The menu costs model developed by Ball and Mankiw (BM)(1994,1995) predicts that inflation is positively related to the skewness of price changes distribution. We test this prediction in different inflationary contexts: Spain (1975-2002) and Argentina (1960-1989). We find a positive inflation-skewness relationship in both countries at low inflation, even though the mean annual inflation rates were very different: 2,2% for Spain and 23% for Argentina. Therefore, the threshold of *low* inflation under which the menu costs model is suitable is determined endogenously, and it depends on the inflationary experience of each economy. In the higher inflation periods skewness is not significant. Finally, our results suggest that the menu-costs model is not suitable beyond certain threshold of inflation.

1. Introduction

In a flexible price framework changes in relative prices should not affect average inflation and, therefore, the prediction is that there is no relationship between inflation and the higher moments of the relative price changes distribution. But empirical evidence does not support this result. On the contrary, inflation and the second and third moments of the relative price changes distribution appear to be positively correlated. However, there is no consensus about the causal mechanism underlying that relationship. On one hand, there is a vast empirical literature studying the relation between inflation and the second moment, the relative price variability (*RPV*), finding that causation runs from inflation to *RPV*. This strand of work dates back to Mills (1927), and since the contributions of Vining and Elwertowski (1976) and specially Parks (1978), a lot of empirical work has been done.

On the other hand, a second line of research proposed by Ball and Mankiw (henceforth BM) (1994,1995) shows that inflation is influenced by the skewness[1]. They argue that, in presence of nominal rigidities, due to the fact that firms face menu costs, changes in the price level and skewness are positively correlated. This paper is focused on this approach, and tries to check if the skewness-inflation relationship holds for different inflationary contexts. More precisely, our goal is to show that there is a threshold of the inflation rate under which the BM approach is suitable, and furthermore that such threshold is determined endogenously in each economy. The hypothesis is that this threshold depends on the inflationary experience of each country.

We test out such statement in two economies with very different inflationary history: Spain, from 1975 to 2002, and Argentina, from 1960 to 1989. The first economy has been historically stable in the last fifty years in comparison with Argentina: along the period studied in this paper the monthly inflation rate moved in a range between -1% and 4%. On the contrary, Argentina shows a very rich inflationary history: in the last forty years its monthly inflation rate fluctuated from -1.7% to 54%.

Our results show that the predictions of menu costs model hold for the lower inflation period in both countries, even though the mean inflation rate in each period differs strongly across them. In fact, the mean annual inflation rate in Argentina along the low inflation period was around 20%, higher than the inflation rate of Spain in the high inflation period. Nonetheless, in neither of them such approach is suitable at high inflation.

[1] Usually menu costs model has been used to explain nominal price rigidity, which implies that demand policies may be effective. BM move away from the traditional approach: they propose a theory of supply shocks. As they argue, supply shocks are changes in certain relative prices and they assert that menu costs model is a plausible framework to explain why those changes affect the price level.

The paper is organized as follows. Section 2 summarises the theoretical framework and the main empirical evidence. Section 3 presents the price data, variables and equations used in the empirical analysis. Section 4 shows the empirical results concerning the inflation-*RPV*-skewness relationship. Finally, section 5 concludes.

2. Theoretical framework and empirical literature

BM (1994,1995) use a menu costs model to explain how the economy responds to shifts in relative prices that, in a flexible price setting, would leave the price level unchanged. Within a menu costs framework, price adjustments are costly. Hence, when firms experience a shock to their desired relative prices, they only change their prices if the profit from the adjustment is larger than the menu cost. These menu costs give rise to a band of inaction in response to relative prices shocks. In that framework, a relationship between the inflation rate and the higher moments of the distribution of the desired price changes arises. The features of that relationship depend on the inflationary context.

On one hand, BM (1995) state that in an economy with no trend inflation, the average inflation rate is positively related to the skewness of the distribution of relative price changes. The intuition behind this result is illustrated in figures 1.a to 1.c[2], presented in appendix I. Those figures show how the skewness of the distribution of desired price changes influences the price level. As it was aforementioned, the presence of menu costs implies that firms have a range of inaction in response to shocks to their desired prices. If there is no trend inflation, such range is assumed to be symmetric around zero and it is between the upper (*U*) and the lower (*L*) cut-off prices. In figure 1.a the distribution of desired price changes is symmetric. In this case, if the desired changes are in the upper tail of the distribution –i.e., above *U*- firms will raise their prices, and if the desired changes are in the lower tail –i.e., under *L*-, firms will lower their prices. As the distribution is symmetric, both tails are equal and the net effect of the shock on the average inflation is zero. In figure 1.b the distribution of desired changes is skewed to the right (but still has mean zero); thus, the upper tail is larger than the lower tail. In this case, more prices rise than fall, so that the overall price level increases. In figure 1.c the distribution of shocks is skewed to the left, so the lower tail is bigger than the upper tail, which implies that more firms are lowering prices than raising them and the price level falls.

Moreover, a larger *RPV* will magnify the effects of skewness: if the distribution of shocks is symmetric, an increase in the variance of shocks increases the size of both tails by the same amount, so the price level remains unchanged. However, if the distribution is skewed to the right (left), a larger variance increases both tails, but the absolute increase in the upper (lower) tail is larger. Therefore the price level increases (decreases) by a larger amount. In short, *RPV* has no independent effect on inflation, but it interacts positively with skewness: a larger *RPV* is inflationary when the distribution is skewed to the right and deflationary when it is skewed to the left.

On the other hand, BM(1994) examine the effects of changes in relative prices in presence of a positive trend inflation, given a symmetric distribution of the desired price changes, concluding that price adjustments become asymmetric. In this context, when firms face a negative shock, they can either pay the menu cost and lower their prices or let inflation erodes their relative prices until the desired level. The higher the inflation, the faster the erosion process and the less likely the firms will pay menu costs. Therefore, a positive trend inflation will reduce the lower tail of the distribution, i.e., the size of the zone in which firms pay menu costs and lower their price. On the contrary, a positive shock implies that if the firm does not pay the menu cost, the gap between current and optimal price will widen. The firms are more likely to pay menu costs and raise their prices, increasing the upper tail of the distribution. Therefore, in a positive trend inflation framework, downward price rigidity appears. In other words, a positive trend inflation moves the range of inaction to the left (see figure 2.a). Finally, figure 2.b shows that an increase in *RPV* moves the distribution to the dotted line; hence, in absolute values the upper tail increases in relation to the lower one, so that inflation increases even if the distribution is symmetric[3].

As for some periods, both for Argentina and Spain, the features of inflation and the higher moments of the relative price changes distribution do not fit the aforementioned assumptions, we have to consider some additional cases:

1. Negative trend inflation and a symmetrical distribution of the desired price change: In this context, upwards rigidity appears and therefore the range of inaction moves to the right –see figure 3.a-, due to analogous reasons to those explained in an economy with positive trend inflation. Figure 3.b shows the impact of an increase in *RPV*. Such increase magnifies the lower tail of the distribution, which implies a negative relation between inflation and *RPV*.

2. Positive trend inflation and a distribution of desired price changes with positive skewness[4]: As it has been pointed out, a positive trend inflation moves the range of

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[2] These figures are based on BM (1995).

[3] Figures 2.a and 2.b are based on Amano and Macklem (1997).

[4] BM(1995) argue that if we combine the asymmetries in the distribution of the desired price changes with the asymmetric price adjustment derived endogenously by BM(1994) in an economy with trend inflation, we

inaction to the left, thus the upper tail will be bigger and the effects of an increase in *RPV* will be magnified –see figures 4.a and 4.b-.

3. Negative trend inflation and a distribution of desired price changes with positive skewness: In this case, the band of inaction moves to the right and, therefore, the right skewness might balance the impact of an increase in *RPV*, so that the negative inflation-*RPV* relationship can even disappear (see figures 5.a and 5.b).

Table 1 summarises the testable implications of menu costs model under the different assumptions considered in this section.

TABLE 1: TESTABLE IMPLICATIONS OF MENU COSTS MODEL*

A) NO TREND INFLATION
BM(1995)

DISTRIBUTION OF DESIRED PRICE CHANGES

SYMMETRICAL	SKEWED TO THE RIGHT	SKEWED TO THE LEFT
No *RPV*- π relation	Positive *S* - π relation Positive *RPV*- π relation *RPV* magnifies effect of *S*	Positive *S* - π relation Negative *RPV*- π relation *RPV* magnifies effect of *S*

B) TREND INFLATION

POSITIVE		NEGATIVE	
DISTRIBUTION OF DESIRED PRICE CHANGES		DISTRIBUTION OF DESIRED PRICE CHANGES	
SYMMETRICAL BM(1994)	SKEWED TO THE RIGHT	SYMMETRICAL	SKEWED TO THE RIGHT
Positive *RPV*- π relation	Positive *S* - π relation Positive *RPV*- π relation *S* magnifies effect of *RPV*	Negative *RPV*- π relation	Weak negative *RPV*- π relation Effect of *RPV* can be balanced by *S*

* π denotes inflation and *S* denotes skewness

The empirical evidence in this area is mixed. In general positive inflation-skewness and inflation-*RPV* relationships are supported by the data, but results are not conclusive about which relation is stronger in different inflationary contexts. On one hand, in low inflation contexts, the inflation-skewness relationship seems to be stronger than the inflation-*RPV* relationship. In this sense, Lourenco and Gruen (1995), for Australia, show that for periods with an annual inflation rate lower (higher) than 4%-5%, the inflation-skewness relation is stronger (weaker) than the inflation-*RPV* one. Studies for periods under that limit show similar results – see, among others, Ball and Mankiw (1995), for the US, Amano and Macklem (1997), for Canada, Aucremanne *et al.* (2002), for Belgium and Caraballo and Usabiaga (2004, 2007) for Spain. However, as an exception to this general result, Assarsson (2004) finds that in Sweden both relationships are positive and strong, and neither of them is stronger than the other.

expect that skewness still have a direct effect on inflation but there is also a direct effect of variance, however they do not specify the sign and relevance of such effects.

On the other hand, for studies covering periods with changing inflation rate, the evidence is mixed. For example, Hall and Yates (1998), for the 1975-1996 period in the United Kingdom, find a weaker inflation-skewness relationship than the inflation-*RPV* one. Döpke and Pierdzioch (2003), for the 1969-2000 period in Germany, find that both relations are positive, but none of them is stronger. Finally, Raftai (2004) for Hungary shows that there is a positive association between inflation and skewness along a period of an annual inflation rate moving from 15% to 30%.

In short, it seems that recent empirical evidence supports the existence of a positive association between inflation and the higher moments of price change distribution. Nevertheless, it is clear that the features of such relation change depending on the different rates of inflation. In order to give a wider evidence on that relationship, this paper analyses the relation between inflation and the higher moments of price change distribution in two different inflationary contexts: Spain and Argentina.

3. Price data and empirical methodology

3.1. Price data[5]

We use monthly price data for both countries. For Argentina, price series have been extracted from the statistical bulletins of the Instituto Nacional de Estadísticas y Censos, from January 1960 to March 1989. Individual price data correspond to the items of the national Wholesale Price Index (WPI), at the level of WPI groups (i.e. three digits of the International Standard Industrial Classification). Since the structure of WPI in Argentina changed in July 1984, we use 87 price indexes for the January 1960-June 1984 period and 64 for the July 1984-March 1989 period.

For Spain we use 24 categories of disaggregated price data of the Producer Price Index (PPI). They were extracted from the Instituto Nacional de Estadística for the January 1975-December 2002 period[6]. Along this period inflationary and deflationary processes can be found. There was a stagflation peaked in 1977 with 26% of annual inflation, while since

[5] As BM(1995) point out, one limitation of the theoretical framework explained in previous section is that it concerns the distribution of the *desired* price changes, which is unobservable. In order to give empirical content to their predictions, they use the distribution of *actual* price changes in place of the unobserved distribution of desired price changes. Following those authors, we also use the distribution of actual price changes.

[6] We have used WPI for Argentina and PPI for Spain because similar price indexes for both countries are required in order to compare results. Nonetheless, as the degree of disaggregation of price data is clearly different, we have checked if this fact could affect the results. In this sense, we have done the same empirical work performed in this paper using the Spanish CPI and comparing the results obtained for two different levels of disaggregation: 57 and 110 categories. The conclusions achieved for both cases are quite similar. These data and results are available from the authors upon request.

1986 the adjustment process, required for admission into the European Economic Community, was associated with a lower annual inflation (which was under 4%-5%).

For Argentina, the WPI price data do not present seasonality problems, because most of prices, and specially the prices of industrial and imported products, do not have a seasonal component. On the contrary, for the Spanish case, PPI price data present a seasonal component, which has been removed by means of the TRAMO-SEATS method. Thus, all the results of the estimations presented along the paper are referred to non-seasonal variables for Argentina and seasonally adjusted variables for Spain.

3.2. Empirical methodology

As it is common in this strand of the literature, we use the second and third moment of the price changes distribution: *RPV* and the skewness (S):

$$RPV_t = \left[\sum_{i=1}^{n} w_i \left(\pi_{it} - \pi_t\right)^2 \right]^{0,5} \quad [1]$$

$$S_t = \frac{\sum_{i=1}^{n} w_i \left[\pi_{it} - \pi_t\right]^3}{\left(RPV_t\right)^3} \quad [2]$$

where w_i is the weight of price *i* in the price index, π_{it} is the inflation rate of price *i* in period *t* and π_t is the inflation rate in period *t*. For the Spanish PPI, weights are calculated according to the importance of the branches of activity and the products in 1990, with the help of information provided by the Industrial Survey. For Argentina, w_i denotes the average expenditure share of the i^{th} good in the price index. As usual, weights are nonnegative and sum to one.

For Argentina, we use a slight variation of *RPV,* because in a high inflation economy expression [1] can be spuriously correlated with the mean of the distribution, i.e. the inflation rate. To avoid such problem, we use a coefficient of variations, as follows:

$$RPV_t = \frac{\left[\sum_{i=1}^{n} w_i \left(\pi_{it} - \pi_t\right)^2 \right]^{0,5}}{\left(1 + \pi_t\right)^2} \quad [3]$$

We estimate four equations to analyse the relationships among variables:

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \varepsilon_t \quad [4]$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_2 RPV_t + \varepsilon_t \quad [5]$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_3 S_t + \varepsilon_t \quad [6]$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_2 RPV_t + \beta_3 S_t + \varepsilon_t \quad [7]$$

Lagged inflation is included to capture persistence. As a preliminary step, we have applied the classical ADF test to the series (see appendix II)[7]. Price data present a deterministic trend both for Argentina and Spain, positive for the former and negative for the latter. These features of the data have been included in the regressions. Moreover, given that the explanatory variables are the higher moments of the price change distribution, multicollinearity could appear. To tackle this issue, the correlation coefficients between *RPV* and *S* have been calculated, as their values are under 0.3 -see appendix III- we have considered that both variables can be included jointly in the regressions.

4. Inflation, *RPV* and Skewness

This section presents the main empirical results. Along the paper, we estimate equations by means of Ordinary Least Squares (OLS) and we test for first and up to twelfth order autocorrelation in residuals using the Breusch-Godfrey (BG) Lagrange multiplier test. If no autocorrelation appears at a 5% level of significance, we present the results of the OLS estimate. If autocorrelation is detected, we estimate by Non Linear Least Squares and, previously, we model the structure of the residuals attending to the autocorrelation properties shown by the residuals series. As usual, the value of the t-statistic (p-values in brackets in the tables) is corrected of heteroscedasticity by means of the White method.

We run the regressions specified in equations (4) to (7). Table 2 shows the results for the total period in Argentina and Spain.

TABLE 2. TOTAL PERIOD

	ARGENTINA (1960:01-1989:03)*				**SPAIN (1975:02-2002:12)***			
Equations	**(4)**	**(5)**(‡)	**(6)**	**(7)** (‡)	**(4)**	**(5)**	**(6)**	**(7)**
Constant	-0.32 (0.31)	-2.23 (0.00)	-0.60 (0.04)	-2.29 (0.00)	0.64 (0.00)	0.22 (0.00)	0.59 (0.00)	0.21 (0.00)
Π_{t-1}	0.63 (0.00)	0.01 (0.91)	0.62 (0.00)	-0.009 (0.95)	0.46 (0.00)	0.36 (0.00)	0.45 (0.00)	0.36 (0.00)
RPV_t		2.06 (0.02)		2.96 (0.00)		0.41 (0.00)		0.39 (0.00)
S_t			0.18 (0.01)	0.05 (0.52)			0.02 (0.00)	0.01 (0.01)
Adjusted R^2	0.56	0.64	0.57	0.64	0.49	0.65	0.51	0.65
BG p-value	0.65	0.99	0.58	0.96	0.39	0.61	0.16	0.68

* Regressions include a deterministic trend
(‡) Estimates including a MA(3) structure

[7] The classical ADF test has low power under the presence of a structural break in the series. As a consequence, the test may falsely detect a unit root. For our data this problem with classical ADF test does not appear in the sense that the classical ADF has not detected unit roots. Moreover, for Spain the unit root test has been applied to the seasonally adjusted series. The methods of adjustment for seasonality introduce persistence, reducing the power of the test, in a way that tests are not able to reject non-stationarity. According to Ghysels (1990), this problem arises when seasonality is stochastic, therefore this problem should not affect our data -see Ghysels and Perron (1993) for literature related to the unit root test applied to seasonally adjusted series-.

In both countries the values of adjusted R^2 suggest that *RPV* appears to be more significant than *S* to explain the inflation rate. Moreover, in Argentina when both variables are included jointly in the regression (fourth column) skewness is not significant.

As our goal is to show that, under certain threshold of the inflation rate, menu costs model could be suitable even in high inflation economies, the next step is to test the stability of the coefficients along the whole period. In order to do that, we have employed two methods: the test for one or more unknown structural breakpoints proposed by Andrews (1993) and Andrews and Ploberger (1994) – *AP* test from now on- and the recursive residuals and the recursive coefficients estimates.

4.1. Stability tests

On the one hand, the *AP* test allows us to test for one or more unknown structural breakpoints in the sample. This test performs a single Chow Breakpoint Test at every observation between two dates t_1 and t_2. The *k* test statistics from those Chow tests are summarised into one test statistic for a test against the null hypothesis of no breakpoints between t_1 and t_2. From each individual Chow Breakpoint Test two statistics can be obtained: the Likelihood Ratio *F*-statistic and the Wald *F*-statistic. The former is based on the comparison of the restricted and unrestricted sums of squared residuals and the latter is computed from a standard Wald test of the restriction that the coefficients on the equation parameters are the same in all subsamples. When equations are linear, both statistics are equal.

The individual test statistics can be summarised into a *AP* statistics that is the simple average of the individual *F*-statistics:

$$AP = \frac{1}{k}\sum_{t=t_1}^{t_2} F(t) \qquad [8]$$

We also report the Maximum statistic which shows the maximum of the individual Chow *F*-statistics, allowing us to detect the most likely breakpoint location:

$$MaxF = \max_{t_1 \le t \le t_2} F(t) \qquad [9]$$

The distribution of both test statistics is non-standard. Andrews (1993) developed their true distribution, and Hansen (1997) provided approximate asymptotic *p*-values. We report the Hansen *p*-values. Moreover, the distribution of these statistics becomes degenerate as t_1 approaches the beginning of the equation sample, or t_2 approaches the end of the equation sample. To compensate for this behavior, the ends of the equation sample are not included in the testing procedure. We have considered two levels for this "trimming" , 15% and 10%, in a symmetric way, i.e., we remove the first and last 7.5% and 5%, respectively, from the observations.

We have applied this methodology to estimations of equation (7) for both countries and we have tested if there have been structural changes in the two coefficients we are interested in, that is, those associated to *RPV* and *S*.

We are going to consider linear equations for both countries[8], therefore, as it was mentioned above, the Likelihood Ratio *F*-statistic and the Wald *F*-statistic will be identical, that's why in the tables only one F-statistics for each case appears. We obtain three F-statistics for one or more unknown structural breakpoints: one for each regressor, *RPV* and *S*, and a third one when we consider the two regressors jointly. Tables 3 and 4 summarise the results.

TABLE 3. *AP* STATISTICS

	ARGENTINA			SPAIN		
Varying regressors	**trimming**	**AP-Statistics**	**Hansen p-value**	**trimming**	**AP-Statistics**	**Hansen p-value**
RPV_t	15%	6,30	0,00	15%	26,74	0,00
	10%	5,61	0,00	10%	23,83	0,00
S_t	15%	3,79	0,02	15%	1,61	0,16
	10%	4,76	0,00	10%	1,79	0,13
RPV_t and S_t	15%	7,87	0,00	15%	18,57	0,00
	10%	7,93	0,00	10%	17,17	0,00

TABLE 4. *MaxF* STATISTICS

	ARGENTINA				SPAIN			
Varying regressors	**trimming**	**MaxF-Statistics**	**Hansen p-value**	**Most likely breakpoint**	**trimming**	**MaxF-Statistics**	**Hansen p-value**	**Most likely breakpoint**
RPV_t	15%	19,83	0,00	1975.01	15%	142,40	0,00	1985.12
	10%	19,83	0,00	1975.01	10%	142,40	0,00	1985.12
S_t	15%	23,96	0,00	1984.10	15%	7,09	0,09	1985.06
	10%	28,12	0,00	1985.04	10%	7,09	0,11	1985.06
RPV_t and S_t	15%	24,94	0,00	1975.01	15%	90,25	0,00	1985.12
	10%	25,04	0,00	1975.01	10%	90,25	0,00	1985.12

Table 3 shows that for Argentina there are structural breaks for both coefficients when they are considered independently and when they are considered jointly. For Spain, the coefficient of *RPV* shows structural breaks but this evidence is very weak for *S*. From table 4, it can be seen that the most likely break point according to this test is 1975.01 for Argentina, when we consider only *RPV* or both regressors jointly, and 1985.12 for Spain.

[8] That is, we exclude the MA(3) structure for Argentina that was included for estimation of equation (7). Results concerning the value and significance of the coefficients don't change.

On the other hand, we have obtained the recursive residuals for those estimations in table 2 including the higher moments of the distribution as regressors[9] (i.e., equations [5] to [7]). Results show structural changes for Argentina around 1975 and around 1985, and for Spain around 1986 -see figures 1 and 4 in appendix IV for Argentina and Spain, respectively-. On the other hand, the recursive coefficients estimates have been calculated for the estimations in table 1 including both *RPV* and *S* (i.e., equation [7]). Results show that coefficients are not stable -see figures 2 and 3 for Argentina and figures 5 and 6 for Spain in appendix IV-: In Argentina, the coefficient of *RPV* increases in 1975 and the coefficient of *S* decreases slightly around 1975 and decreases again in a more pronounced way in 1985. In Spain, *RPV* increases and *S* decreases around 1985. As we will see in the following section, these results show that for both countries the coefficient of *RPV* is higher in the period with a higher mean inflation and the coefficient of skewness is higher in the lower inflation period. Moreover, results obtained with recursive residuals and recursive coefficients estimates reinforce those obtained in tables 3 and 4.

4.2. Inflation regimes

Our results do suggest the existence of structural changes in the estimations. The intuition behind these results is that the changes in the coefficients of the estimations correspond to a significant change in the inflation regime. This section is focused on this issue.

In order to determine the inflation regimes, we analyse the inflation series of each country by applying the same procedure as in Caraballo et al. (2006). This method captures only persistent changes, disregarding transitory variations in inflation levels. As the inflationary experiences of Argentina and Spain are very different, we have used different criterions to classify the inflation regimes. For Argentina we follow Leijonhufvud (1990)´s criterion: an economy is considered to be in a moderate inflation regime when monthly inflation rate is under 2%. High inflation corresponds to the 2%-10% range and very high inflation to the 10%-50% range. In turn, as in Spain the range of the inflation, and inflation rate itself, is substantially lower than in Argentina, we have considered a low inflation period when annual inflation rate is under 5%, and high inflation regimes otherwise. We have chosen this threshold for Spain because the empirical literature finding clear conclusions about positive relationship inflation-skewness is related with economies moving around that rate of inflation –see, for example, Ball and Mankiw (1995),

[9] Appendix IV includes the figures corresponding to estimations including both *RPV* and *S* as regressors. Results for estimations including only one of those two variables are very similar -they are disposable from

Lourenco and Gruen(1995), Amano and Macklem(1997), Aucremanne *et al* .(2002) and Caraballo and Usabiaga(2004,2007)-.

According to these criterions and by applying the method developed by Caraballo et al. (2006), we have obtained two main regimes in each country. In Argentina, the most relevant break is observed in February 1975. Therefore, two main regimes can be distinguished: a moderate inflation period from January 1960 to January 1975 and a high and very high inflation period from February 1975 to March 1989. In Spain, the most relevant break in the inflation series is observed in January 1986, so we can distinguish a high inflation period from January 1975 to December 1985, and a low inflation period from January 1986 to December 2002. These results imply that the changes in the coefficients of the estimations shown in section 4.1 correspond to a change in the inflation regime.

Once the two inflation regimes for both countries were distinguished, we analyse the main features of the variables into these regimes. Previously, we have applied the classical ADF test to the inflation rate. This test shows a positive deterministic trend for the Argentinean high inflation period and a negative deterministic trend for the Spanish high inflation period, while for both countries low inflation periods have no trend inflation[10].

As far as for the moments of the distribution is concerned, they show similar features in both countries. On one hand, *RPV* is higher and the range of oscillation is wider in high inflation periods than in low inflation periods. On the other hand, the distribution of price changes is clearly skewed to the right for both regimes in Argentina and for the high inflation period in Spain. On the average, skewness is lower and the range of oscillation is wider in low inflation than in high inflation periods. These features and the predictions of the menu costs model given such features are summarised in table 5.

the authors upon request-.

[10] Given that the deterministic trend appears only for the high inflation period for both countries, we have estimated again the regressions in table 2 taking this new result into account. We have not included them in this paper because there are no relevant changes.

TABLE 5. SUMMARY STATISTICS

COUNTRY		ARGENTINA		SPAIN	
INFLATION REGIME		LOW INFLATION NO TREND	HIGH INFLATION POSITIVE TREND	LOW INFLATION NO TREND	HIGH INFLATION NEGATIVE TREND
ANNUAL INFLATION (Π)	MEAN MIN. MAX.	23% 21% 58%	162% 46% 602%	2.2% -0.7% 6.4%	14% 7.9% 20.1%
MONTHLY INFLATION (Π)	MEAN MIN. MAX.	1.95% -1.70% 13.70%	10.95% 0.94% 54.05%	-0.17% -1% 1.6%	1.08% 0.0% 4.5%
RPV	MEAN MIN. MAX.	0.36 0.04 4.16	0.78 0.06 9.12	0.73 0.25 3.67	1.24 0.37 5.76
SKEWNESS (S)	MEAN MIN. MAX.	1.67 -8.35 8.72	2.64 -5.18 8.12	0.16 -16.79 18.57	1.62 -3.95 9.22
	% RIGHT* % LEFT*	75% 25%	89% 11%	60% 40%	89% 11%
PREDICTIONS OF THE MENU COSTS MODEL		▪ Positive *S-Π* relation ▪ Positive *RPV- Π* relation ▪ *S-Π* relation stronger than *RPV- Π* relation	▪ Positive *S-Π* relation ▪ Positive *RPV-Π* relation ▪ *RPV- π* relation stronger than *S- Π* relation	▪ Positive *S-Π* relation ▪ Positive *RPV-Π* relation ▪ *S-Π* relation stronger than *RPV-Π* relation	▪ Negative *RPV-Π* relation ▪ No prediction for *S-Π* relation ▪ Weak *RPV-Π* relation, it can be balanced by *S*

* Percentage of months in which the distribution of price changes is skewed to the right or to the left

The next step is to check if the predictions of the model hold in both periods. This leads us to estimate equations [4] to [7] for the low and high inflation regimes. Table 6 and 7 presents the results.

TABLE 6. LOW INFLATION REGIME

	ARGENTINA (1960:01-1975:01)				SPAIN (1986:01-2002:12)			
Equations	**(4)**	**(5)**	**(6)** (‡)	**(7)**	**(4)**	**(5)**	**(6)**	**(7)**
Constant	0.90 (0.00)	0.59 (0.00)	0.85 (0.00)	0.36 (0.06)	0.09 (0.00)	0.11 (0.03)	0.08 (0.00)	0.12 (0.01)
Π_{t-1}	0.54 (0.00)	0.51 (0.00)	0.30 (0.00)	0.42 (0.00)	0.46 (0.00)	0.46 (0.00)	0.47 (0.00)	0.47 (0.00)
RPV_t		1.00 (0.02)		0.75 (0.09)		-0.02 (0.72)		-0.04 (0.50)
S_t			0.30 (0.00)	0.29 (0.00)			0.02 (0.00)	0.02 (0.00)
Adjusted R^2	0.28	0.31	0.50	0.51	0.21	0.21	0.28	0.28
BG p-value	0.83	0.45	0.43	0.07	0.14	0.21	0.06	0.06

(‡) Estimate including a MA(1) term

TABLE 7. HIGH INFLATION REGIME

	ARGENTINA (1975:02-1989:03)*				SPAIN (1975:02-1985:12)*			
Equations	**(4)**	**(5)** (‡)	**(6)**	**(7)**	**(4)**	**(5)**	**(6)**	**(7)**
Constant	3.86 (0.02)	6.25 (0.02)	4.83 (0.02)	3.31 (0.08)	0.91 (0.00)	0.25 (0.04)	0.79 (0.00)	0.30 (0.01)
Π_{t-1}	0.61 (0.00)	-0.15 (0.35)	0.61 (0.00)	0.46 (0.00)	0.31 (0.00)	0.17 (0.01)	0.03 (0.00)	0.17 (0.00)
RPV_t		2.33 (0.03)		2.45 (0.02)		0.55 (0.00)		0.58 (0.00)
S_t			-0.24 (0.40)	-0.46 (0.17)			0.06 (0.04)	-0.04 (0.07)
Adjusted R^2	0.36	0.49	0.36	0.44	0.12	0.62	0.14	0.63
BG p-value	0.55	0.85	0.52	0.06	0.39	0.08	0.99	0.05

*A deterministic trend has been included for all regressions
(‡) Estimate including a MA(3) structure

Table 6 shows that in both cases results change drastically in comparison to those in table 2. According to the adjusted R^2, *S* seems to be much more relevant than *RPV*, and when both variables are included in the regressions *RPV* is not significant at 5% level. Thus, BM (1995) approach holds for low and stable inflation periods. In turn, the threshold under which menu costs model is suitable differs according to the inflationary history of each country.

Table 7 shows that in both cases the inflation-*RPV* relationship seems to be stronger than the inflation-skewness one. The contribution of *RPV* to the adjusted R^2 is larger than the contribution of *S*. In turn, in Spain the latter is not significant when both of them are included in the regression, meanwhile in Argentina only *RPV* is significant at a level of confidence of 5%. On one hand, results for Argentina do not corroborate the predictions obtained by BM(1994): although the inflation-*RPV* relationship is stronger than the inflation-skewness relation[11], skewness is not significant to explain the inflation rate. Results for Spain do not support menu costs predictions either. *RPV* coefficients were expected to be negative but they are positive; and *RPV*-inflation relation was expected to weaken once skewness were included, while table 7 shows that the coefficient is positive and *RPV* is still significant when skewness is included. The intuition is that nominal rigidities tend to disappear in higher inflation periods.

[11] To check these results in another high inflation country, we have done a similar analysis for Peru. We used 168 individual prices from the CPI for the January 1980-April 1994 period. By applying the criterion used for Argentina, the Peruvian inflation presents two periods of high inflation with a mean monthly inflation rate around 5%, and a very high inflation period with a mean monthly inflation rate of 44%. *RPV* is significant to explain inflation for the total period, and the adjusted R^2 increases from 0.18 to 0.91 when such variable is included in the regression, while *S* is not significant. In turn, BM approach is not suitable in the high inflation period (1991-1994), even tough this is the lower inflation period in Peru: *RPV* is significant, but the contribution to the adjusted R^2 is smaller and *S* is not significant. Hence, these results point out that there is also a limit from which BM approach doesn't work. In particular, it is not suitable beyond certain thresholds of inflation.

To sum up, the menu costs model holds for low inflation in Argentina when annual inflation is around 20%, while it is not suitable for the Spanish higher inflation period, even though the average inflation rate was only around 14%. These results are favourable to our hypothesis: it seems that there is a threshold of inflation under which the predictions of menu costs model hold, and such limit depends on the inflationary history of each economy. In particular, menu costs approach is suitable in the low inflation periods of two countries with very different inflationary experiences, even though their inflation rates in such periods were substantially different.

4.3 Alternative measures of skewness

This section test the relevance of skewness in low inflation regimes defining alternative measures of skewness. BM(1995) relate inflation with the size of the tails of the price changes distribution; therefore, it seems more accurate to define a variable to measure the tails and also to capture the magnifying effect of *RPV* on skewness. Specifically, for a cut-off X chosen arbitrarily, SX_t is defined as:

$$SX_t = \sum_{i=1}^{n} w_i (\pi_{it} - \pi_t) D_i^- + \sum_{i=1}^{m} w_i (\pi_{it} - \pi_t) D_i^+ \quad [10]$$

where D_i^- and D_i^+ are dummy variables. The former term takes the value one when i_{th} industry's relative price change falls in the lower X per cent of the distribution and zero otherwise, and the latter term is one when i_{th} industry's relative price change falls in the upper X per cent of the distribution and zero otherwise. Therefore, SX_t subtracts the mass in the upper tail of the distribution of prices changes from the mass in the lower tail. This variable is zero for a symmetrical distribution of relative price changes and positive (negative) when the right (left) tail is larger than the left (right) tail. Moreover, for a given skewness, the larger the *RPV* the larger the tails; thus the same variable combines the effects of skewness with its interaction with *RPV*. As the choice of X is arbitrary, we have chosen X=10 and X=25 in order to compare our results with those of BM (1995) and Amano and Macklem (1997).

Finally, instead of giving full weight to the price changes above a cut-off and zero weight otherwise, as with SX_t, BM(1995) define a new variable which increases the weights linearly with the size of the adjustment, as follows:

$$Q_t = \sum_i w_i \left|\pi_{it} - \pi_t\right| \left(\pi_{it} - \pi_t\right) \quad [11]$$

Therefore, Q_t is a weighted average of the product of each relative price change and its own absolute value, with the properties of SX_t: it is zero for a symmetrical distribution and positive (negative) for a right (left) skewed distribution. In turn, its value is magnified with a larger *RPV*.

We estimate the following equations:

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_4 S10_t + \varepsilon_t \quad [12]$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_5 S25_t + \varepsilon_t \quad [13]$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_6 Q_t + \varepsilon_t \quad [14]$$

Again, we carry out the estimations for the total period and test the stability of the parameters. Finally, we estimate equations [12] to [14] for the two inflation regimes. Tables 8 and 9 show the results for Argentina and Spain, respectively,

TABLE 8. ALTERNATIVE MEASURES OF SKEWNESS. ARGENTINA*

	TOTAL PERIOD			**LOW INFLATION PERIOD**			**HIGH INFLATION PERIOD**		
Equations	**(10)**	**(11)**	**(12)**	**(10)** (‡)	**(11)**	**(12)**	**(10)**	**(11)**	**(12)**
Constant	2.12 (0.00)	2.29 (0.00)	2.41 (0.00)	1.17 (0.00)	0.98 (0.00)	0.85 (0.00)	3.33 (0.06)	3.61 (0.02)	3.75 (0.02)
Π_{t-1}	1.62 (0.00)	0.65 (0.00)	0.62 (0.00)	0.29 (0.00)	0.51 (0.00)	0.50 (0.00)	0.60 (0.00)	0.63 (0.00)	0.60 (0.00)
$S10_t$	0.02 0.08			0.01 (0.00)			0.02 (0.17)		
$S25_t$		0.01 (0.23)			0.01 (0.03)			0.01 (0.17)	
Q_t			0.11 (0.39)			1.18 (0.00)			0.05 (0.67)
Adjusted R^2	0.58	0.57	0.56	0.38	0.31	0.37	0.38	0.38	0.36
BG p-value	0.12	0.39	0.61	0.96	0.87	0.07	0.23	0.35	0.54

*A positive deterministic trend for the high inflation period has been included for all regressions
(‡) Estimate including a MA(1) term

TABLE 9. ALTERNATIVE MEASURES OF SKEWNESS. SPAIN*

	TOTAL PERIOD			**LOW INFLATION PERIOD**			**HIGH INFLATION PERIOD**		
Equations	**(10)**	**(11)**	**(12)**	**(10)**	**(11)**	**(12)**	**(10)**	**(11)**	**(12)**
Constant	0.58 (0.00)	0.64 (0.00)	0.56 (0.00)	0.09 (0.00)	0.03 (0.35)	0.09 (0.00)	0.83 (0.00)	0.90 (0.00)	0.80 (0.00)
Π_{t-1}	0,39 (0.00)	0,46 (0.00)	0.44 (0.00)	0.42 (0.00)	0.41 (0.00)	0.47 (0.00)	0.20 (0.02)	0.31 (0.00)	0.29 (0.00)
$S10_t$	1.83 (0.00			0.97 (0.04)			2.19 (0.00)		
$S25_t$		-0.03 (0.88)			0.34 (0.00)			-0.68 (0.40)	
Q_t			1.36 (0.00)			3.91 (0.03)			1.47 (0.00)
Adjusted R^2	0.56	0.49	0.55	0.25	0.26	0.27	0.30	0.13	0.25
BG p-value	0.73	0.39	0.52	0.13	0.11	0.10	0.64	0.73	0.42

* A negative deterministic trend for the high inflation period has been included for all regressions

In general the results boost the conclusions obtained previously. On one hand, results shown in tables 8 and 9 show that the three alternatives measures of skewness are significant in the low inflation period for both countries. This implies that skewness is significant in low inflation and its effect is magnified by *RPV*. In turn, for the Argentinean high inflation periods none of the variables is significant, which implies that menu costs model is not suitable.

On the other hand, no clear conclusions arise for total and high inflation periods in Spain. The fact that for those two periods $S10_t$ is significant and $S25_t$ is negative and non-significant is implying that the choice of the cut-off may be relevant in order to explain inflation, against conclusions obtained by Ball and Mankiw (1995) and Amano and Macklem (1997).

5. Conclusions

This paper analyses the relevance of menu costs models, performed by BM (1994,1995), in two countries with very different inflationary experiences: Argentina and Spain. For low and stable inflation periods, BM (1995) approach predicts a strong positive relation between inflation and skewness, which can be magnified by *RPV*. Our results show that such relation holds in the lower inflation periods of both countries, even though their inflation rates are very different. Therefore, these results seem to verify our hypothesis: the limit of *low* inflation differs in order to apply BM framework. For Spain, that barrier could be 4%-5% of annual inflation rate, whereas for Argentina it reaches 20%. The intuition is that such limit depends on the inflationary experience of the economy.

For high inflation periods both countries present a deterministic trend. In this context BM (1994) assert that both *RPV* and skewness are significant but *RPV* is more significant than skewness in order to explain inflation. Our results show that the inflation-*RPV* relationship is stronger than the inflation-skewness one in both countries, but skewness is not significant in any of them. Such results suggest the relevance of inflation regime in explaining both relationships and state that nominal rigidities disappear at high inflation. In short, beyond an upper threshold menu costs model is not suitable, and that limit seems to be endogenous to the inflationary history of the economy.

Finally, a natural extension of this paper is to take a higher number of countries, with different inflationary experiences, in order to determine if our results hold in an expanded sample of cases.

APPENDIX I. THEORETICAL FRAMEWORK
FIGURE 1. NO TREND INFLATION

FIGURE 1.a: SYMMETRIC DISTRIBUTION

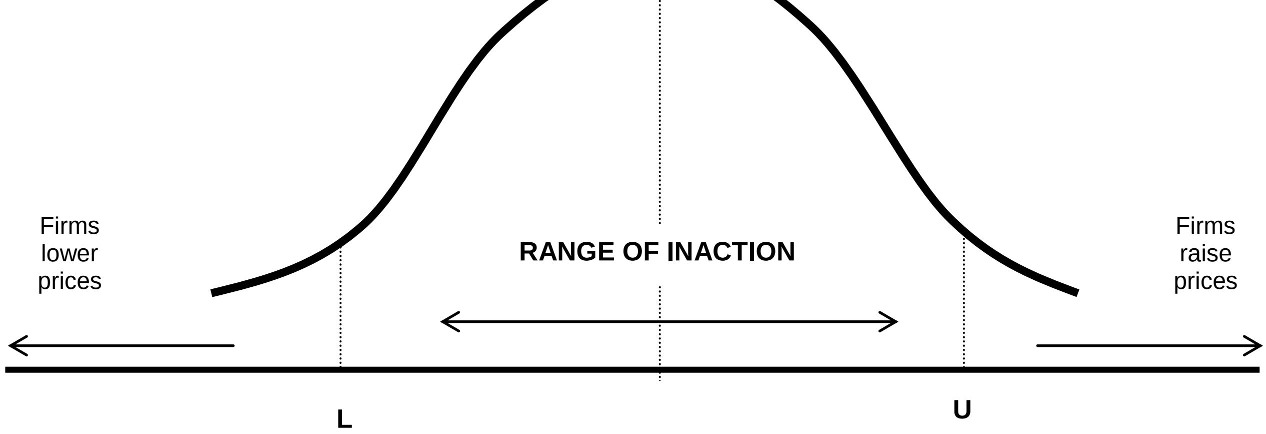


FIGURE 1.b: SKEWED TO RIGHT

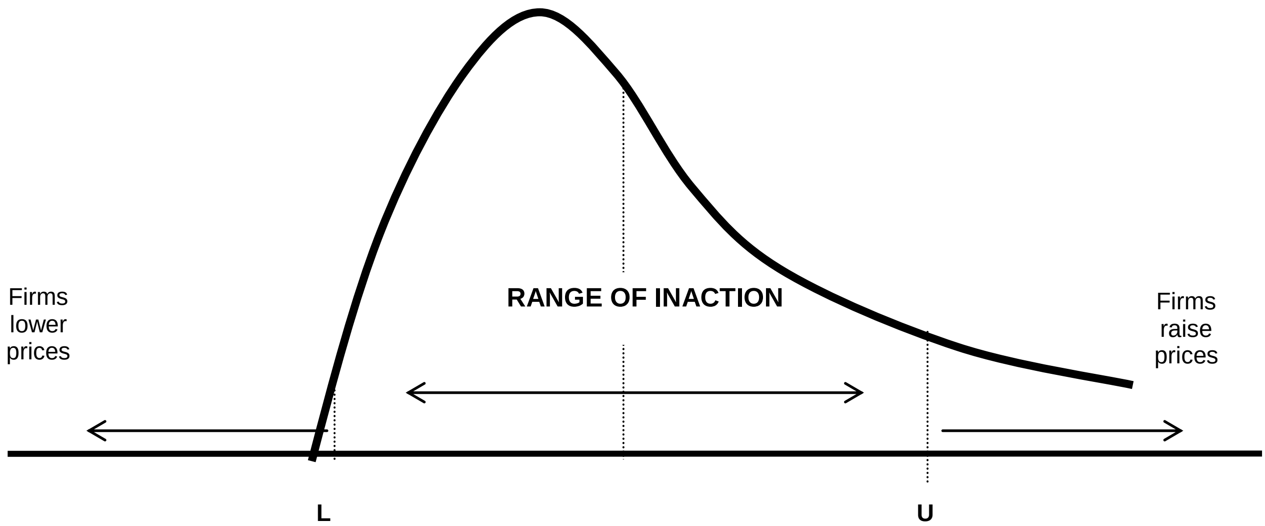


FIGURE 1.c: SKEWED TO LEFT

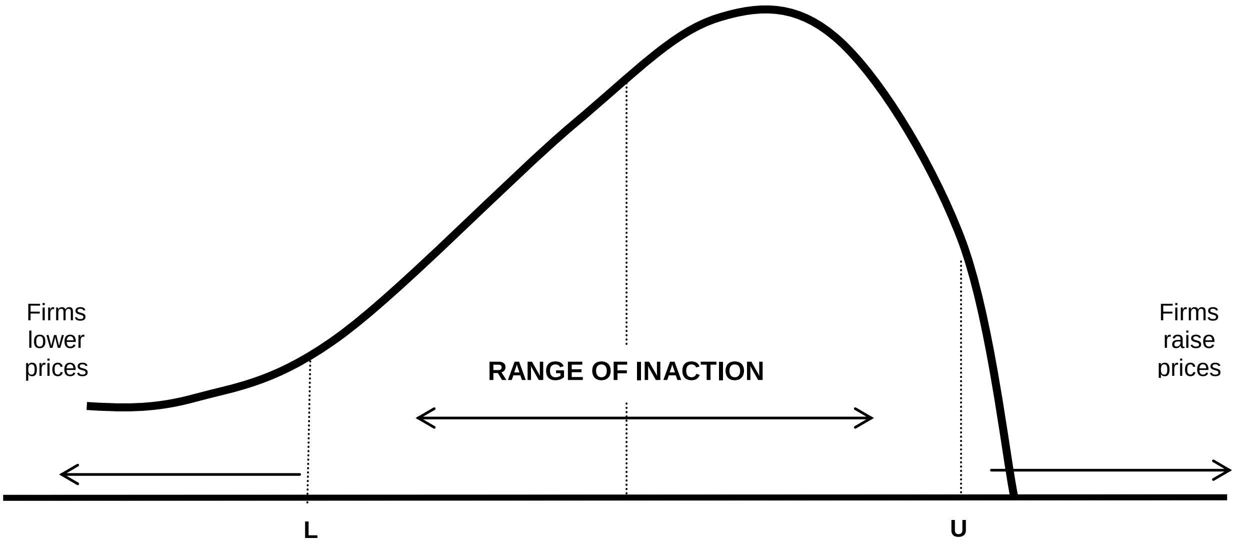


FIGURE 2 . POSITIVE TREND INFLATION. SYMMETRIC DISTRIBUTION

FIGURE 2.a

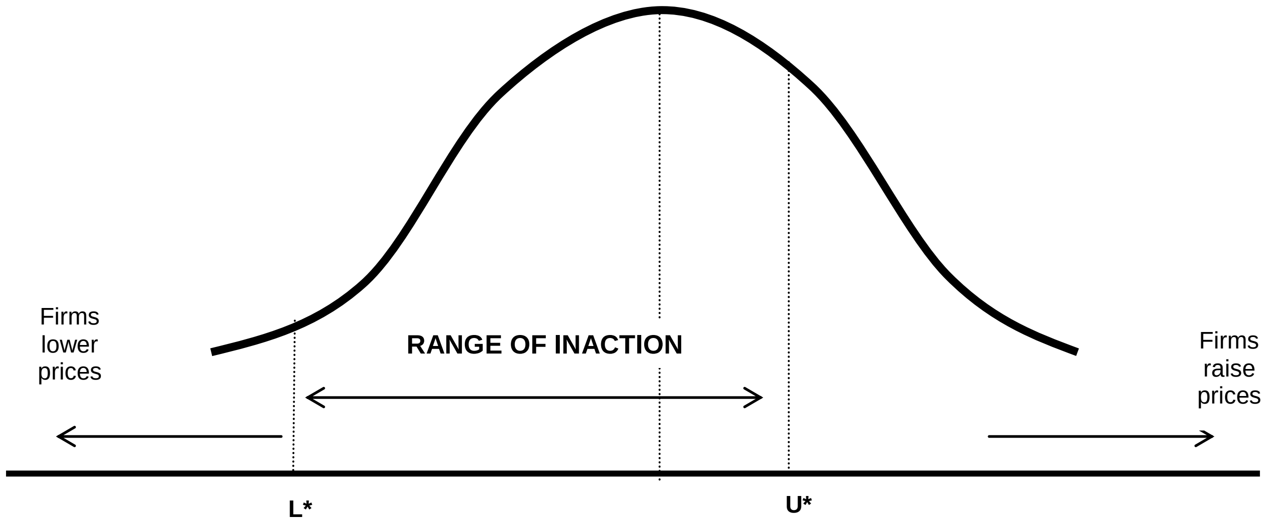


FIGURE 2.b

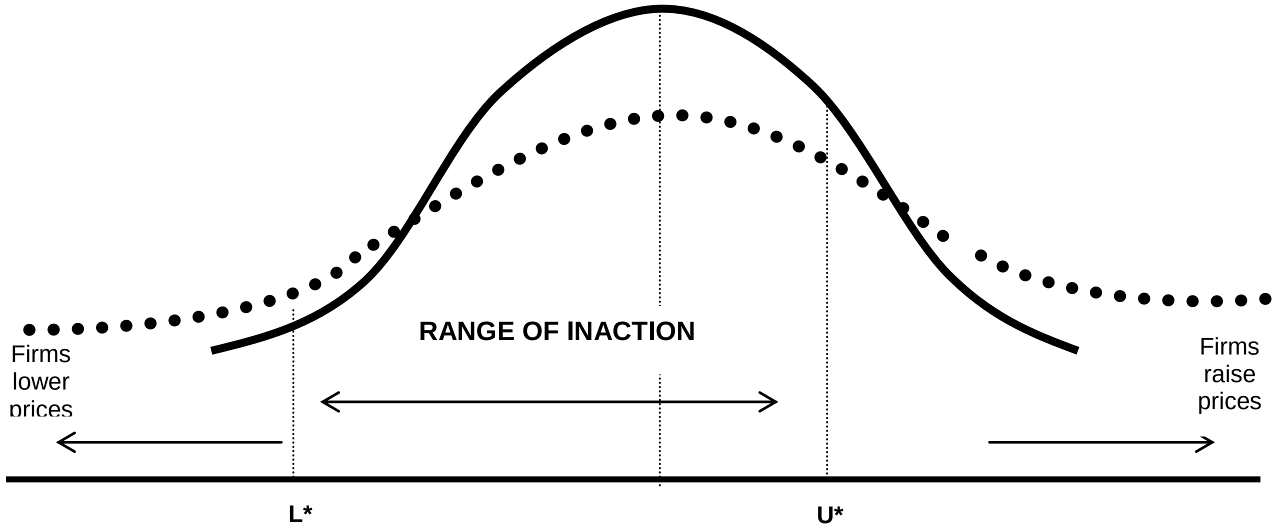


FIGURE 3 . NEGATIVE TREND INFLATION. SYMMETRIC DISTRIBUTION

FIGURE 3.a

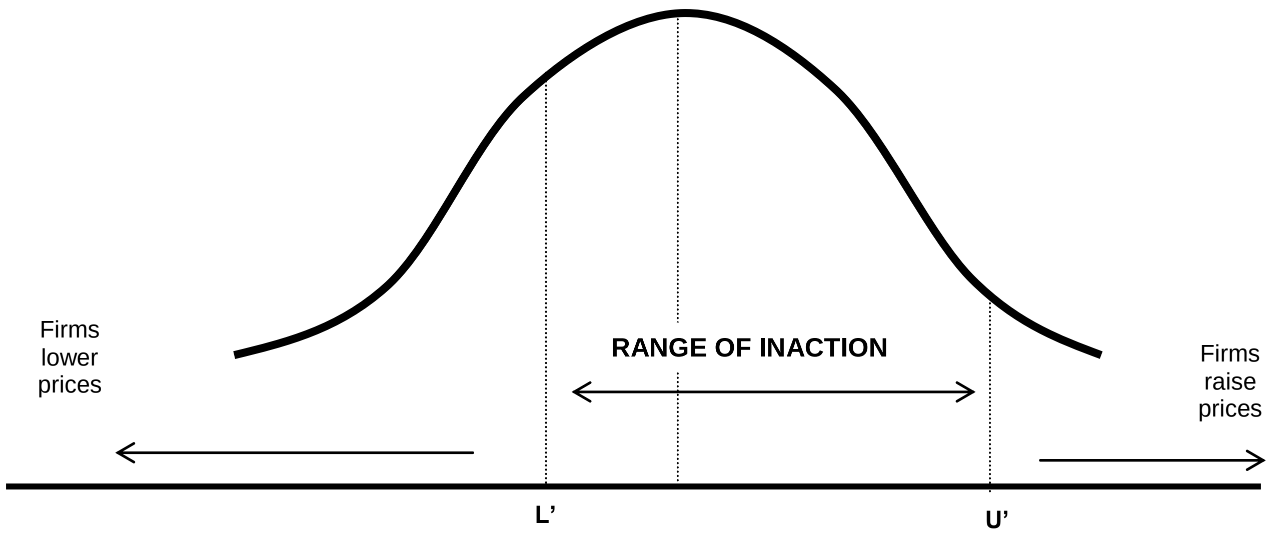


FIGURE 3.b: EFFECTS OF AN INCREASE IN RPV

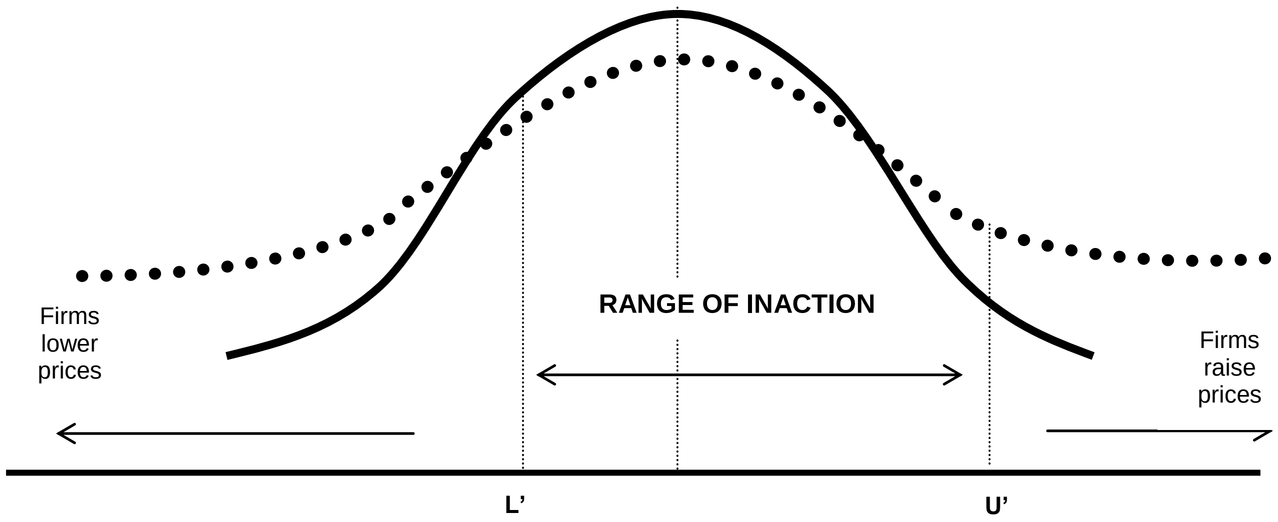


FIGURE 4. POSITIVE TREND INFLATION. RIGHT SKEWED DISTRIBUTION

FIGURE 4.a

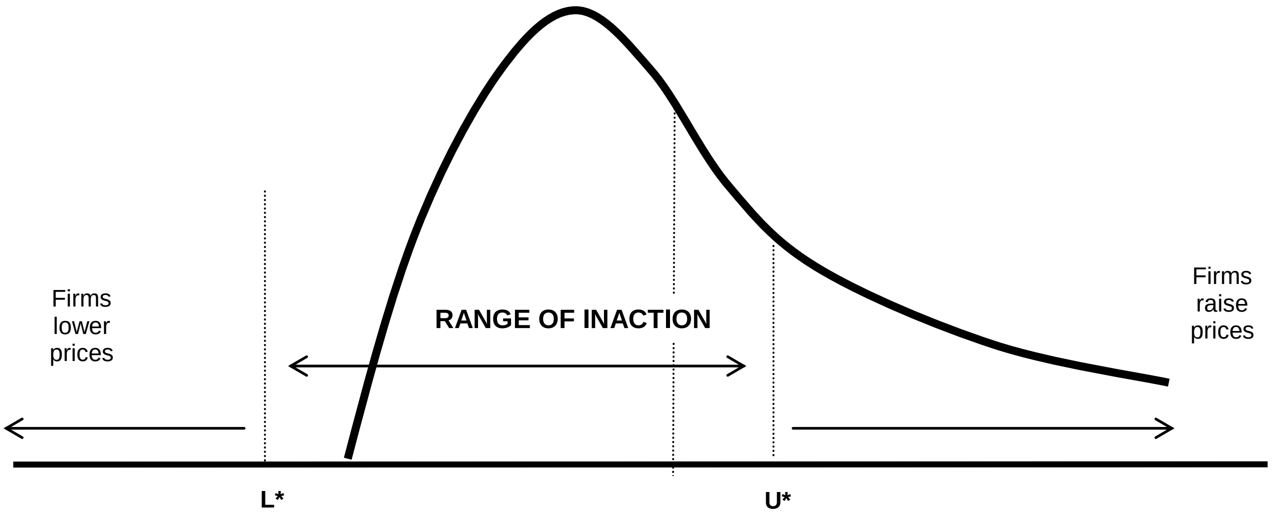


FIGURE 4.b: EFFECT OF AN INCREASE IN RPV

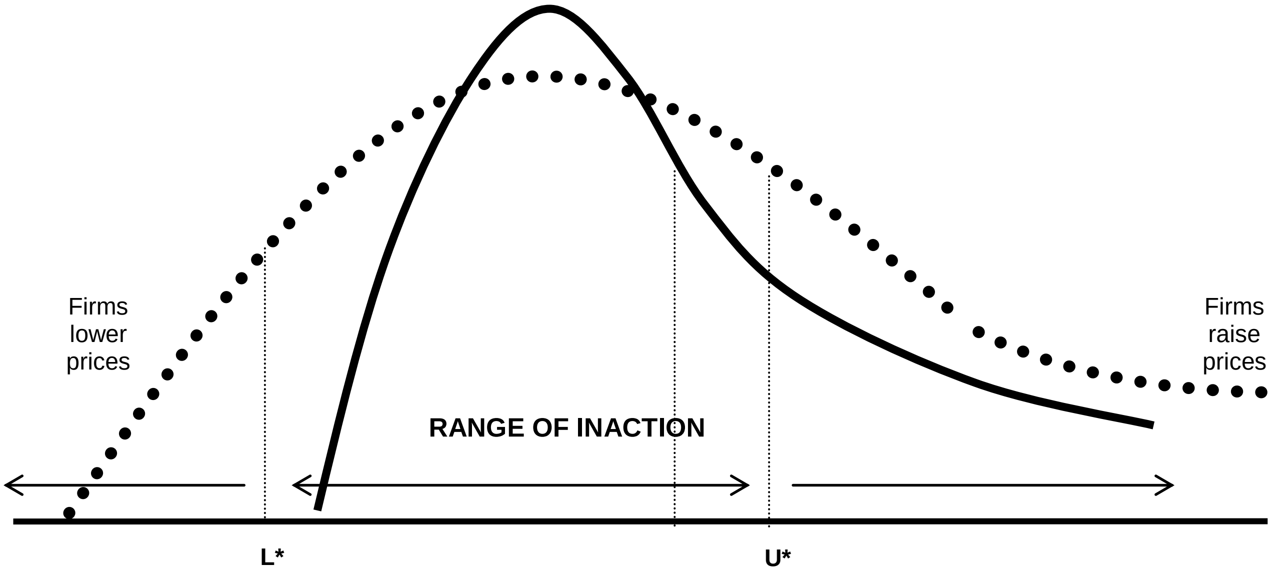


FIGURE 5. NEGATIVE TREND INFLATION. RIGHT SKEWED DISTRIBUTION

FIGURE 5.a

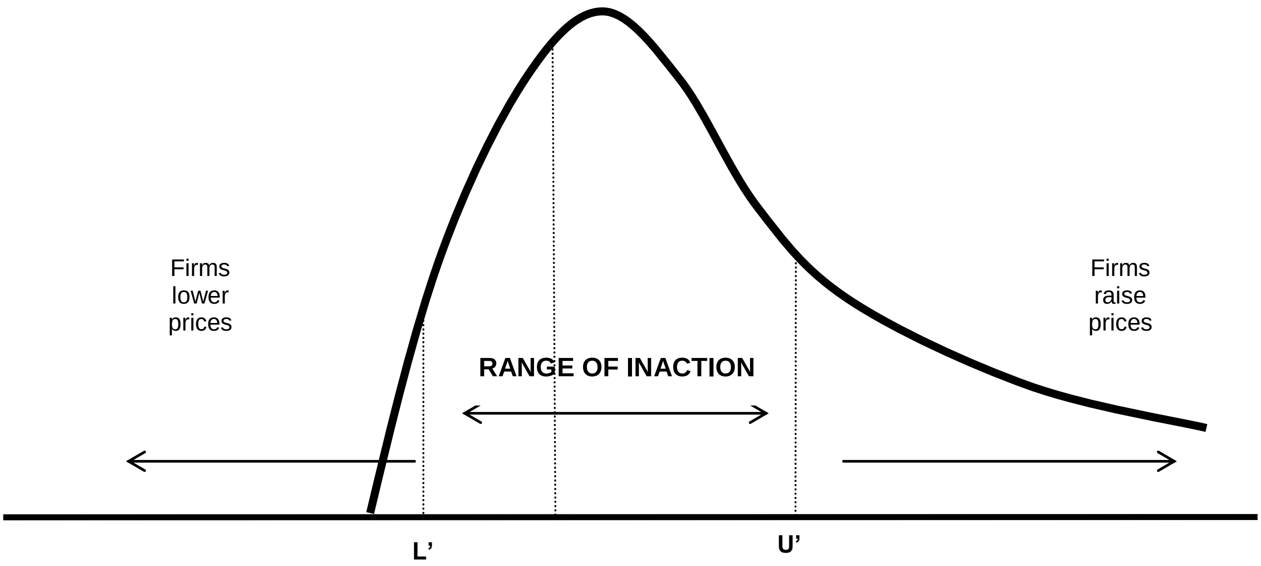
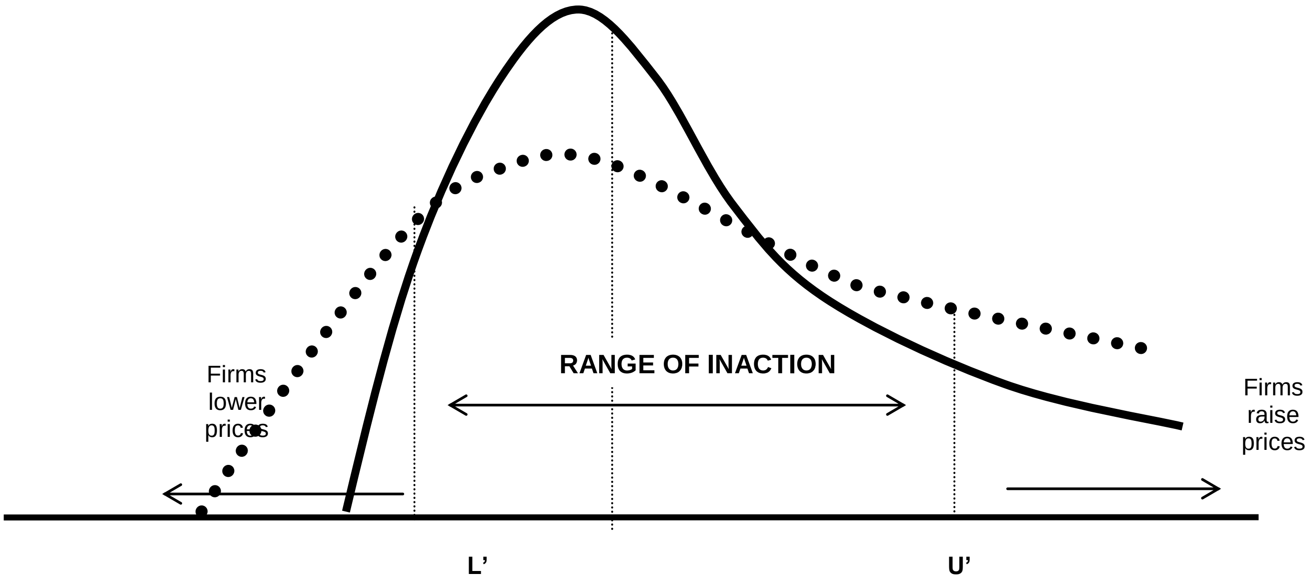


FIGURE 5.b: EFFECT OF AN INCREASE IN RPV



APPENDIX II. UNIT ROOT TEST

The specific testing procedure adopted is the Augmented Dickey-Fuller test with Akaike Information criterion used for selecting the number of lags included in the ADF regressions. Moreover, results have been checked using Schwartz criterion. By default, the maximum number of lags allowed in the tests is 12. For both countries a deterministic trend appears for the total period, but when the inflation series is divided by periods the deterministic trend disappears for the low inflation period. We show the results for the total period (results for each period are disposable from the authors upon request).

Unit root test. Spain. Total period.

Variable	Number of lags	Constant	Trend	ADF statistic	p-value
Π_t	6	yes	yes	-4.14	0.00
RPV_t	9	yes	yes	-3.69	0.02
S_t	2	yes	yes	-8.25	0.00
$S10_t$	2	yes	yes	-6.59	0.00
$S25_t$	2	no	yes	-3.12	0.01
Q_t	0	yes	yes	-6.43	0.00

Unit root test. Argentina. Total period

Variable	Number of lags	Constant	Trend	ADF statistic	p-value
Π_t	11	no	yes	-3.73	0.02
RPV_t	6	yes	no	-3.74	0.03
S_t	12	yes	no	-3.53	0.00
$S10_t$	2	yes	no	-9.32	0.00
$S25_t$	11	no	no	-4.42	0.00
Q_t	4	no	no	-4.32	0.00

APPENDIX III

CORRELATION COEFFICIENTS BETWEEN RPV_t AND S_t

	TOTAL PERIOD	HIGH INFLATION PERIOD	LOW INFLATION PERIOD
ARGENTINA	0,12	0,19	0,03
SPAIN	0,20	0,30	0,08

APPENDIX IV. TEST OF STRUCTURAL CHANGE.

We present the results for recursive residuals for estimations of equation (7), for Argentina and Spain respectively. Residuals outside the standard error bands suggest instability in the parameters of the equation.

ARGENTINA

FIGURE 1. RECURSIVE RESIDUAL

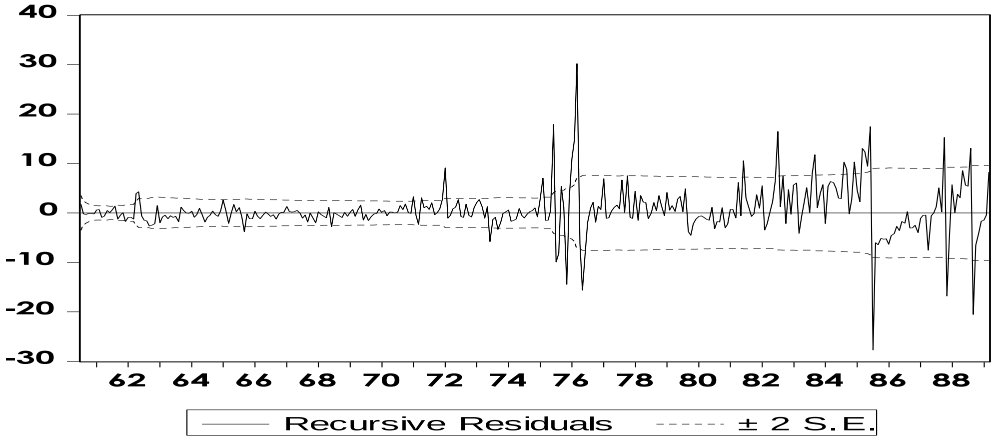


FIGURE 2. RECURSIVE COEFFICIENTS. RPV_t

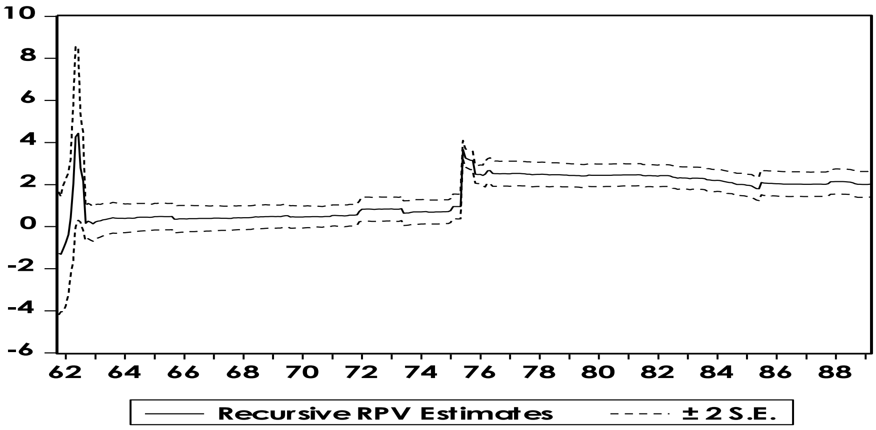
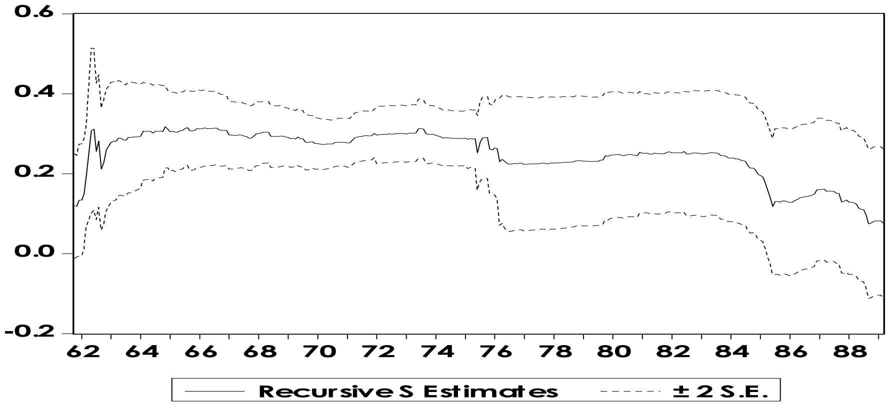


FIGURE 3. RECURSIVE COEFFICIENTS. S_t



SPAIN

FIGURE 4. RECURSIVE RESIDUALS

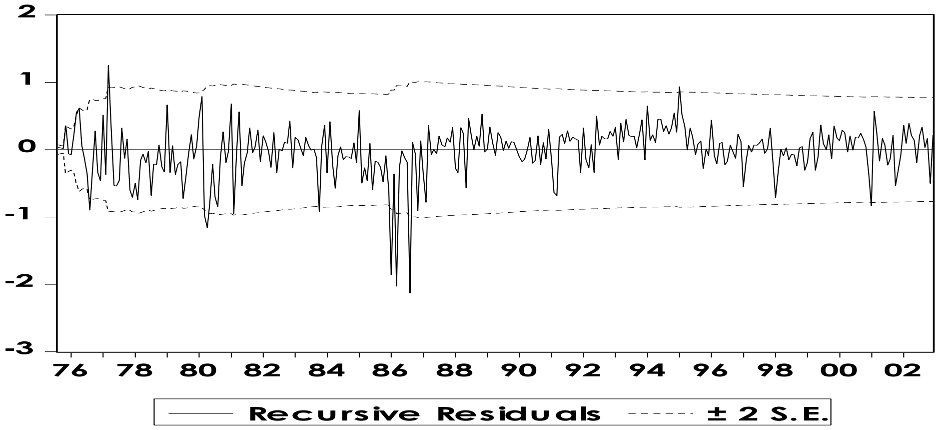


FIGURE 5. RECURSIVE COEFFICIENTS. RPV_t

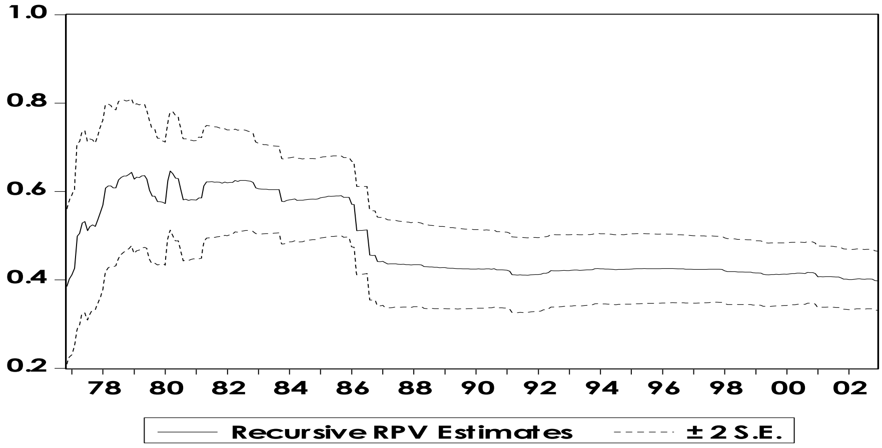
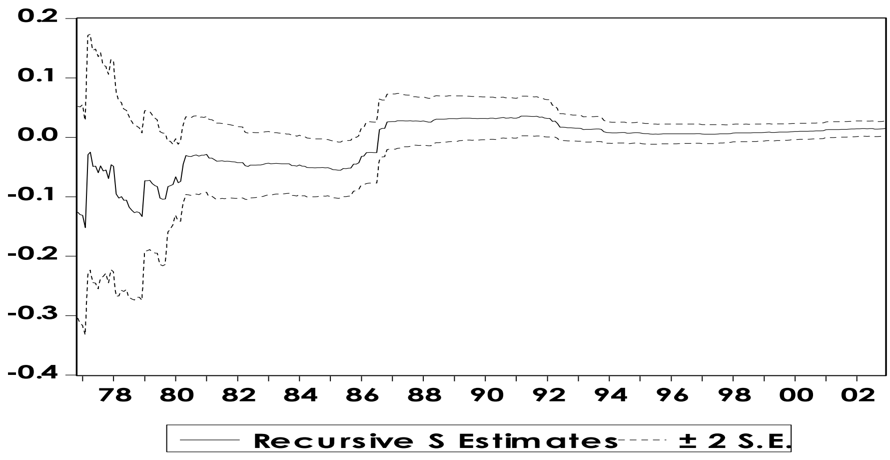


FIGURE 6. RECURSIVE COEFFICIENTS. S_t



ANEXO I .1 . ESTADÍSTICA DESCRIPTIVA DE LOS DATOS DE LAS COMUNIDADES AUTÓNOMAS

	π_{it}	S_{it}	SG_{it}	DS_{it}	DSN_{it}	A_{it}	AG_{it}	DA_{it}	DAN_{it}	K_{it}	U_{it}	IPI_{it}
ANDALUCÍA / MANCHA												
Media	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Mediana	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Desv. típica	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Varianza	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Curtosis	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Asimetría	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Mínimo	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Máximo	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]

References

Amano, R.A. and Macklem, R. (1997): "Menu Costs, Relative Prices, and Inflation: Evidence for Canada", Bank of Canada, Working Paper 97-14.

Andrews, D.W.K.(1993): "Test for Parameter Instability and Structural Change with Unknown Change Point", *Econometrica*, 61(3), pp. 821-856.

Andrews, D.W.K. and Ploberger, W.(1994): "Optimal Tests when Nuisance Parameter is Present under the Alternative", *Econometrica*, 62(6), pp. 1383-1414.

Assarsson, B. (2004): "Inflation and Relative-Price Changes in the Swedish Economy", *Sveriges Riksbank Economic Review*, 3, pp. 43-62.

Aucremanne, L., Brys, G., Hubert, M., Rousseeuw, P.J. and Struyf, A. (2002): "Inflation, Relative Prices and Nominal Rigidities", National Bank of Belgium, Working Paper 20.

Ball, L. and Mankiw, N.G. (1994): "Asymmetric Price Adjustment and Economic Fluctuations", *Economic Journal*, 104 (423), pp. 247-261.

Ball, L. and Mankiw, N.G. (1995): "Relative-Price Changes as Aggregate Supply Shocks", *Quarterly Journal of Economics*, 110 (1), pp. 161-193.

Caraballo, M.A., Dabús, C. and Usabiaga, C.(2006): "Relative Prices and Inflation: New Evidence from Different Inflationary Contexts", *Applied Economics*, 38, pp. 1931-1944.

Caraballo, M.A. and Usabiaga, C.(2004): "Inflation and Relative Prices: Empirical Evidence for the Spanish Economy", *Problems and Perspectives in Management. Chapter 1: Macroeconomic Processes and Regional Economies Management*, 3, pp. 59-71.

Caraballo, M.A. and Usabiaga, C.(2007): "The Relevance of Supply Shocks for Inflation: The Spanish Case", *Applied Economics*, Forthcoming.

Dabús, C. (2000): "Inflationary Regimes and Relative Price Variability: Evidence from Argentina", *Journal of Development Economics*, 62 (2), pp. 535-547.

Döpke, J. and Pierdzioch, C. (2003): "Inflation and Skewness of the Distribution of Relative Price Changes: Empirical Evidence for Germany", *Jahrbücher für Nationalökonomie und Statistik*, 223 (2), pp. 136-159.

Ghysels, E. (1990): "Unit Root Tests and the Statistical Pitfalls of Seasonal Adjustment: the Case of U.S. Postwar Real Gross National Product", *Journal of Business and Economic Statistics*, 8 (2), pp. 145-152.

Ghysels, E. and Perron, P. (1993): "The Effect of Seasonal Adjustment Filters on Tests for a Unit Root", *Journal of Econometrics*, 55, pp. 57-98.

Hall, S. and Yates, A. (1998): "Are There Downward Nominal Rigidities in Product Markets?", Bank of England, Working Paper Series, 80.

Hansen, B. (1997): "Approximate asymptotic p-values for structural change tests", *Journal of Business and Economic Statistics*, 15, pp. 60-67.

Leijonhufvud, A. (1990): *Extreme Monetary Instability: High Inflation*, Lecture, University of Hoheneim (Germany), mimeo.

Lourenço, R.D.A. and Gruen, D. (1995): "Price Stickiness and Inflation", Reserve Bank of Australia, Research Discussion Paper 9502.

Mills, F. (1927): *The Behaviour of Prices*, Arno, New York.

Parks, R.W. (1978): "Inflation and Relative Price Variability", *Journal of Political Economy*, 86 (1), pp. 79-95.

Rattai, A.(2004): "Inflation and Relative Price Asymmetry", European Central Bank, Working Paper Series, nº 301.

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