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Inflation and Relative Prices: Empirical Evidence for the Spanish Economy

María Ángeles Caraballo[1], Carlos Usabiaga[2]

1. Introduction

Nominal price rigidity assumption has generated a vast literature in recent macroeconomics. There are several theories trying to explain it. This paper focuses on New Keynesian menu cost models, and more precisely on the empirical testing strand proposed by Ball and Mankiw (1994, 1995), who explain why a relative price shock, which serves as a measure of a supply shock, can affect the average inflation rate, whereas in a flexible price framework changes in relative prices don't affect average inflation.

Basically, those papers assume a monopolistic competition model characterized by firms that face menu costs while adjusting prices. In this setting, a distribution of the desired relative price changes for the firms can be defined[3]. When firms experience a shock to their desired relative price, they change their prices only if the profit derived from the adjustment is larger than the menu cost. We will analyse the implications of this behavior under two alternative scenarios with respect to inflation: no trend inflation (Ball and Mankiw, 1995) and trend inflation (Ball and Mankiw, 1994):

a) No trend inflation or stable inflation, near zero. In this case, if the distribution of the desired relative price changes of the firms of a sector or geographical area is symmetrical, around zero, as the price adjustment costs differ among firms, there will be an inaction range of the firms around zero; but the firms that are located in the right tail of the distribution will increase their price, whereas those located in the left tail will reduce it. As the distribution is symmetrical, a shock that affects that sector or geographical area won't affect the general price level, because the price increases will balance the reductions. However, if the distribution of price changes is skewed to the right (left), the left tail will be smaller (bigger) than the right one, so the net effect of a shock will be a price increase (reduction). Therefore, a testable implication of the menu cost model is the positive association between average inflation and the skewness of the distribution of the desired price changes. We can also point out that a larger variability[4] will magnify the effects of skewness, increasing the relative weight of the tails; however, if the distribution of desired prices is symmetrical, an increase in variability strengthens both tails equivalently, so the change in variability doesn't affect inflation.

b) Trend inflation. In this context, for a negative shock, firms have two options: to pay the menu cost or let inflation erode their relative prices until the desired level. The higher the inflation is, the faster the erosion process and the smaller the probability of firms paying menu costs are. Therefore, positive inflation reduces the range of the zone in which firms pay menu costs, and reduces their price. On the contrary, for a positive shock, as inflation is also positive, if the firm doesn't pay the menu cost the gap between current and optimal price will increase, so the firms are more likely to pay menu costs and increase their price, strengthening the right tail of the distribution. In conclusion, in a trend inflation framework downward price rigidity appears. Finally, an increase in the variability of the distribution of the shocks increases in absolute values the right tail in relation to the left one, so inflation increases independently of the skewness of the distribution of the shocks.

To sum it up, under this theoretical framework the conclusion is that if average inflation is stable and near zero the inflation-skewness relation is stronger than the inflation-variability relation, whereas in a trend inflation framework the inflation-variability relation is stronger.

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[3] A problem arises when we try to test these models because the desired price changes for the firms can´t be observed, so a proxy like the observed price changes must be used.

[4] Relative price variability is referred to the variance or the standard deviation of the distribution.

The empirical evidence in this area is mixed. In general, positive associations inflation-skewness and inflation-variability are supported by the data, but the results are not conclusive about which relation is stronger[1]. This ambiguity of the results and the lack of studies in this area for the Spanish economy have motivated this paper. In essence, our work tries to clarify whether menu cost models are plausible for the Spanish economy, following the methodology of the moments *à la* Ball and Mankiw.

The rest of the paper is organised as follows. Section 2 describes the data and variables used in our analysis and the empirical methodology followed. Section 3 presents the results of our first approximation to Ball and Mankiw (1995). In section 4 alternative measures of skewness are considered. Section 5 tries to detect if there is a sector in the economy that determines the results obtained along the paper. Section 6 presents feasible extensions of the paper, and section 7 concludes.

2. Data, basic variables and empirical methodology[2]

Our period of analysis is 1993.01-2001.12; period characterized by a mean monthly inflation rate around 0.25%, and lower than 5% in annual terms. We have chosen this low inflation period because around an annual 4-5% is placed the upper limit for which the model predicts a strong inflation-skewness relation.

Our study is based both on consumer and producer prices (CPI -consumer price index- and PPI -producer price index-). For both of them, the data come from the series of monthly change rates of price indices, disaggregated by goods and services (33 subgroups) for consumer prices and by 25 sectors for producer prices, elaborated by the Instituto Nacional de Estadística (INE).

For consumer prices, the weight of each subgroup is offered by the INE (proportion of expenditure made on that article in relation to total expenditure made by households). The weight is kept constant by the INE within the period analysed. When price rigidity models are trying to be tested this constancy is an advantage, because if this is not the case the price indices could change simply due to changes in the weights and not because of flexibility.

For producer prices, the weight of each sector is not available, so the results achieved must be taken cautiously[3]. Anyway, to some extent, it can be meaningful to use non-weighted data, in the sense that if there exists price rigidity in some sectors of the economy but not in others, with the weighted data the presence of rigidity will be captured just in those sectors with a larger weight under the criterion used.

We are aware of the limitations of our data, because the indices hide information when we aggregate; however, at this moment those problems don't have an easy solution, because we haven´t got other kind of data.

We now define some of the essential variables in our analysis[4]. As a proxy of the moments of the shocks distribution we use the second and third cross-sectional moment of the distri-

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[1] Ball and Mankiw (1995), for the US, and Amano and Macklem (1997), for Canada, find a positive inflation-skewness association stronger than the inflation-variability one. Lourenco and Gruen (1995), for Australia, show that for an annual inflation rate lower than 4-5%, the inflation-skewness relation is stronger than the inflation-variability one, but as inflation increases above that rate, the inflation-variability relation is stronger. Hall and Yates (1998), for the United Kingdom, find an inflation-skewness association weaker than the inflation-variability one. Aucremanne et al. (2002), for Belgium, assert that the inflation-skewness association is positive independently of mean inflation. Finally, Döpke and Pierdzioch (2003), for Germany, find both associations positive, but none of them is stronger.

[2] Along this work we refer to some results that haven't been presented explicitly in it in order to avoid a too long exposition. Those results, as well as the data used, before and after their seasonal adjustment, are available from the authors upon request.

[3] Caraballo et al. (2003) compare the results for the weighted and non-weighted variables (CPI) of the estimation of section 3, concluding that the significance of the variables doesn´t change, although the values of the coefficients are slightly larger for skewness and slightly smaller for standard deviation in the case of non-weighted variables.

[4] We have to introduce a precision about notation. It is usual in this literature to distinguish between relative price variability and relative price dispersion: the former is defined as the variance or the standard deviation of the distribution of the price change rate, whereas the latter refers to the distribution of the price levels. In general, attending to the disposability of data, it is common to use the relative price variability, which is the variable used in this paper.

bution of price changes. The expressions of the standard deviation (S_t) and the skewness (A_t) are as follows:

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

where π refers to inflation rate, i to goods, and t to time. Consequently, π_t: is average inflation rate in period t; π_{it}: is inflation rate of subgroup i in period t; and w_i is the weight of each subgroup i used by INE to build the CPI within this period. Obviously, for producer prices we consider the same weight for every sector.

As an interaction measure between S_t and A_t we define: $M_t = S_t \cdot A_t$. If the distribution is symmetrical, M_t takes a null value independently of the standard deviation, but for other value of skewness the value of this variable – in absolute terms – is positively correlated with standard deviation; in other words, the standard deviation magnifies the value of skewness.

There are some issues to point out with respect to the empirical methodology:

1) As data are monthly, they present an important seasonal component, that has been eliminated by means of an X-12 ARIMA method. All the results presented in this work refer to seasonally adjusted variables. In order to choose the method for seasonal adjustment, we have started applying TRAMO-SEATS and the X-12 ARIMA method; so we run the regressions presented in Tables 1, 2 and 3 using the seasonally adjusted variables obtained under both methods. Additionaly, as seasonality has a stable pattern along our period of analysis, we have introduced dummies in those regressions instead of seasonal adjustment, choosing those dummies that are significant. Finally, we included dummies for the months 1, 2, 10, 11 and 12 for CPI, and dummies for the months 1 and 9 for PPI. We have observed that introducing dummies and including 12 lags for the dependent variable we still have autocorrelation of order 12 in the residuals. Using seasonally adjusted variables with TRAMO-SEATS we need to introduce 12 lags for the dependent variable and we have autocorrelation problems at least of order 5. With the X-12 ARIMA method we obtain the best results, because we need just one lag for the dependent variable in order to remove autocorrelation in the residuals. So, finally, for the whole paper we have chosen the X-12 ARIMA method. Anyway, the results concerning the value and significance of the variables are very similar for the three aforementioned methods.

2) We have checked the stationarity of the series applying the ADF test (see the appendix). As we have used the seasonally adjusted series, we have applied the unit root test to those series, and not to the original ones[1].

3) As in our analysis we basically use the inflation rate as dependent variable and the higher moments of the distribution of the inflation rate as regressors, multicollinearity problems could appear. To tackle this problem we calculate the correlation coefficients between standard deviation and skewness. As their values are under 0.3, we have decided to introduce both variables jointly in the regressions[2].

4) The regressions are estimated by using ordinary least squares (OLS)[3] and, as usual, the value of the t statistic (in brackets in the tables) is corrected of heteroscedasticity by means of the White method.

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[1] In this sense a problem could arise, because 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. As Ghysels (1990) asserts that this problem arises when seasonality is stochastic, we think that this problem doesn´t affect our data (see Ghysels and Perron (1993) for literature related to the unit root test applied to seasonally adjusted series).

[2] Specifically, the correlation between SCPIt and ACPIt is 0.16, and the correlation between SPPIt and APPIt is 0.145.

[3] As it is well known, if the lagged endogenous variable is not correlated with the error term, the properties of the OLS estimators hold. In order to verify that there is no correlation we have estimated the model using OLS and we have checked that there isn´t autocorrelation in the residuals.

3. Estimation of the baseline specifications

Following Ball and Mankiw (1995), 5 regressions have been estimated, and we basically compare the adjusted R^2 in order to study the contribution of the skewness to the estimations.

The regressions of Tables 1 and 2, following the order of the columns, are as follows:

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \varepsilon_t, \tag{1}$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_2 S_t + \varepsilon_t, \tag{2}$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_3 A_t + \varepsilon_t, \tag{3}$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_2 S_t + \beta_3 A_t + \varepsilon_t, \tag{4}$$

$$\pi_t = \alpha + \beta_1 \pi_{t-1} + \beta_4 M_t + \varepsilon_t, \tag{5}$$

We have introduced the lags of the dependent variable required to remove autocorrelation in the residuals. In order to choose the number of lags, we start including one lag for inflation, and the Breusch-Godfrey test of autocorrelation is applied until the order 12; if autocorrelation appears we include a second lag, and the Breusch-Godfrey test is run again, etc. For our estimations just one lag has been enough to remove autocorrelation.

In the previous equations π is the rate of inflation, of CPI or PPI depending on which index is being used. In the tables we add the labels CPI or PPI to the aforementioned variables. For example, $ACPI_t$ is the skewness of the distribution of the inflation corresponding to CPI.

In Table 1 the dependent variable is the rate of inflation of CPI, whereas the regressors are the moments of the distribution of CPI, and in Table 2 the dependent variable is the rate of inflation of PPI, whereas the regressors are the moments of the distribution of PPI.

With respect to Table 1 (CPI), some results arise. $SCPI_t$ is significant when it is the only moment included in the regression, but it is not when $ACPI_t$ is added. On other contrary, $ACPI_t$ is always significant. All the coefficients of the table are positive. As far as the relative contribution of skewness and standard deviation is concerned, the R^2 increases when one of them is included in the regression, but it is higher with $ACPI_t$; and the best R^2 is obtained when both variables are included. It can also be observed that the coefficient of the standard deviation is a bit larger than the coefficient of the skewness. Finally, as it was expected, the interaction variable is significant.

With regard to Table 2 (PPI), the results are very similar. $SPPI_t$ is not significant in any case, whereas $APPI_t$ is significant; therefore the contribution of skewness to the R^2 is greater than the contribution of standard deviation. Finally, the interaction variable is significant.

To sum it up, the significance, both of the standard deviation and the skewness, shows the presence of rigidities in the price adjustment. It is also observed that skewness is more significant than standard deviation, as predicted by Ball and Mankiw (1995).

In the regressions of both tables the residuals are normally distributed and there are not problems of autocorrelation. A problem may arise because the results of Tables 1 and 2 can be distorted because of the small sample bias argued by Bryan and Cecchetti (1996,1999). According to them, even for a zero-mean sample extracted from a zero-mean symmetrical distribution, an additional single draw from the extreme positive tail of the distribution makes both the sample mean and the sample skewness positive; therefore a positive association mean-skewness is obtained. In this line, Amano and Macklem (1997) argue that this problem arises when the regressors are the higher moments of the distribution of a set of consumer (producer) prices and the dependent variable is the aggregate inflation rate obtained from the same set of consumer (producer) prices. Consequently, the sample bias can be avoided by using different sets of prices in order to calculate average inflation and the higher moments of the distribution. Thus, in order to control the small sample bias mentioned by Bryan and Cecchetti (1999), we have used the rate of inflation of

CPI as dependent variable and the moments of the PPI distribution as regressors[1]. The results of this estimation appear in Table 3.

As it can be seen from Table 3, skewness is significant, whereas standard deviation is not significant; these results support again the Ball and Mankiw (1995) conclusions. Anyway, these results must be taken cautiously, because in the equations that contain skewness the residuals are not normally distributed.

Table 1

Estimations of the baseline specifications (CPI)[2]

α	0.0120 (0.000)	0.049 (0.210)	0.095 (0.001)	0.044 (0.238)	0.110 (0.000)
CPI_{t-1}	0.546 (0.000)	0.527 (0.000)	0.548 (0.000)	0.534 (0.000)	0.539 (0.000)
$SCPI_t$		0.099 (0.015)		0.074 (0.507)	
$ACPI_t$			0.023 (0.001)	0.021 (0.003)	
$MCPI_t$					0.015 (0.012)
Adjusted R^2	0.292	0.314	0.364	0.374	0.335
BG	0.731	0.921	0.771	0.780	0.904
JB	0.093	0.077	0.255	0.176	0.342

Table 2

Estimations of the baseline specifications (PPI)

α	0.053 (0.028)	-0.021 (0.729)	0.049 (0.029)	-0.004 (0.929)	0.049 (0.027)
PPI_{t-1}	0.734 (0.000)	0.726 (0.000)	0.629 (0.000)	0.627 (0.000)	0.641 (0.000)
$SPPI_t$		0.086 (0.235)		0.062 (0.350)	
$APPI_t$			0.040 (0.000)	0.039 (0.000)	
$MPPI_t$					0.031 (0.000)
Adjusted R^2	0.540	0.549	0.611	0.614	0.618
BG	0.562	0.485	0.691	0.815	0.975
JB	0.182	0.170	0.092	0.287	0.282

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[1] We have decided this option because PPI anticipates CPI (see Quilis (1999)).

[2] From now on, for every table BG denotes the p-value of a Breusch-Godfrey test on autocorrelation of first order (recall that the null hypothesis is no autocorrelation, so if the p-values are over 0.05 we can´t reject the null hypothesis at a level of significance of 5%), and JB denotes the p-value of a Jarque-Bera test on normality of the residuals (recall that the null hypothesis is the normality of the residuals, so if the p-values are over 0.05 we can´t reject the null hypothesis at a level of significance of 5%).

Table 3

Estimations of the baseline specifications (CPI as dependent variable, moments of PPI as regressors)[1]

α	0.113 (0.008)	0.112 (0.000)	0.116 (0.006)	0.111 (0.000)
CPI$_{t-1}$	0.546 (0.000)	0.536 (0.000)	0.536 (0.000)	0.541 (0.000)
SPPI$_t$	0.007 (0.838)		-0.004 (0.900)	
APPI$_t$		0.016 (0.002)	0.016 (0.002)	
MPPI$_t$				0.014 (0.001)
Adjusted R^2	0.286	0.335	0.329	0.355
BG	0.769	0.724	0.694	0.790
JB	0.105	0.019	0.016	0.080

4. Alternative measures of skewness

Up to this moment, it seems that skewness is a more relevant variable than standard deviation in order to explain inflation. To check this result, following Ball and Mankiw (1995), in this section alternatives measures of skewness are defined. Those authors assert that their theory relates inflation with the size of the tails of the distribution of relative price changes, so they define a variable to measure the tails and also to capture how the effects of skewness are magnified by variability. Specifically, for a cut-off X, AX_t is defined as:

$$AX_t = \sum_{i=1}^{n} w_i(\pi_{it} - \pi_t)D_i^- + \sum_{i=1}^{m} w_i(\pi_{it} - \pi_t)D_i^+ ,$$

where D_i^- and D_i^+ are dummy variables. The former variable 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 variable is one when i^{th} industry's relative price change falls in the upper X per cent of the distribution and zero otherwise. Therefore, this variable substracts the mass in the upper tail of the distribution of price changes from the mass in the lower tail. AX_t is zero for a symmetrical distribution of relative price changes and positive (negative) when the right (left) tail is larger than the left (right) one. Moreover, for a given skewness, the larger the variability is, the larger the tails are; so with the same variable we are combining the effects of skewness with its interaction with the variability. The choice of X is arbitrary; we have chosen 10 and 25 in order to compare our results with those of Ball and Mankiw (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 we have done with AX_t, it can be defined a variable which increases the weights linearly with the size of the adjustment:

$$Q_t = \sum_{i} w_i |\pi_{it} - \pi_t|(\pi_{it} - \pi_t),$$

As it can be observed, Q_t is a weighted average of the product of each relative price change and its own absolute value, and it has the same properties of AX_t: it is zero for a symmetrical distribution, positive for a right skewed distribution, and negative for a left skewed distribution; and it is magnified with a larger variability.

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[1] Again, we have introduced the lags of the dependent variable required to remove autocorrelation (one lag is enough for our data).

These new variables are seasonally adjusted by means of the X-12 ARIMA method, and the ADF test has been applied as well (the results can be found in the appendix). Recall that we introduce the lags of the dependent variable required to remove the autocorrelation in the residuals[1].

Table 4

Estimations with alternatives measures of skewness (CPI)

α	0.120 (0.000)	0.091 (0.005)	0.112 (0.000)	0.052 (0.163)	0.109 (0.000)	0.122 (0.014)
CPI_{t-1}	0.381 (0.000)	0.378 (0.000)	0.510 (0.000)	0.498 (0.000)	0.485 (0.000)	0.485 (0.000)
$SCPI_t$		0.041 (0.231)		0.085 (0.028)		-0.019 (0.756)
$ACPI10_t$	1.715 (0.000)	1.661 (0.000)				
$ACPI25_t$			0.963 (0.041)	0.861 (0.050)		
$QCPI_t$					0.067 (0.015)	0.070 (0.021)
Adjusted R^2	0.515	0.515	0.323	0.337	0.399	0.394
BG	0.550	0.476	0.880	0.756	0.640	0.636
JB	0.133	0.100	0.126	0.094	0.912	0.946

Table 5

Estimations with alternative measures of skewness (PPI)

α	0.086 (0.000)	0.013 (0.780)	0.069 (0.007)	-0.016 (0.772)	0.043 (0.032)	0.047 (0.323)
PPI_{t-1}	0.616 (0.000)	0.608 (0.000)	0.705 (0.000)	0.690 (0.000)	0.572 (0.000)	0.572 (0.000)
$SPPI_t$		0.083 (0.099)		0.101 (0.127)		-0.005 (0.924)
$APPI10_t$	0.960 (0.000)	0.957 (0.000)				
$APPI25_t$			0.490 (0.040)	0.486 (0.044)		
$QPPI_t$					0.096 (0.000)	0.097 (0.000)
Adjusted R^2	0.646	0.656	0.553	0.567	0.686	0.683
BG	0.393	0.263	0.641	0.525	0.946	0.946
JB	0.009	0.106	0.300	0.215	0.050	0.050

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[1] In Tables 4 and 5 one lag is enough. However, in the regressions of columns (5) and (6) of Table 6 there are problems of autocorrelation of 7 to 10 order, but as they don´t disappear introducing more lags, we present the results of the estimation with one lag for the dependent variable.

Table 6

Estimations with alternative measures of skewness
(CPI as dependent variable, moments of PPI as regressors)

α	0.126 (0.000)	0.123 (0.000)	0.125 (0.000)	0.111 (0.010)	0.112 (0.000)	0.154 (0.000)
CPI_{t-1}	0.535 (0.000)	0.535 (0.005)	0.551 (0.000)	0.552 (0.000)	0.496 (0.000)	0.488 (0.000)
$SPPI_t$		0.004 (0.906)		0.016 (0.654)		-0.048 (0.145)
$APPI10_t$	0.375 (0.002)	0.374 (0.003)				
$APPI25_t$			0.293 (0.077)	0.301 (0.065)		
$QPPI_t$					0.049 (0.000)	0.053 (0.000)
Adjusted R^2	0.350	0.344	0.315	0.310	0.439	0.447
BG	0.538	0.510	0.787	0.670	0.283	0.337
JB	0.108	0.123	0.375	0.458	0.432	0.468

In these regressions the residuals are normally distributed (except for regression 1, in Table 3) and there are not problems of autocorrelation. As it can be seen from these tables the relevance of skewness holds[1], so this result is robust across the different skewness measures. In conclusion, there are several ways to measure the skewness in relative price changes, though there doesn't exist a clear criterion for choosing among them.

5. The relevance of the oil sector

As Bryan and Cecchetti (1996, 1999) have pointed out, the skewness is very sensitive to outliers, so our results could be heavily determined by a sector or sectors. This issue has given rise to an extensive literature. For instance, Fischer (1981) asserts that the relation between variability and inflation in the US for CPI is dominated by energy and food shocks. Bomberger and Makinen (1993) show that Parks' findings for the 1948-1975 period about the strong relation between the rate of inflation and the variability of relative price changes depend on the inclusion of a single observation: the oil price shock in 1974; when this data is removed the relation doesn't hold. However, Jaramillo (1999) responds that Parks' results can be maintained excluding 1974, but extending Parks' sample to 1996 and allowing for a much higher degree of disaggregation.

In this sense, because our sample is short, we haven't removed outliers; in other words, we have not eliminated a period because there is a big shock. And, in order to check the relevance of the different sectors in our type of analysis, we have calculated the standard deviation, the skewness and the alternative measures of skewness of the inflation distribution of CPI excluding each time a different group. Then we have estimated the above regressions for CPI as dependent variable and these new variables as regressors, and there are not remarkable changes in the results. On the contrary, following the same procedure for PPI, we have concluded that the oil sector is determining our results, in the sense that none of the moments and the alternative measures of skewness are significant when they are constructed excluding the oil sector.

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[1] Except for APPI25t in Table 6, where it is only significant at a 10% level.

Table 7

Estimations excluding the oil sector (PPI)[1]

α	0.026 (0.517)	0.051 (0.049)	0.008 (0.869)	0.052 (0.035)	0.043 (0.058)	0.041 (0.093)
PPI_{t-1}	0.729 (0.000)	0.734 (0.000)	0.725 (0.000)	0.734 (0.000)	0.689 (0.000)	0.683 (0.000)
$SPPI^*_t$	0.072 (0.295)		0.066 (0.349)			
$APPI^*_t$		0.007 (0.503)	0.005 (0.661)			
$APPI10^*_t$				-0.034 (0.896)		
$APPI25^*_t$					-0.596 (0.002)	
$QPPI^*_t$						0.067 (0.219)
Adjusted R^2	0.541	0.537	0.537	0.535	0.582	0.577
BG	0.538	0.585	0.503	0.557	0.381	0.890
JB	0.280	0.214	0.303	0.175	0.025	0.313

Table 8

Estimations excluding the oil sector (CPI as dependent variable, moments of PPI as regressors)[2]

α	0.058 (0.135)	0.117 (0.000)	0.058 (0.137)	0.120 (0.000)	0.121 (0.000)	0.105 (0.000)
CPI_{t-1}	0.521 (0.000)	0.552 (0.000)	0.520 (0.000)	0.542 (0.000)	0.518 (0.000)	0.531 (0.000)
$SPPI^*_t$	0.100 (0.061)		0.101 (0.057)			
$APPI^*_t$		0.003 (0.622)	-0.0006 (0.925)			
$APPI10^*_t$				-0.033 (0.848)		
$APPI25^*_t$					-0.197 (0.190)	
$QPPI^*_t$						0.055 (0.000)
Adjusted R^2	0.322	0.287	0.316	0.286	0.303	0.387
BG	0.825	0.725	0.820	0.701	0.449	0.856
JB	0.269	0.102	0.275	0.072	0.000	0.144

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[1] The asterisk denotes that the variable doesn´t include the oil sector. In this table the autocorrelation in the residuals is removed with one lag of the dependent variable.

[2] In the column (6) we have autocorrelation of order higher than one, but as it doesn´t dissappear introducing more lags, we present the results of the estimation with one lag for the dependent variable.

Both tables show that the strong relation between inflation and skewness for PPI vanishes when the oil sector is removed.

6. Extensions

In this section we describe several research lines as feasible extensions of our paper. As it can be observed, we have made preliminary work in most of them.

1) One of the feasible extensions of this paper is the causality analysis. In this sense, the Ball and Mankiw (1995) approach reverses the causality relation between inflation and relative price variability generally accepted, departing from an important branch of literature inspired in Lucas (1972), that studies the effects of a higher inflation on price variability[1]. Ball and Mankiw (1995) also establish that, at least theoretically, it is possible to think in a direction of causality from inflation to skewness, but they assert that the US data don't support this fact[2]. Amano and Macklem (1997) and Hall and Yates (1998) also analyse the causality relation between inflation and skewness with the Granger causality test, concluding that in general there is empirical evidence supporting the causality from skewness to inflation. Following the line of these authors we have applied the Granger causality test to our data as well, and we don't obtain clear conclusions, so we think that alternative techniques must be used in order to deepen in this issue[3].

2) This kind of studies can help us to choose between sticky and flexible price models. In this sense, the correlation between mean inflation and skewness shown in this paper can be explained not only with the sticky price model developed by Ball and Mankiw (1995) but also, for example, with the flexible price model proposed by Balke and Wynne (2000). These authors set up a multisectoral general equilibrium model with flexible prices, which yields a positive correlation between inflation and skewness. However, sticky and flexible price models differ in the predictions of the short and long run effects. In the Ball and Mankiw model the correlation arises because of short run considerations, as menu costs, but as soon as the firms adjust their prices the correlation should disappear. Therefore, as the length of time over which price changes are measured increases (from months to quarters, from quarters to years) the correlation should weaken. On the contrary, for the Balke and Wynne model the correlation doesn't disappear in the long run, because it is due to real factors; even it might become more pronounced. In this sense, we think that it could be interesting to test these conclusions with our data. Our preliminary results in this line, using different time units (months, quarters, years), point out in the direction of the Ball and Mankiw model.

3) A strand of research very interesting for us is the comparison of the results among different inflationary regimes. In this sense, we are working in two directions: considering a longer period for the Spanish data (1974-2002) – period characterized by a higher mean inflation –, and comparing with other countries characterized by inflationary regimes very different from European standards; in this second line we want to consider several Latin-American countries, beginning with Argentina. Our preliminary results in this line also point out in the direction of Ball and Mankiw, because for those two alternative options (characterized by a higher average inflation) we observe a weaker inflation-skewness relation and a stronger inflation-standard deviation relation.

4) Another extension of this work could be the inclusion of real variables, in order to study, for example, the Phillips curve or the aggregate supply function. We have already done some work in this line, introducing as regressor, on one hand, the change of the monthly unem-

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[1] The question of causality in this area was also studied by Fischer (1981), for US quarterly data the period of 1948-1972. The results offered by this author don't show a clear causality pattern between the variables. In this same line, some recent works aren't conclusive.

[2] In order to evaluate this possibility these authors attend to the historical evolution of US inflation, and observe anomalies in the data: in 1975 inflation decreased while skewness was positive, and in 1982 inflation also decreased while the distribution of price changes was symmetrical. Ball and Mankiw argue that in both cases the reduction of inflation was due to a restrictive monetary policy, in other words to an exogenous demand factor, and the fact that skewness wasn't affected shows that inflation changes don't affect that variable.

[3] Caraballo and Usabiaga (2003, section 5) pay attention to this topic.

ployment rate[1] and, on the other hand, the cyclical unemployment measured as the deviation of the observed unemployment rate from the trend unemployment rate calculated with the Hodrick-Prescott filter[2]. For both cases unemployment is not significant. We would like to introduce in our analysis additional real variables as regressors.

5) Finally, a research line that could be explored in depth is the relation between our results and the level of aggregation of the data. For example, we have observed that for a higher level of disaggregation – we consider 57 "rúbricas" for CPI – the inflation-skewness relation is stronger. There is a lot of international literature in this area about the relation of the results with the level of aggregation of the data.

7. Conclusions

The significance of the standard deviation and, especially, of the skewness of the distribution of price changes, obtained in our work from different perspectives, shows the "vulnerability" of Spanish inflation in terms of relative price shocks.

Our analysis for Spanish data corroborates the results of Ball and Mankiw (1995) for the US, about the relevance of the skewness in explaining inflation. Their results in the line that the standard deviation coefficient is higher than the skewness coefficient, and that the estimations containing skewness present a higher R^2 are also confirmed.

Those results invite us to deepen in the plausibility of price rigidity models for the Spanish economy (menu cost models in the case of Ball and Mankiw).

Our results also point out that the oil sector is very relevant, in the type of analysis implemented in this paper, for the PPI.

References

1. Amano, R.A. and Macklem, R. (1997): "Menu Costs, Relative Prices, and Inflation: Evidence for Canada", Bank of Canada, Working Paper 97-14.
2. 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.
3. Balke, N. and Wynne, M. (2000): "An Equilibrium Analysis of Relative Changes and Aggregate Inflation", *Journal of Monetary Economics*, 45, pp. 269-292.
4. Ball, L. and Mankiw, N.G. (1994): "Asymmetric Price Adjustment and Economic Fluctuations", *Economic Journal*, 104, March, pp. 247-261.
5. Ball, L. and Mankiw, N.G. (1995): "Relative-Price Changes as Aggregate Supply Shocks", *Quarterly Journal of Economics*, 110 (1), pp. 161-193.
6. Bomberger, W.A. and Makinen, G.E. (1993): "Inflation and Relative Price Variability: Parks´ Study Reexamined", *Journal of Money, Credit and Banking*, 25 (4), pp. 854-861.
7. Bryan, M.F. and Cecchetti, S.G. (1996): "Inflation and the Distribution of Price Changes", NBER, Working Paper 5793.
8. Bryan, M.F. and Cecchetti, S.G. (1999): "Inflation and the Distribution of Price Changes", *Review of Economics and Statistics*, 81 (2), pp. 188-196.
9. Caraballo, M.A., Sánchez, A. and Usabiaga, C. (2003): "Análisis Regional de la Rigidez de Precios en la Economía Española", *VI Encuentro de Economía Aplicada*, Granada.
10. Caraballo, M.A. and Usabiaga, C. (2003): "Análisis de la Estructura de la Inflación de las Regiones Españolas: La Metodología de Ball y Mankiw", centrA, Documento de Trabajo E2003/44.
11. 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.

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[1] The unemployment rate has been obtained by dividing the monthly number of unemployees (Instituto Nacional de Empleo) by the labor force of the corresponding quarter (Encuesta de Población Activa, INE).

[2] We use the standard smoothing parameters for the Hodrick-Prescott filter.

12. Fischer, S. (1981): "Relative Shocks, Relative Price Variability, and Inflation", *Brookings Papers on Economic Activity*, 2, pp. 381-431.
13. 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.
14. 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.
15. Hall, S. and Yates, A. (1998): "Are There Downward Nominal Rigidities in Product Markets?", Bank of England, Working Paper Series, 80.
16. Jaramillo, C.F. (1999): "Inflation and Relative Price Variability: Reinstating Parks´ Results", *Journal of Money, Credit and Banking*, 31 (3), pp. 375-385.
17. Lourenco, R.D.A. and Gruen, D. (1995): "Price Stickiness and Inflation", Reserve Bank of Australia, Research Discussion Paper 9502.
18. Lucas, R.E. (1972): "Expectations and the Neutrality of Money", *Journal of Economic Theory,* 4, pp. 103-124.
19. Quilis, E.M. (1999): "Nota sobre la Relación Cíclica entre los Índices de Precios al Consumo (IPC) e Industriales (IPRI)", *Boletín Trimestral de Coyuntura*, 73 (septiembre).

Appendix: Results of the unit root test

We have applied the Akaike criterion in order to select the number of lags. The null hypothesis is the existence of a unit root. Once the significance of the constant, the trend, or both of them is analysed, the value of the ADF statistics as well as the corresponding p-value in order to reject the null hypothesis is offered.

Table 9

Unit root test (CPI)

Variable	Number of lags	Constant	Trend	ADF statistic	p-value
CPI_t	0	yes	no	-5.536	0.000
$SCPI_t$	0	yes	no	-9.295	0.000
$ACPI_t$	0	yes	yes	-10.077	0.000
$MCPI_t$	0	yes	no	-9.813	0.000
$ACPI10_t$	0	yes	no	-8.024	0.000
$ACPI25_t$	0	yes	no	-8.347	0.000
$QCPI_t$	0	yes	no	-9.763	0.000

Table 10

Unit root test (PPI)

Variable	Number of lags	Constant	Trend	ADF statistic	p-value
PPI_t	2	no	no	-2.267	0.023
$SPPI_t$	1	yes	yes	-8.860	0.000
$APPI_t$	2	no	no	-3.435	0.000
$MPPI_t$	1	no	no	-7.009	0.000
$APPI10_t$	2	no	no	-3.981	0.000
$APPI25_t$	2	no	no	-3.872	0.000
$QPPI_t$	0	yes	no	-7.995	0.000
$SPPI^*_t$	1	yes	yes	-8.577	0.000
$APPI^*_t$	3	no	no	-3.020	0.002
$APPI10^*_t$	2	yes	no	-4.624	0.000
$APPI25^*_t$	0	no	no	-6.468	0.000
$QPPI^*_t$	0	yes	no	-9.061	0.000