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The Relevance of Supply Shocks for Inflation: The Spanish Case

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RESUMEN

Este trabajo analiza los efectos de las perturbaciones de oferta sobre la tasa de inflación española. La metodología aplicada está basada en Ball y Mankiw (1995). Estos autores asumen que una buena "proxy" para las perturbaciones de oferta la constituye el tercer momento de la distribución de los cambios de precios, y muestran que las rigideces nominales implican una relación positiva entre la inflación y la asimetría, que es amplificada por la varianza de la distribución. Los principales datos usados son los índices de precios de consumo mensuales de cada región, desagregados en 57 categorías, para el periodo 1993-2005. Estimamos la relación entre la inflación media y los momentos superiores de la distribución, incluyendo numerosas variables de control. El análisis ha sido desarrollado de dos formas: en primer lugar cada región es analizada separadamente y en segundo lugar hemos usado técnicas de datos de panel para contrastar la homogeneidad entre las regiones. Nuestros resultados apuntan que las regiones españolas muestran un patrón común respecto a las rigideces nominales detectadas en el comportamiento de la inflación, y que la economía española es vulnerable a las perturbaciones de oferta.

Palabras clave:
Inflación, rigideces nominales, asimetría, perturbaciones de oferta, regiones españolas.

ABSTRACT

This paper analyses the effects of supply shocks on the Spanish inflation rate. The methodology applied is based on Ball and Mankiw (1995). These authors assume that a good proxy for supply shocks is the third moment of the distribution of price changes, and show that nominal rigidities imply a positive relation between inflation and skewness, that is magnified by the variance of the distribution. The main data used are the monthly consumer price indexes of each region, disaggregated in 57 categories, for the 1993-2005 period. We estimate the relation between mean inflation and the higher moments of the distribution, including several control variables. The analysis has been carried out in two ways: firstly, each region is analysed separately and, secondly, we have used panel data techniques in order to test the homogeneity across regions. Our results point out that Spanish regions show a common pattern with regard to the nominal rigidities detected, and that the Spanish economy is vulnerable to supply shocks.

Keywords: Inflation, nominal rigidities, skewness, supply shocks, Spanish regions
JEL classification: E31

1. Introduction[1]

The idea of the present contribution is based on several factors: i) Spain is a country characterised by a persistent moderate inflation differential with the core EU countries –see European Central Bank (2003). ii) The figures of Spanish inflation have slightly increased in recent years and the Spanish Government faces problems to control inflation –see Bank of Spain (2006). iii) The irregular evolution of oil prices in recent years, with several adverse supply shocks, deserves a lot of international attention –see Kilian (2005). Our paper tries to shed some light jointly on these factors, from the Spanish perspective, proposing some explanations mainly based on the use of Ball and Mankiw's (1994, 1995) approach. In order to implement panel data techniques and provide additional information at a regional level, we pay special attention to the Spanish regional inflation data, although we also include in our analysis several control variables. The main contributions of this paper, in comparison with previous ones in this area for the Spanish economy –Caraballo and Usabiaga (1994a, 1994b), Caraballo and Dabús (2005)–, are the following: we work with a higher degree of disaggregation in the data, a very important feature in this kind of literature based on price changes distribution functions; we extend and update the period of analysis; and we incorporate as control variables the main economic variables related to this topic available for the Spanish economy with a monthly frequency.

Empirical evidence shows that inflation and higher moments of the distribution of relative prices are positively correlated, against the theoretical predictions of the flexible price model. Ball and Mankiw (1994,1995) show that inflation is mainly influenced by skewness, arguing 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; effect that can be magnified by the standard deviation of the distribution, denoted as relative price variability (RPV) in this strand of the literature. This paper tries to check if the skewness-inflation relation holds for Spain and if the behaviour of Spanish regions is homogeneous in this respect. The analysis of such relation can be relevant in the sense that these authors show that skewness is a

[1] The authors would like to thank the suggestions received from Carlos Dabús and Diego Romero-Avila, participants at the EEFS 5th Annual Conference (Crete, Greece, 2006), an anonymous referee of ERSA 46th Congress (Volos, Greece, 2006), and an anonymous referee of Centro de Estudios Andaluces. The authors also acknowledge financial support from Junta de Andalucía: Centro de

proxy for supply shocks, and therefore that relation is explaining how sensitive the inflation is when a supply shock affects the economy, and if a supply shock affects to the same extent all regions.

Positive inflation-skewness and inflation-RPV relations are supported by the data, but results are not conclusive about which relation is stronger. On one hand, for periods with an annual inflation rate lower (higher) than 4%-5%, the inflation-skewness relation is stronger (weaker) than the inflation-RPV one –see Lourenco and Gruen (1995) for Australia, Ball and Mankiw (1995) for the United States, Amano and Macklem (1997) for Canada, Aucremanne *et al.* (2002) for Belgium and Caraballo and Usabiaga (2004a,b) for Spain, among others.[2] 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 relation than the inflation-RPV one, or Döpke and Pierdzioch (2003), for the 1969-2000 period in Germany, find that both relations are positive, but none of them is stronger than the other. 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%. Finally, Caraballo and Dabús (2005) focus on Spain and Argentina, concluding that the predictions of menu costs models 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 an approach is suitable at high inflation levels.

This mixed evidence can be due to different reasons, and specially to the fact that the relation between inflation and the higher moments of the distribution of changes in relative prices is very sensitive to changes in the features of mean inflation. Generally in low inflation countries both variables are significant but depending on the trend of inflation a relation can be more significant than the other.[3]

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[2] However, as an exception to this general result, Assarsson (2004) finds that in Sweden both relations are positive and strong, and neither of them is stronger than the other.

[3] See Caraballo and Dabús (2005) for further details.

The rest of the paper is organised as follows. Section 2 presents the main data and variables. In section 3 we develop a preliminary analysis for the 17 Spanish regions. Section 4 performs a panel data analysis. In section 5 several control variables are included, and section 6 concludes.

2. Main data and variables

Our analysis refers to the 1993.02-2005.12 period. We are aware of the shortness of this sample period (13 years) in comparison with other studies, but it is not possible to extend it, due to the important data requirements of our analytical methodology, with a high degree of disaggregation in the data, as well as the use of several control variables. Only the period considered fulfils all this data matching. However, we have to take into account that the data are monthly, a frequency which is not commonly used in the literature, and consequently we get 155 observations of each series. Our sample period can be clearly divided into two subperiods. The first one goes from 1993.02 to 1998.12, and is characterised by a negative trend inflation, and a mean monthly inflation rate around 0.28%. The second one is the 1999.01-2005.12 period, in which no trend inflation is found and presents a mean monthly inflation rate around 0.26%.

The main data used are the series of monthly change rates of consumer price indexes, disaggregated by goods and services (57 categories), for the 17 Spanish regions elaborated by the Instituto Nacional de Estadística (INE). The weight of each subgroup offered by the INE is defined as the proportion of expense made on that article in relation to total expenditure made by households. The weight is kept constant by the INE along the 1993.02-2001.12 period, but since 2002 there has been a change in the methodology and the weights change every year. This fact is taken into account when the moments of the distribution of inflation are calculated. Another change in the methodology is the introduction of sales in the index. In order to avoid the problems caused by this change, we remove the seasonal component using the TRAMO-SEATS method.

As control variables we use the rate of unemployment, the industrial production index, the general retail trade index, the shopping mall retail trade index, the oil prices and the industrial price index. We provide information about them in the corresponding section.

As far as the construction of the main variables is concerned, we use the second and third cross-sectional moment of the distribution of price changes. The expressions of the standard deviation for each region (*RPV_{jt}*) and the skewness for each region (*S_{jt}*) are as follows:

$$RPV_{jt} = \left[\sum_{i=1}^{n} w_{ij} \left(\pi_{ijt} - \pi_{jt} \right)^2 \right]^{0,5}; \quad S_{jt} = \frac{\sum_{i=1}^{n} w_{ij} \left[\pi_{ijt} - \pi_{jt} \right]^3}{\left(S_{jt} \right)^3}$$

where π refers to inflation rate, i to goods, j to regions and t to time periods. Therefore, π_t: Spanish inflation in period t; π_{jt}: inflation of region j in period t; π_{ijt}: inflation of subgroup i in region j in period t; and w_{ij} is the weight of each subgroup i and region j used by INE.

3. Inflation, RPV and skewness: preliminary analysis on a regional basis

In this section a preliminary region-by-region analysis is perfomed. In order to implement it, we run the following regression for each region

$$\pi_{jt} = \alpha_j + \beta_1^j \pi_{j,t-1} + \beta_2^j S_{jt} + \beta_3^j RPV_{jt} + \varepsilon_{jt} \qquad [1]$$

The lagged inflation term is included in order to capture the persistence of the series.

Before running the regressions we have checked the stationarity of the series.[4] For the 17 regions inflation presents a negative deterministic trend for the 1993-1998 period, but there is no trend in the 1999-2005 one. This feature of inflation is included in the regressions.

The regressions are estimated by ordinary least squares (OLS).[5] As usual, the p-value of the t-statistic (in brackets in the tables) is corrected for heteroscedasticity by means of the White method. We show the results for each subperiod (Tables 1 and 2) and for the whole period (Table 3).

[4] In the Appendix we present the results for a common unit root –Breitung (2000) and Levin *et al.* (2002)-, and the general result is that it does not exist. Results of individual unit root tests are available from the authors upon request. The specific testing procedure adopted is the Augmented Dickey-Fuller (ADF) test with the Schwartz information criterion used to select the number of lags included in the ADF regressions. By default, the maximum number of lags allowed in the tests is 12. In the Appendix we also show the summary statistics for inflation, RPV and skewness.

Table 1: Regional analysis (1993-1998)

Region	*Constant*	$\pi_{j,t-1}$	$S_{j,t}$	$RPV_{j,t}$	*Trend*	*Adjusted R^2*
Andalucía	0.15 (0.00)	0.47 (0.00)	0.03 (0.00)	0.05 (0.02)	-0.002 (0.00)	0.71
Aragón	0.29 (0.00)	0.40 (0.00)	0.02 (0.00)	-0.01 (0.48)	-0.003 (0.00)	0.75
Asturias	0.14 (0.35)	0.50 (0.00)	0.01 (0.00)	0.04 (0.57)	-0.002 (0.00)	0.75
Baleares	0.19 (0.00)	0.26 (0.02)	0.004 (0.02)	0.08 (0.01)	-0.002 (0.00)	0.83
Canarias	0.32 (0.00)	0.01 (0.83)	0.04 (0.00)	0.04 (0.09)	-0.003 (0.00)	0.51
Cantabria	0.50 (0.00)	-0.33 (0.00)	0.00 (0.95)	0.02 (0.35)	-0.004 (0.00)	0.47
Cataluña	0.30 (0.00)	0.27 (0.01)	0.01 (0.00)	0.007 (0.86)	-0.003 (0.00)	0.60
Castilla-León	0.30 (0.00)	0.19 (0.08)	0.005 (0.01)	0.02 (0.44)	-0.004 (0.00)	0.64
Castilla-La Mancha	-0.27 (0.07)	0.43 (0.00)	0.01 (0.00)	0.27 (0.00)	-0.000 (0.59)	0.76
Extremadura	0.08 (0.24)	0.44 (0.00)	0.03 (0.00)	0.09 (0.00)	-0.002 (0.00)	0.77
Galicia	0.28 (0.00)	0.36 (0.00)	0.01 (0.00)	0.00 (0.84)	-0.003 (0.00)	0.78
Madrid	0.28 (0.00)	0.25 (0.02)	0.01 (0.00)	0.01 (0.70)	-0.003 (0.00)	0.61
Murcia	0.23 (0.00)	-0.37 (0.00)	0.01 (0.00)	0.18 (0.00)	-0.004 (0.00)	0.63
Navarra	0.53 (0.01)	0.23 (0.04)	0.007 (0.08)	-0.06 (0.56)	-0.005 (0.00)	0.63
País Vasco	0.40 (0.00)	-0.18 (0.14)	0.00 (0.21)	0.07 (0.05)	-0.004 (0.00)	0.61
La Rioja	0.26 (0.17)	-0.15 (0.22)	0.00 (0.87)	0.12 (0.17)	-0.003 (0.00)	0.35
Valencia	0.05 (0.52)	0.35 (0.00)	0.03 (0.01)	0.12 (0.01)	-0.001 (0.00)	0.60

[5] As well known, if the lagged endogenous variable is not correlated with the error term, the validity of the OLS estimator holds. To prove that there is no correlation, we have estimated the model with OLS and verified that there is not autocorrelation in the residuals.

Table 2: Regional analysis (1999-2005)

Region	*Constant*	$\pi_{j,t-1}$	$S_{j,t}$	$RPV_{j,t}$	*Adjusted* R^2
Andalucía	0.14 (0.00)	0.39 (0.00)	0.03 (0.00)	0.003 (0.89)	0.26
Aragón	0.08 (0.28)	0.38 (0.00)	0.02 (0.00)	0.04 (0.30)	0.23
Asturias	0.11 (0.07)	0.42 (0.00)	0.01 (0.00)	0.01 (0.66)	0.24
Baleares	0.17 (0.00)	0.56 (0.00)	0.00 (0.00)	-0.04 (0.06)	0.47
Canarias	0.28 (0.00)	0.26 (0.00)	0.02 (0.00)	-0.10 (0.06)	0.16
Cantabria	0.35 (0.00)	-0.52 (0.00)	0.01 (0.19)	0.01 (0.46)	0.24
Cataluña	0.20 (0.00)	0.21 (0.04)	0.01 (0.00)	0.01 (0.75)	0.09
Castilla-León	0.12 (0.09)	0.37 (0.00)	0.009 (0.02)	0.02 (0.54)	0.18
Castilla-La Mancha	0.09 (0.40)	0.37 (0.00)	0.02 (0.00)	0.03 (0.58)	0.22
Extremadura	0.05 (0.29)	0.42 (0.00)	0.03 (0.00)	0.03 (0.14)	0.27
Galicia	0.12 (0.02)	0.40 (0.00)	0.009 (0.02)	0.01 (0.50)	0.20
Madrid	0.14 (0.08)	0.21 (0.03)	0.01 (0.00)	0.03 (0.48)	0.18
Murcia	0.25 (0.00)	-0.02 (0.78)	0.01 (0.00)	0.02 (0.44)	0.10
Navarra	0.17 (0.03)	0.20 (0.05)	0.01 (0.00)	0.02 (0.60)	0.09
País Vasco	0.27 (0.00)	-0.01 (0.90)	0.006 (0.02)	-0.001 (0.92)	0.02
La Rioja	0.31 (0.00)	0.01 (0.88)	0.009 (0.04)	-0.001 (0.44)	0.02
Valencia	0.14 (0.03)	0.44 (0.00)	0.02 (0.16)	0.001 (0.97)	0.20

Table 3: Regional analysis (1993-2005)

Region	*Constant*	$\pi_{j,t-1}$	$S_{j,t}$	$RPV_{j,t}$	*Trend (93-98)*	*Adjusted R^2*
Andalucía	0.11 (0.00)	0.52 (0.00)	0.03 (0.00)	0.03 (0.04)	-0.001 (0.02)	0.47
Aragón	0.10 (0.03)	0.46 (0.00)	0.01 (0.00)	0.04 (0.06)	-0.001 (0.01)	0.43
Asturias	0.12 (0.01)	0.49 (0.00)	0.01 (0.00)	0.03 (0.09)	-0.001 (0.00)	0.46
Baleares	0.07 (0.35)	0.71 (0.00)	0.005 (0.14)	0.02 (0.60)	-0.000 (0.34)	0.70
Canarias	0.30 (0.00)	0.11 (0.02)	0.03 (0.00)	0.01 (0.35)	-0.002 (0.00)	0.36
Cantabria	0.43 (0.00)	-0.33 (0.00)	0.004 (0.53)	0.04 (0.06)	-0.001 (0.00)	0.22
Cataluña	0.17 (0.00)	0.34 (0.00)	0.01 (0.00)	0.05 (0.06)	-0.001 (0.00)	0.28
Castilla-León	0.10 (0.06)	0.45 (0.00)	0.005 (0.01)	0.05 (0.06)	-0.001 (0.02)	0.35
Castilla-La Mancha	-0.01 (0.77)	0.44 (0.00)	0.01 (0.00)	0.13 (0.00)	-0.001 (0.01)	0.47
Extremadura	0.04 (0.39)	0.52 (0.00)	0.03 (0.00)	0.07 (0.00)	-0.001 (0.01)	0.56
Galicia	0.14 (0.00)	0.48 (0.00)	0.01 (0.00)	0.03 (0.09)	-0.001 (0.01)	0.46
Madrid	0.08 (0.15)	0.33 (0.00)	0.01 (0.00)	0.09 (0.00)	-0.001 (0.01)	0.32
Murcia	0.22 (0.00)	-0.01 (0.88)	0.01 (0.00)	0.08 (0.00)	-0.001 (0.00)	0.26
Navarra	0.20 (0.00)	0.29 (0.00)	0.01 (0.00)	0.07 (0.01)	-0.002 (0.00)	0.34
País Vasco	0.27 (0.00)	0.12 (0.30)	0.004 (0.15)	0.04 (0.05)	-0.001 (0.00)	0.32
La Rioja	0.32 (0.00)	0.05 (0.52)	0.002 (0.43)	0.02 (0.17)	-0.001 (0.00)	0.12
Valencia	0.08 (0.07)	0.48 (0.00)	0.03 (0.01)	0.05 (0.05)	-0.000 (0.07)	0.37

As it can be seen from the tables, skewness is significant in 13 regions for the 1993-1998 subperiod, in 15 regions for the 1999-2005 subperiod, and in 13 regions for the whole period, and its coefficient remains unchanged for the different sample periods. However, the behaviour of RPV is not so homogeneous across periods, and tables show that it is significant in 7 regions for the first subperiod, it is not significant in any region for the second subperiod, and it is significant in 8 regions for the whole period (in 6 of them it was significant in the first period as well), and its coefficient varies considerably among sample periods. It is also interesting to point out the remarkable changes in the adjusted R^2 depending on the period considered; the existence of a trend can be the key to this result. Finally, according to the coefficients on lagged inflation, it is clear that inflation shows persistence.

In conclusion, these results seem to confirm the predictions of Ball and Mankiw's model regarding the relevance of skewness, and show that RPV is more sensitive to changes in the inflation regime (the two sample periods in our analysis) than skewness.

4. Panel data analysis

In this section, we perform panel data analysis in order to control for the possibility that regional inflation may be affected by common factors, which lead to strong correlation across regional inflation rates. In order to implement it, we attend to the following estimation:

$$\pi_{jt} = \alpha_j + \beta_1 \pi_{j,t-1} + \beta_2 S_{jt} + \beta_3 RPV_{jt} + \varepsilon_{jt} \qquad j = 1...17 \qquad [2]$$

where α_j is a fixed effect for each region. As it can be seen from equation (2), lagged inflation is correlated with the fixed effects. Therefore, within estimators will be biased and inconsistent. This problem cannot be avoided estimating the model in first differences, because although the fixed effect is wiped out, the first-differenced variables are correlated with the random component of the error term. The degrees of inconsistency and bias depend on T; only if $T \rightarrow \infty$ the within estimator is unbiased and consistent.[6] In other words, for a typical panel where N is large in relation to T (T is usually fixed), and where the enlargement of the sample always refers to N and not to T, instrumental variable estimation is required in order to get consistent and unbiased estimators. However, this is not our case because N (regions) is fixed, T is very large in relation to N, and the enlargement of the sample can be referred only to T. Despite the discussion about the number of periods required to get an unbiased and consistent within estimator would deserve a lot of attention, we have considered that the features of our sample allow us to use within estimators.

Now, we estimate (2) for the two subperiods[7] and the total period –see Tables 4, 5 and 6, first column- and we perform a fixed effect test[8] for the null hypothesis $\alpha_j = \alpha$, for all j = 1…17. The test statistic is distributed under the null hypothesis as a $F_{16,1169}$ and its value is 1.53 for the 1993-1998 period, as a $F_{16,1390}$ and its value is 1.15 for the 1999-2005 period, and as a $F_{16,2597}$ and its value is 1.04 for the total

[6] See Baltagi (1995, p. 126).
[7] In order to reinforce the validity of the division in the sample period that we use in our analysis we have implemented a Chow test. The critical value of this test is 3.02 at 1% (the F statistic is 27.49) so we reject the null hypothesis of lack of a break in 1998:12.

period. Therefore, the null hypothesis that α_j are equal cannot be rejected in any case, so we estimate (3) –see Tables 4, 5 and 6, second column–:

$$\pi_{jt} = \alpha + \beta_1 \pi_{j,t-1} + \beta_2 S_{jt} + \beta_3 RPV_{jt} + \varepsilon_{jt} \qquad j = 1...17 \qquad [3]$$

Finally, the instrumental variable estimation suggested by Anderson and Hsiao (1981) is applied –see Tables 4, 5 and 6, third column. We estimate the model in first differences, in order to get rid of the hypothetical individual effects:

$$\pi_{jt} - \pi_{j,t-1} = \beta_1(\pi_{j,t-1} - \pi_{j,t-2}) + \beta_2(S_{jt} - S_{j,t-1}) + \beta_3(RPV_{jt} - RPV_{j,t-1}) + (\varepsilon_{jt} - \varepsilon_{j,t-1}) \quad [4]$$

As ($\pi_{j,t-1}$- $\pi_{j,t-2}$) is correlated with the new error term, we run an instrumental variable estimation using the inflation variable in levels $\pi_{j,t-2}$ as instrument; for the rest of the variables we do not define any instruments.

Table 4: Panel data analysis (1993-1998), with negative trend

Variable	Fixed Effect	OLS	Anderson-Hsiao
$\pi_{j,t-1}$	0.12 (0.00)	0.25 (0.00)	0.12 (0.00)
$S_{j,t}$	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)
$RPV_{j,t}$	0.05 (0.00)	0.04 (0.00)	0.006 (0.03)
Adjusted R^2	0.58	0.57	-

Table 5: Panel data analysis (1999-2005)

Variable	Fixed Effect	OLS	Anderson-Hsiao
$\pi_{j,t-1}$	0.28 (0.00)	0.29 (0.00)	0.57 (0.00)
$S_{j,t}$	0.01 (0.00)	0.01 (0.00)	0.006 (0.00)
$RPV_{j,t}$	0.01 (0.25)	0.01 (0.05)	0.00 (0.7)
Adjusted R^2	0.17	0.16	-

Table 6: Panel data analysis (1993-2005), with negative trend (1993-1998)

Variable	Fixed Effect	OLS	Anderson-Hsiao
$\pi_{j,t-1}$	0.35 (0.00)	0.36 (0.00)	0.39 (0.00)
$S_{j,t}$	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)
$RPV_{j,t}$	0.04 (0.00)	0.04 (0.00)	0.002 (0.23)
Adjusted R^2	0.32	0.34	-

[8] See Baltagi (1995, p. 12).

As it can be observed, there are not remarkable changes with respect to skewness for the three methods of estimation reported, and its coefficient seems to be stable across periods. But this does not hold for RPV and the constant term in the OLS estimation.[9] These results lead us to introduce in the estimation for the total period both a dummy variable (*D93-98*) and a slope dummy (*D93-98*RPV*$_{j,t}$) for the 1993-1998 period, in order to capture the change in the constant and in the coefficient of RPV respectively. Moreover, we have checked that a slope dummy for skewness is not significant. We have run the regression with fixed effects for the whole period, and again the Hausman test leads us to reject the fixed effects, so finally we present the results for the OLS estimation in Table 7:

Table 7: Panel data analysis with dummies (1993-2005). OLS

Constant	$\pi_{j,t}$	$S_{j,t}$	$RPV_{j,t}$	$D93\text{-}98{*}RPV_{j,t}$	***D93-98***	***Trend (93-98)***	***Adjusted*** R^2
0.38 (0.00)	0.28 (0.00)	0.01 (0.00)	0.01 (0.02)	0.02 (0.01)	-0.14 (0.00)	-0.003 (0.00)	0.38

Summarising, our results show a homogeneous behaviour both across regions and periods regarding skewness, which can be revealing the vulnerability of the Spanish economy to supply shocks. As far as RPV is concerned, the predictions of Ball and Mankiw (1995) for no trend inflation are confirmed, given that it is not significant for the 1999-2005 period in any region. This variable appears to be heavily affected by the behaviour of the inflation rate.

5. Introduction of control variables

As it was mentioned in the introduction, in this section we include several control variables. The idea embedded in the inclusion of these variables is twofold: i) to check the robustness of the aforementioned relation between mean inflation on the one hand and skewness and RPV on the other –Ball and Mankiw's approach–; ii) to get some preliminary empirical evidence on the relevance of different macroeconomic relations for the Spanish economy.

Although we have introduced many control variables, we would have liked to include even a higher number, but the monthly frequency imposed an important shortcoming (think for instance in variables related to fiscal policy). With the

[9] Results for the constant term and trend are not included in the tables. They are available from the

exception of the regional unemployment rate, these variables are provided at a national level, because they are nor available, homogeneously, at a regional level. The data sources for our control variables are the following[10]: i) Unemployment rates: Instituto Nacional de Empleo (INEM). ii) Industrial production index: INE (Base year 2000). iii) General retail trade index and shopping mall retail trade index: INE. iv) Interest rate: Bank of Spain. %. 3 months-deposits. Interbank mean rate. v) Oil prices: Reuters. North Sea Brent. Dollars/barrel. vi) Industrial price index: INE

As many of our control variables are clearly related among them, we have opted for including them in the basic expression –see panel data analysis, Table 7– separately, in order to avoid multicollinearity problems, as well as to isolate its effect on mean inflation. The main results of this analysis appear in Table 8.

authors upon request.

[10] A more detailed information about these variables and data sources is available from the authors upon request.

Table 8: Introduction of control variables: unemployment, industrial production, retail trade, interest rates, oil prices and industrial price index

Constant	0.38 (0.00)	0.38 (0.00)	0.39 (0.00)	0.38 (0.00)	0.37 (0.00)	0.44 (0.00)	0.35 (0.00)	0.35 (0.00)
$\pi_{j,t-1}$	0.28 (0.00)	0.28 (0.00)	0.27 (0.00)	0.27 (0.00)	0.27 (0.00)	0.27 (0.00)	0.31 (0.00)	0.26 (0.00)
$S_{j,t}$	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)	0.01 (0.00)
$RPV_{j,t}$	0.01 (0.02)	0.01 (0.02)	0.01 (0.02)	0.01 (0.02)	0.01 (0.00)	0.01 (0.04)	0.01 (0.03)	0.01 (0.09)
$D93\text{-}98{*}RPV_{j,t}$	0.02 (0.01)	0.02 (0.01)	0.02 (0.02)	0.02 (0.00)	0.02 (0.06)	0.01 (0.17)	0.02 (0.01)	0.02 (0.01)
D93-98	-0.14 (0.00)	-0.14 (0.00)	-0.14 (0.00)	-0.14 (0.00)	-0.12 (0.00)	-0.14 (0.00)	-0.12 (0.00)	-0.12 (0.00)
Trend (93-98)	-0.003 (0.00)	-0.003 (0.00)	-0.003 (0.00)	-0.003 (0.00)	-0.003 (0.00)	-0.004 (0.00)	-0.002 (0.00)	-0.002 (0.00)
Spanish cyclical unemployment	-0.006 (0.50)							
Regional cyclical unemployment		-0.002 (0.69)						
Cyclical industrial production index			0.002 (0.01)					
Cyclical general retail trade index				0.01 (0.00)				
Cyclical shopping mall retail trade index					0.008 (0.00)			
Lagged (t-15) change in interest rates						-0.11 (0.02)		
Change in oil prices							0.30 (0.00)	
Lagged (t-3) change in industrial price index								0.06 (0.00)
Adjusted R^2	0.38	0.38	0.39	0.39	0.39	0.34	0.42	0.39

Firstly, we have considered the unemployment variable. In this sense, we have worked both with the national and regional unemployment series. Obviously, the inclusion of this variable tries to capture the Phillips curve relation (in few words, the negative relation between inflation and unemployment). In order to obtain an accurate relation, we have used the cyclical unemployment, defined as the difference between the seasonally adjusted unemployment rate and the same variable filtered following Hodrick-Prescott's (1980) method with the standard smoothing parameter for monthly data. Once we introduce the cyclical aggregate Spanish unemployment rate as a control variable, we obtain a small negative coefficient, which is not significant, and the adjusted R^2 does not change. The same holds for the cyclical regional unemployment rate. In other words, the evidence in favour of the Phillips curve relation is not conclusive at all. This conclusion accords with many other contributions for the Spanish economy –see the survey of Gómez and Usabiaga (2001).

Secondly, we focus on the industrial production index. In this case the underlying relation would be of the aggregate supply type (in few words, the positive relation between inflation and production). Following the methodology previously applied to unemployment, we implement the analysis using the cyclical industrial production index. The results obtained in this case are similar in spirit to those for unemployment. The coefficient is positive and significant, but small, and the adjusted R^2 hardly changes. To sum up, the evidence in favour of an aggregate supply relation is very weak.

Thirdly, we have considered two well known demand indicators, which are related to retail trade: the general retail trade index and the shopping mall retail trade index. In our analysis we use the cyclical indexes following the aforementioned methodology (seasonally adjusted variables and Hodrick-Prescott). Despite the common use of both indicators to capture the demand strength, the results are very similar to the case of the industrial production index. In conclusion, the response of inflation to these two demand proxies is not noteworthy.

Fourthly, we have included the interest rate variable in our analysis. More precisely, we have considered the monthly change in interest rates. The idea is to capture the incidence of the management of monetary policy (reflected in the behaviour of interest rates) on mean inflation. From the seminal papers by Friedman –see Friedman (1968)–, and the subsequent more technical contributions on this topic (SVAR analysis and so on), it is well known that the maximum effect of interest rate policy can be very delayed, mainly due to the relevant "external" lag of this kind of policy –this is the opposite case of fiscal policy, in which the "internal" lag is the predominant one. Several studies on the main effects of monetary policy on output find a lag of even two or three years –see for instance Bryant *et al.* (1988). In this sense, in principle we would expect that the increase of interest rates (restrictive monetary policy) would help to control inflation, although with a considerable lag.

Having these ideas in mind, in our analysis we introduce the interest rate change with different increasing lags, and only with fifteen months we get a negative and significant coefficient for that variable. However, due to the data requirements of the inclusion of this relevant lag, the adjusted R^2 is lower than in the previous cases. In other words, the explanatory power of monetary policy is not very convincing in this respect.

Fifthly, we have to note that in the title of our work, as well as in the underlying idea in Ball and Mankiw's model, supply shocks are the key. We have to highlight that Ball and Mankiw (1995) emphasise the importance of the skewness of the price changes distribution as a proxy for supply shocks. The main supply shock considered in the related literature is the change in oil prices, so we include the monthly change in oil prices as an additional control variable in our analysis. In principle, we expect an increase in oil prices (an adverse supply shock) to cause an inflation upturn. In that direction, our results show that, even contemporaneously, the coefficient on the change in oil prices is clearly positive and significant, and the adjusted R^2 is the highest in comparison with those obtained with other control variables.

Finally, it is well known that the industrial price index is commonly used to anticipate the behaviour of the consumer price index. Several studies have tried to calibrate or estimate the exact lag between both indexes. In general, we can conclude from the review of the literature that the industrial price index anticipates the consumer price index in just a few months –see Quilis (1999) for the Spanish case. That explains why we have included the monthly change in the seasonally adjusted industrial price index with a lag of three months. Although we get a positive and significant coefficient, it can be observed that it is small and that the adjusted R^2 remains almost unchanged.

In this section, we conclude that oil prices seem to be the most important control variable in our analysis, highlighting the role of the supply shocks in comparison with the demand shocks in this field. The evidence presented in this section also reinforces the relevance of the methodology developed by Ball and Mankiw for the analysis of the Spanish inflation, because the coefficients and the adjusted R^2 of the expression imported from the previous section –Table 7– are almost unaffected by the introduction of the different types of control variables. To sum up, despite the inclusion of the control variables, the lagged inflation and the higher moments of the distribution of price changes maintain their relevance in the explanation of mean inflation.

6. Concluding remarks

In this paper we try to contribute to a better understanding of Spanish inflation mainly by means of the application of Ball and Mankiw's (1995) approach. These authors assume that the third moment (skewness) of the distribution of changes in relative prices is a good proxy for supply shocks, and show that, for no trend inflation regimes, nominal rigidities imply a positive relation between inflation and skewness, which is magnified by the variance of the distribution.

The main data used in our analysis are the monthly consumer price indexes of each region, disaggregated in 57 categories, for the 1993:02-2005:12 period, given that they fulfil the features required to apply the aforementioned methodology. On the one hand, we estimate the basic relation between inflation and the higher moments of the distribution. This analysis has been carried out in two ways: firstly, each region is analysed separately and, secondly, we use panel data techniques to test the homogeneity across regions. On the other hand, on the basis of the aforementioned panel data analysis, we add several control variables (unemployment, industrial production, retail trade, interest rate, oil price and industrial price) separately in order to avoid multicollinearity and isolate their effects.

The results from our regional analysis seem to confirm the predictions of Ball and Mankiw's (1995) model regarding the relevance of skewness, and show that the RPV is more sensitive than skewness to changes in the inflation regime. Our panel data analysis shows a homogeneous behaviour both across regions and periods regarding the importance of skewnes. As far as the RPV is concerned, the predictions of Ball and Mankiw's (1995) model for no trend inflation regimes are confirmed, given that this variable is never significant for the 1999-2005 period (neither in the regional analysis nor in the panel one). We can conclude that the relevance of skewness is very robust, whereas the role of the RPV appears to be heavily affected by the inflation context.

As it was previously mentioned, the relevance of skewness in our analysis can be interpreted as a sign of the vulnerability of the Spanish economy to supply shocks.[11] Along similar lines, in our analysis with several control variables, we conclude that oil prices seem to be the most important control variable, highlighting the importance of supply shocks in comparison with demand shocks. This conclusion could explain the great attention paid internationally to the evolution of oil prices and other related factors. These results also open a new line of explanation of the Spanish inflation differential with respect to the euro area, additional to the traditional explanations based on inertial elements associated with price and wage rigidities or on dual inflation (the prices of non-tradable goods are more rigid than prices in sectors exposed to international competition), or explanations that focus on the expansion of aggregate demand (biased towards spending on services and housing).[12]

The evidence presented also reinforces the relevance of the methodology developed by Ball and Mankiw for the analysis of the nominal rigidities of Spanish inflation, because the coefficients and the adjusted R^2 of our basic panel data analysis result almost unaffected by the introduction of the different types of control variables. In other words, despite the inclusion of the control variables, the contribution of lagged inflation and the higher moments of the distribution of price changes in the explanation of mean inflation remains unchanged.

We think that the promising evidence gathered in this paper should invite us to deepen in the use of Ball and Mankiw's (1995) methodology in order to explain the aforementioned Spanish inflation differential, using different datasets, extending the consideration of control variables and connecting the results to some microeconomic features of the Spanish economy.

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[11] Moreover, another worrying feature found in our analysis is the positive value of the skewness variable in the 63% of the sample period. This feature may indicate the presence of inflationary tension in the menu cost models framework that we use.

[12] See, among others, Dolado and Jimeno (1997), Estrada and López-Salido (2002) and López-Salido *et al.* (2005).

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Appendix

Table A1: Panel data unit root analysis (1993-1998, with trend) and summary statistics

Variable	Levin, Lin and Chu (2002)		Breitung (2000)	
	Statistic	*Prob.*	*Statistic*	*Prob.*
$S_{j,t}$	-30.62	0.00	-17.84	0.00
$RPV_{j,t}$	-14.29	0.00	1.58	0.00
$\pi_{j,t}$	-12.18	0.00	-3.70	0.00
	Mean	***Max.***	***Min.***	
$S_{j,t}$	0.58	12.98	-9.84	
$RPV_{j,t}$	1.46	3.56	0.52	
$\pi_{j,t}$	0.28	0.81	-0.22	

Table A2: Panel data unit root analysis (1999-2005) and summary statistics

Variable	Levin, Lin and Chu (2002)		Breitung (2000)	
	Statistic	*Prob.*	*Statistic*	*Prob.*
$S_{j,t}$	-25.50	0.00	-13.98	0.00
$RPV_{j,t}$	-4.29	0.00	-2.82	0.00
$\pi_{j,t}$	-27.57	0.00	-18.19	0.00
	Mean	***Max.***	***Min.***	
$S_{j,t}$	0.43	10.48	-12.48	
$RPV_{j,t}$	1.62	2.88	0.53	
$\pi_{j,t}$	0.26	0.74	-0.27	

Table A3: Panel data unit root analysis (1993-2005, with trend) and summary statistics

Variable	Levin, Lin and Chu (2002)		Breitung (2000)	
	Statistic	*Prob.*	*Statistic*	*Prob.*
$S_{j,t}$	-48.38	0.00	-26.92	0.00
$RPV_{j,t}$	-4.20	0.00	0.42	0.66
$\pi_{j,t}$	-24.53	0.00	-5.86	0.00
	Mean	***Max.***	***Min.***	
$S_{j,t}$	0.50	12.90	-12.21	
$RPV_{j,t}$	1.55	3.56	0.52	
$\pi_{j,t}$	0.27	0.81	-0.27	