RELATIVE-PRICE CHANGES AND INFLATION: EVIDENCE FROM SPAIN[#]

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Abstract: This paper explores the connection between inflation and its higher-order moments in the case of the Spanish economy from 1975 to 1999. We found evidence of a strong positive correlation between aggregate inflation and the distribution of relative-change prices. Also, a standard Phillips curve is used to study aggregate supply shocks. The results strongly support our model from 1975 onwards and we are able to document, for the first time, the dynamic features of the relative-price adjustment process. Small sample problems are addressed.

Keywords: Inflation; Cross-section distribution of prices; Spain

JEL Classification: E31

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I. INTRODUCTION.

Evidence of a significant statistical relationship between inflation and the higher cross-sectional moments (variance and skewness) of the distribution of prices is amply available in the literature. Based on Vining and Elwetowski (1976) seminal paper, different avenues of inquiry have been established to both ratify the existence of this relationship and to determine its origins[1]. Most attention has been directed towards the study of the relationship between inflation and its second higher moment[2], although recently the exploration of the relationship between inflation and its third higher moment has gained momentum. Ball and Mankiw (1995) and Balke and Wynne (2000) have built on previous work by Batchelor (1981), Blejer (1983), and Mizon, Safford, and Thomas (1990) to study the nature of this empirical finding. Although the existence of this empirical regularity has been reported under a variety of circumstances for a number of different countries[3], its categorization as a macroeconomic stylized fact has been questioned by the work of Bryan and Cecchetti (1999a) and, in some measure, by Verbrugge (1999.) These authors attribute to small-sample bias the reason for the detected correlation between inflation and its higher-order moments[4]. Simultaneously, the question of the origin of this correlation is also open to debate. The most frequently cited Neo-Keynesian argument, invoking the existence of menu costs to justify the apparent sluggishness of the relative price adjustment processes, has only recently been questioned by Balke and Wynne (2000.) These authors argue that technology shocks are, instead, responsible for this empirical regularity.

In this paper we apply the analysis developed by Ball and Mankiw to study the relationship between aggregate inflation and the cross-sectional distribution of relative-price changes in the context of contemporary Spain, while addressing the relevant points raised by Bryan and Cecchetti and contesting the real-business-cycle interpretation of this phenomenon given by Balke and Wynne.

The last 25 years in the history of Spain have been extraordinarily interesting and present an illustrative case study in matters pertaining to inflationary and deflationary processes. Throughout this quarter of a century, political and social changes have gone hand in hand with a profound transformation of the economy. The political transition from an authoritative regime to a parliamentary democracy, the full inclusion of Spain into the European Union, and later its participation into the European Monetary Union, have provided powerful stimuli for the rapid modernization of the country. During this time, world-wide economic shocks have also generated additional jolts to the price-adjustment mechanism, usually translating into relatively high price levels. The persistence of inflation rates consistently higher than those of other EU members has been extensively studied. Attention has been focused on wage behavior (Mauleon and Bara (1996)), inflation expectations (Sobczak (1998)), monetary policies and shocks (Malo de Molina et al (1998), Ayuso et al (1998), Andres et al (1998)), the domestic price gap (Garcia Herrero and Pradham (1998)), and only marginally on nominal price rigidities (Borondo and González (1997)). Our approach is based on Ball and Mankiw (1995), linking the mean inflation and the cross-sectional skewness in the distribution of price changes. This would be an indicator of the existence of significant menu costs in the Spanish economy. Additionally, we will expand our analysis to study the tradeoff between inflation and unemployment in Spain, contributing to the work of Dolado, López-Salido and Vega (2000), Dolado, Gónzalez-Páramo and Viñals (1999), Andres et al (1998), Dolado, López-Salido and Vega (1996), and Rodriguez-Prado (1995).

Using monthly two-digit and three-digit Producer Price Index (PPI) data, we examine the cross-sectional distribution of price changes from January 1975 to January 2000. We find substantial variation in the third moment of this distribution: in the years prior to Spain's admission into the EEC it is sharply skewed to the right and exhibits great variability; after 1986 there is a dramatic drop in variability and the distribution becomes fairly symmetrical. The menu cost paradigm provides the theoretical framework for the positive relationship between changes in aggregate inflation and the skewness in relative-price variation. Our empirical analysis confirms that connection. We also find it to be highly significant and dynamically consistent. The theory indicates that current aggregate inflation should be negatively linked to past distributions of shocks since firms face a range of inaction when adjusting their price schedules. Previous research efforts in this area haven't been able to document this feature of price-setting processes. We find it present and significantly robust in the case of Spain. Lastly, by turning our inflation equations into Phillips curves by incorporating unemployment we contribute to the study of a tradeoff that has proven difficult to assess in the recent economic history of the country.

The remainder of the paper is divided into six sections. Section II summarizes the evolution of the Spanish economy during the last 25 years and points out the critical events that support our line of analysis. In Section III we present our data set, variables, and provide some descriptive statistics. Section IV contains a brief description of our theoretical framework of analysis. Section V presents the equations to be estimated and discusses their alternative formulations. There, we also address the potential small sample bias. Estimation results and diagnostic tests are reported in Section VI, followed by their discussion. Section VII summarizes our findings.

II. ECONOMIC DEVELOPMENTS IN SPAIN: 1975-1999.

The Spanish economy from 1975 to 1999 experienced a variety of internal changes and external shocks vividly represented in its inflation levels. High rates of economic growth during the late 60s and early 70s (an average of 6%) were accompanied by demand-pull inflation that reached an annual rate of 14% just before the first oil crisis in December 1973[5]. The oil shocks and the political uncertainty of the transitional period from Franco's dictatorship to parliamentary democracy devastated the industrial base that had arisen during the economic miracle of the 60s and 70s. Its key elements -iron and steel, shipbuilding, cement- were all energy intensive, so they were especially hard hit by the hikes in oil prices. Acute cost-push inflation ensued. The unsuccessful policy responses designed to tackle this problem were developed within a political and social framework of instability that took over the country until the early 80s. Stagflation peaked in 1977, with 2.4% GDP growth and 26% inflation. Unemployment became a growing concern due to its threefold increase from 1974 to 1978, reaching 10% at the end of the decade.

Admission into the European Economic Community (EEC), later to become the European Union (EU), was seen as the way out of this quagmire. It would provide an economic stimulus to private entrepreneurs and bring price stability and faster growth for a Spanish economy facing increasing unemployment. The process of trade liberalization associated with EU membership would be the fourth, and the most dramatic one, in a sequential wave of tariff reform processes started in 1959[6]. Remarkably, although industrial growth began again in the early 80s, the level of unemployment remained the highest in Europe. After an almost decade-long negotiation process, drawn out by the political resistance of some members, admission was finally achieved on January 1, 1986. The terms were demanding. Full membership into the EEC would be phased in over seven years and complete participation into the Common Agricultural Policy would be phased in over a ten-year period. The adjustment process required for admission, and the "discipline effect" associated with being a full member of the EEC, proved beneficial. Spanish growth rates from 1985 to 1989 exceeded those of the EEC-12 by 1.4% and unemployment fell from 21.5% to 16%[7].

During this period, Spain's central bank achieved some degree of de facto independence from the central government when it joined the Exchange Rate Mechanism of the European Monetary System. The convergence criteria set by the Maastricht Treaty in 1992 to access the European Monetary Union by the end of the decade reinforced fiscal and monetary discipline. Not without sacrifices in employment and growth levels, the strict upper limits on government spending, debt, and inflation rates were met. The global economic slowdown of the early 90s fueled the unemployment rate -that peaked at 24.2% in 1994. Subsequently, the manufacturing, construction and services sectors have shown clear signs of recovery, bringing down the unemployment rate to levels not registered since the early 70s. It must be pointed out that throughout these years the inflation rate was constantly in the decline, though at a very slow pace. Today, unemployment remains well above 10% and inflation has yet to go below its historical minimum of 1.8% (1998.)

Since the persistence of a positive inflation differential with respect to other members of the European Monetary Union is a serious concern for the Spanish economy, we propose a new explanation to this situation. We argue that firms facing menu costs have a range of inaction when adjusting their nominal prices. Therefore, only large relative price adjustments warrant paying such cost. The dramatic oil price increases of the 70s provided such a supply shock, with inflationary connotations. Partnership in the EEC, later to become the European Union (EU), produced a deflationary shock by exposing the economy to external competition. The accelerated process of real convergence required for membership in the European Monetary Union also imposes fiscal and monetary policy discipline and is conducive to lower inflation.

III. DATA AND DESCRIPTIVE STATISTICS.

Our data set encompasses the period from January 1975 to January 2000[8]. Due to the 1993 change in data collection methodology by the Spanish Instituto Nacional de Estadistica (INE) the only consistent time series of the disaggregated monthly Producer Price Index (PPI) are available at the two-digits level with 1990 base year[9]. For additional analysis at the three-digit level we will refer to the monthly PPI with 1974 base year that was compiled until September 1992. Following the literature, we computed the measure of the distribution of price changes as the cross-sectional skewness of PPI inflation in logs. We also found useful to estimate such distribution by computing the cross-sectional skewness of PPI in levels and both specifications are alternatively tested.

The PPI observations were organized on a 301x24 matrix for the two-digit sectors, and on a 213x93 matrix for the three-digit sectors. The measures of cross-sectional skewness for the January 1975 to January 2000 period and the January 1974 to September 1992 period were calculated by computing:

$$s_t = \frac{\frac{1}{N}\sum_{j=1}^{N}(p_{t_j} - \bar{p}_t)^3}{\hat{s}_{p_t}^{3/2}} \qquad t=1, 2\ldots T$$

where N is the number of sectors of activity (24 at the two-digit disaggregation level and 93 at the three-digit disaggregation level), π is either the PPI in levels (rendering s_t^P) or the PPI inflation in logs (rendering s_t^p), and $\hat{s}_{p_t}^{3/2}$ is the cubic root of the estimated variance.

To illustrate the levels of asymmetries we present Figure 1 with the evolution of the skewness of PPI in levels from 1975 to 1992[10]. Substantial variation in the third moment of the cross-sectional price distribution is evident. The figure shows that there is considerable positive variation (1.65 on average) in the distribution of price changes prior to 1986, the date of Spain's formal admission to the EEC, and that theretofore it is sharply reduced (-0.44 on average for the 1985-2000 period). Our sample period does not record the first OPEC shock but the second one is clearly noticeable in 1978. Our Figure 2 presents the inflation rate, measured as the log difference of the PPI for all commodities. A significant drop in its mean value (from 0.0013 to 0.0019) and standard deviation (from 0.0077 to 0.0032) after 1986 is also clearly observable. The evolution of the PPI inflation skewness is reported in Figure 3. Notice that its mean value drops from 1.55 to 0.24 after 1985. Lastly, Figure 4 plots the unemployment rate. Figures for the monthly unemployment rate were obtained from the Encuesta de Población Activa (also recorded by the INE) after applying a cubic spline with the last observation matched to the originally quarterly data. Two distinct business cycles can be observed during the 1976-1999 period, with an acute recession straddled into the second subsample.

The basic empirical prediction of our model is apparent in Figures 1 and 2. The skewness of price changes varies substantially over time and it does so together with the inflation rate. Years of substantial positive skewness (1980, 1983) are years of increasing inflation whereas years of negative skewness (1986 and following) are years of significant deflation.

IV. THEORETICAL FRAMEWORK.

In this section we summarize the theoretical framework of our analysis. We argue that adjustments to a firm's price schedule can be costly[11]. Therefore, a firm contemplating a change in its nominal prices –in order to adjust them to changing relative prices, will have to face a menu cost *C*. Borrowing from the existing literature we can describe the price-adjustment process as follows.

We consider an economy with a continuum of small monopolistically competitive firms, organized into a series of industrial sectors. Within each industrial sector, firms face heterogeneous menu costs, distributed across firms with distribution function $G(\cdot)$ Some firms can adjust their prices frequently, i.e. with little cost, while some others are more sluggish, i.e. face a larger cost. One would expect economy-wide nominal shocks to impact industrial sectors heterogeneously, too. The shock θ has a density function $f(\cdot)$ across industries, and signals the new optimal level of nominal prices. It follows that firms only adjust their nominal prices if the magnitude of the price shock is larger than their menu cost. For simplicity, we assume that before the price shock nominal prices are set optimally and normalized to zero. We can express changes in aggregate inflation π -in logs, then, as a function of the new desired prices that firms want to set θ and the menu cost *C*.

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V. ESTIMATION METHODOLOGY.

This section describes the methodology we will follow for empirical contrastation of our theories. According to the existing literature on menu costs, since firms adjust their relative prices more quickly to large nominal shocks than to small ones, we should find a positive association between inflation and its higher moments. In this paper we focus on the relationship between aggregate inflation and the skewness in relative-price changes[12]. We test two alternative measures of cross-sectional price changes[13]: skewness of PPI inflation in logs s_t^p, and skewness of PPI in levels s_t^P. In all our regressions the left-hand side variable is the PPI monthly inflation rate in log differences p_t. On the right-hand side of the equations are variables describing the distribution of relative-price changes[14]. In order to capture the dynamic features of the price adjustment process we include lagged inflation and lagged skewness terms. Lastly, we turn our inflation equations into Phillips curves by including unemployment. The most generic equations that we estimate take the following form:

$$a(L)p_t = c + b(L)s_t^j + d_t'g + u_t d + e_t$$

where s_t^j is either s_t^P or s_t^p, $a(L) = 1 - a_1 L - a_{12} L^{12}$, $b(L) = b_0 + b_1 L$, d_t is a 2x1 vector of dummy variables, u_t is the unemployment rate, and e_t is the regression error.

Table 2 and Table 5 present our baseline results at the two-digit and three-digit level of PPI disaggregation, respectively[15]. Dynamic versions of our equations are subsequently estimated and their results reported in Table 3 and Table 6[16]. Estimations for the Phillips curves are shown in Table 4 and Table 7.

VI. ESTIMATION RESULTS.

Equations 1 to 6 and 20 to 26 test the basic prediction about the inflationary effects of skewness in relative-price changes. The positive relationship is confirmed and turns out to be highly significant. Structural-break Chow test statistics within the 1% confidence interval[17] warrant the inclusion of a dummy variable to capture the effects of Spain's 1986 admission into the European Economic Community (EEC.) The estimated negative sign of this variable represent the monetary and fiscal policies 'discipline' effect discussed in Section II and is consistent with the findings of Sobczak (1998) and Garcia-Herrero and Pradhan (1998). The R^2 improves significantly and even more so when incorporating one-period dummy variables. These capture the one-time beginning-of-the-year adjustments to government-regulated prices of several utilities (e.g. power, water and natural gas.) Both measures of PPI skewness, in levels and in log differences, perform comparably well. Our baseline regressions show a small degree of influence of skewness on aggregate inflation but the unrefined nature of these benchmark specifications is likely to mask the true nature of the relationship. Finally, reflecting the post-1985 drop in variability levels of cross-sectional skewness shown in Figure 1 we include a highly significant interactive dummy variable.

The basic model is now expanded in equations 7 to 12 to incorporate some dynamic features of the short-run behavior of inflation. To avoid the estimation distortions produced by the serial correlation of inflation we proceed to add one-period and twelve-period lagged inflation into our equations. This 'inflation persistence' shows up in the positive sign of the estimated parameters. Last-period's inflation turns out to be a major determinant of current inflation and even last-year's inflation displays relatively high estimated parametric values. At the same time, the presence of significant menu costs would induce an incomplete price adjustment to past shocks since current adjustments depend not only on current shocks but also on the distribution of initial prices. If our predictions about current inflation depending on both the current distribution of shocks and on past distributions are to be dynamically consistent we should register a negative relationship between current inflation and lagged skewness. This connection, pointed out by Ball and Mankiw (1995,) was not present in their data set for the U.S. economy from 1949 to 1989 but appears robustly in ours. It is worth noticing the positive long-run net effect of skewness on inflation shown in the equations that incorporate PPI skewness in levels and the significant, but close to zero in value, of the lagged PPI inflation skewness.

So far, our analysis has identified three determinants of aggregate inflation: inflation persistence, a 'discipline' effect, and cross-sectional distributional effects (skewness). The existence of inflation persistence, displayed in highly significant lagged values of inflation, is well documented in the literature. Since our estimates record it with the expected sign and magnitude we will not expand on its discussion at this point. The 'discipline' effect, captured through the dummy variables, is derived from institutional factors. The magnitude of this phenomenon is very significant, although particular to the fiscal and monetary efforts of Spain to conform to the requirements of the EEC. Lastly, we identify a strong influence of cross-sectional distributional effects on aggregate inflation. This cross-sectional skewness is very significant when explaining the volatility of inflation (i.e. the R^2 values dramatically increase when skewness is included in the equation.) To what extent this is a result of the particular sample period and/or the 'discipline' effect is not clear. However, in estimating subsample regressions (1975-1985, 1986-2000) we still find strong evidence of a positive, albeit small in magnitude, effect of skewness on inflation. In both subsamples the R^2 remains essentially the same.

Finally, we turn our inflation equations into Phillips curves by including unemployment. A large body of literature on the dynamics of inflation takes the Phillips curve as the main framework to study the contribution of multiple factors towards aggregate inflation. We enrich that line of analysis by incorporating a novel measure of supply shocks in the form of cross-sectional price skewness. The model specification in equations 13 to 19 and 34 to 39 yields the standard results: inflation is positively related to lagged inflation and negatively related to unemployment. Again, past inflation appears to be responsible for most of the variation in current inflation. Our measure of supply shocks performs significantly well and appears with the expected positive sign, though its impact on aggregate inflation is small. By incorporating lagged skewness we complete the description of the inflation-unemployment tradeoff. The range of price stickiness associated with the existence of menu costs contributes to the rise of unemployment and so lagged skewness moves in the same direction as the unemployment rate. Our estimates of the unemployment-inflation sacrifice ratio range from 1.6% to 1.9% (depending on the measure of asymmetries being used) and are consistent with the ones reported by Sobczak (1998) for the 1990-1997 period.

VII. CONCLUSION.

This study uses the analytical framework of the menu cost literature to show how positive and negative supply shocks influenced the overall price level in Spain from 1975 onwards. We show that as a result of menu costs, firms face a range of nominal price adjustment inaction when responding to changes in relative prices. When the cross-sectional distribution of prices is skewed to the right the economy experiences positive aggregate inflation since only large enough adjustments will warrant paying such menu costs. Similarly, when the distribution is skewed to the left we register significant deflation. The menu cost paradigm can provide an explanation for such episodes in the case of Spain. The oil shocks of the 70s, the admission into the EEC and the convergence process towards the EMU can be seen as sequentially negative and positive supply shocks with significant effects on aggregate inflation.

By constructing a data set of monthly observations over 25 years, the first of its kind in this field of research, we have been able to sidestep some of the strongest arguments raised by Bryan and Cecchetti (1999a). Based on the work by Verbrugge (1999) we have eliminated the possibility of small-sample bias being responsible for the robustness of this stylized fact of the literature. Our statistical work achieves, then, solid foundations.

Besides documenting the stylized fact that links the first and third moments of industry price changes we are able to confirm the 'discipline' effect that membership into an economic union brings to fiscal and monetary policies. Sobczak (1998) identifies a similar 'credibility' shock from Spain's incorporation to the EMU as the driving force for the 1996-1997 deflationary episode. Garcia-Herrero and Pradhan (1998) also point to domestic factors as the underlying determinants of inflation. Even though they fail to link the changes in those factors to the entry into the EEC and EMU they implicitly admit their significance by dividing their 1970-1996 sample period into two after 1989. Mauleon and Bara (1996) follow a similar approach and list the 'discipline' effect resulting from the increased openness as one of the three factors driving down inflation during that period. Finally, Ayuso and Escriba (1998) describe in detail how the aforementioned events re-directed Spanish monetary policy and acknowledge their relevance.

Our empirical research documents, also for the first time in the literature, the dynamic features of the price adjustment process under the presence of menu costs. Ball and Mankiw (1995, page 173) could not find in their data set the negative relationship between present inflation and past distribution of shocks that the menu-cost theory points out. On the other hand, Balke and Wynne (2000) did register a robust, although consistently positive, relationship between inflation and the lagged skewness of the distribution of price changes, vaguely attributing to the combination of technology shocks and flexible prices its meaning and significance. Since the negative correlation between inflation and past skewness is significant and consistently present in our results the analysis thus becomes dynamically consistent under the menu cost paradigm.

Lastly, by turning our inflation equations into Phillips curves by incorporating unemployment we contribute to the study of a tradeoff that has been proven difficult to assess in the recent history of Spain. Our results are consistent with the 2.0 sacrifice ratio (in terms of higher unemployment per permanent percentage point reduction in inflation) quoted by Dolado, González-Páramo and Viñals (1999) for the 1964-1995 period. For the 1975-1992 subsample, encompassing the stagflactionary oil shocks of the late 1970s, we report a slightly higher sacrifice ratio (2.5 – 3.0) also consistent with the findings of Dolado, López-Salido and Vega (2000).

A further test of the robustness of our conclusions could be achieved by studying the evolution of aggregate inflation in Spain after the physical introduction of the Euro in 2002 as the common currency in the European Monetary Union. The technical and logistical changes required to undertake this process were very likely to generate significant menu costs in the Spanish economy capable of impacting the aggregate price level in a positive way. The necessary accumulation of enough empirical data in this particular would relegate this study to a future research agenda.

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[1] For extensive literature reviews see Marquez and Vining (1984) and, more recently, Golob (1993).

[2] Fischer (1981) and Fischer (1982,) for example, are frequently referred studies on the relationship bewteen inflation and the variance of price changes.

[3] Vining and Elwertowski (1976), Ball and Mankiw (1995), and Balke and Wynne (2000) for the United States; Dopke and Pierdzioch (2003) for Germany; Amano and Macklem (1997) for Canada; De Abreu Lourenco and Gruen (1995) for Australia.

[4] This debate, on purely statistical grounds, have been continued by the works of Ball and Mankiw (1999) and Bryan and Cecchetti (1999b), although Balke and Wynne (2000) also offer an alternative economic view on the issue.

[5] All figures reported in this section are taken from Lieberman (1995) and Scobie et al (1998).

[6] For an easily readable account of Spain's trade liberalizing episodes see Dehesa et al (1987.)

[7] See Mendez-Carbajo and Thomakos (2003) for a detailed discussion of the impact of the adhesion to the EEC on Spanish productivity growth rates..

[8] The Spanish monthly Producer Price Index (PPI), with its current basic methodology, is not available prior to 1974. We chose to cut off our time series in January 2000 so that our data analysis is limited to the period prior to the physical introduction of the Euro in January 2002. See Mendez-Carbajo (2004) for an extension of this paper's analytical framework to the impact of the single currency on Spanish inflation.

[9] The admission into the European Union required the INE's adoption of certain statistical standards for reasons of data collection homogeneity. At the two-digit level, the PPI was expanded from 25 to 27 sectors and, at the three-digit level, from 93 to 113 sectors. A linked 1975-2000 data set composed of 24 sectors with 1990 as the base year was subsequently constructed and this will be our primary data set.

[10] Descriptive statistics for the complete 1975-2000 sample period are reported in Table 1.

[11] The literature is abundant in the formalization of these costs. See Friedman and Han (1990), Chapter 15 and Chapter 19 for a detailed discussion of these models and an extensive literature review.

[12] We have calculated the third moment of such distribution of prices at two and three levels of product disaggregation. At the two-digit level the 1975-2000 monthly PPI is composed of 24 sectors and at the three-digit level the 1975-1992 monthly PPI is composed of 93 sectors.

[13] Both measures are unweighted due to the lack of consistent weights in the composition of the PPI during the 1975-2000 period -as the elaboration of this price index went under a number of revisions. Note that all the existing literature indistintively employs weighted and unweighted measures and obtains almost identical results in both cases.

[14] Recently, Bryan and Cechetti (1999a) have argued that the observed positive correlation between the mean and the cross-sectional skewness of price changes suffers from small-sample bias. Using Monte Carlo experiments they claim to be able to fully account for the correlation present in the data as a result of the mentioned bias. They conclude that when price-change distributions are asymmetrical on average there will be a small-sample bias in the mean-variance correlation. In such case, one of the stylized facts in the literature of aggregate price behavior would turn out to be the result of defective statistical analysis. The response to this argument by Ball and Mankiw (1999) and Verbrugge (1999) has been twofold. On the one hand they criticize the construction of the Monte Carlo experiments for failing to capture the true nature of the cross-sectional sampling involved in the construction of a measure of aggregate inflation. On the other hand

they argue that the use of monthly data will sidestep the small-sample bias. Since our analysis, for the first time in the literature, employs monthly data Verbugge's (1999) caveat will fit our research. However, this issue remains open to debate (see Bryan and Cecchetti (1999b)) and we shall not try to draw any conclusion at this point.

[15] In Tables 2 to 4 we report our estimation results for the complete 1975-2000 sample period at the two-digit disaggregation level. For purposes of checking the robustness of our findings we replicate the analysis using the 1974-1992 dataset at the three-digit disaggregation level. Those results are reported in Tables 5 to 7 and are identical in interpretation and significance in both cases.

[16] Residual diagnostic tests are also reported. The Breusch-Godfrey LM test with twelve lags indicate the presence of serial correlation in the error term. We also test for the homocedasticity of the disturbance variance by calculating the ARCH statistic. Our results indicate that such feature is not present in the error term. Engle (1982, 1983) amd Cragg (1982) have found evidence of this type of heterocedasticity in inflation time series in which the variance of the forecast error depends on the size of the preceding squared disturbance. Since the modeling of the autoregressive, conditionally heterocedastic features of inflation falls outside the scope of this study we will limit ourselves to the re-estimation -using the ARCH model- of those equations central to our analysis. The results reported in Table 8 should be seen as a robustness check and as a way of obtaining more appropriate standard errors. The interpretation and significance of the estimation parameters is not affected.

[17] See Table 2 and Table 5 with estimation results in Appendix.

APPENDIX: DESCRIPTIVE STATISTICS AND REGRESSION RESULTS.

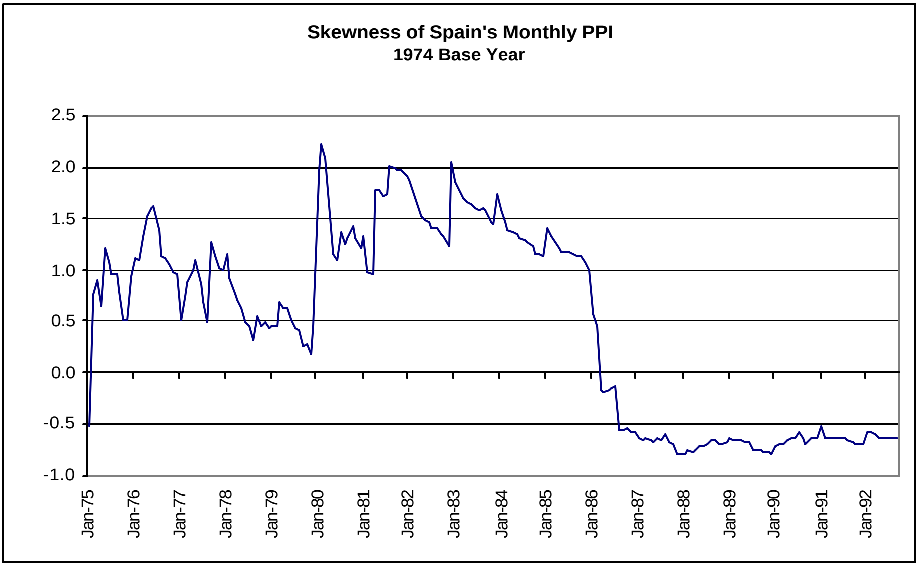

Figure 1. **Skewness of Spain's Monthly PPI. Base Year 1974.**

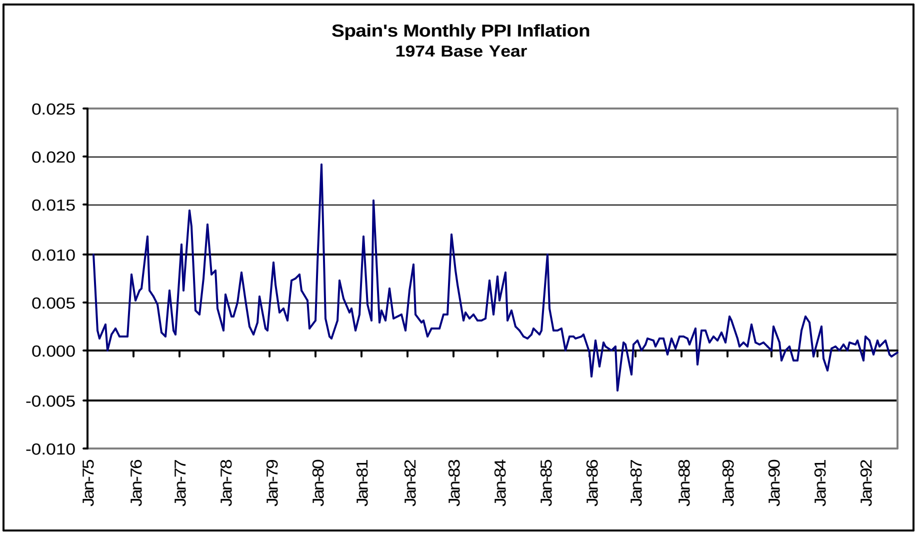

Figure 2. **Spain's Monthly PPI Inflation. Base Year 1974.**

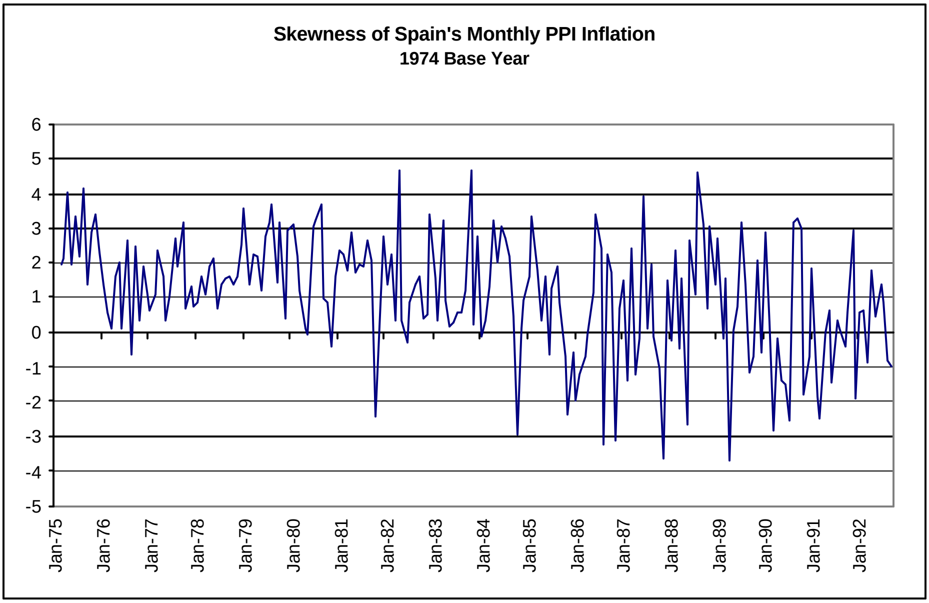

Figure 3. **Skewness of Spain's Monthly PPI Inflation. Base Year 1974.**

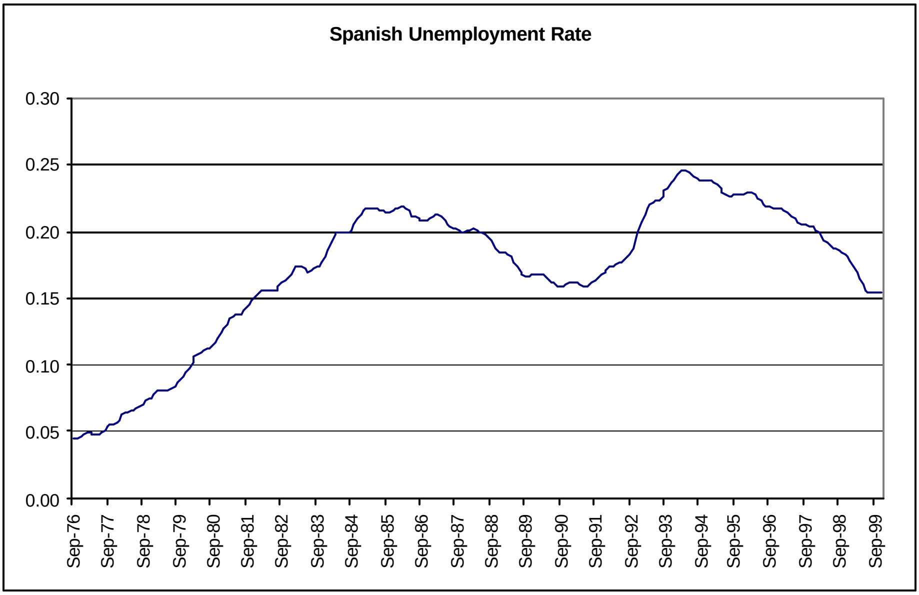

Figure 4. **Spain's Monthly Unemployment Rate.**

Table 1. PPI Inflation, Skewness, and Unemployment Descriptive Statistics.

	Jan. 1975 - Jan. 2000	Jan. 1975 - Jan. 1985 Subsample	Feb. 1985 - Jan. 2000 Subsample
Monthly PPI Inflation			
Mean	0.0056	0.0113	0.0019
Sd. Dev.	0.0071	0.0077	0.0032
Skewness	1.7778	1.4638	0.5560
Kurtosis	7.4904	5.5171	5.6781
Skewness of PPI Inflation			
Mean	0.7672	1.5528	0.2435
Sd. Dev.	2.4063	1.5380	2.7222
Skewness	-0.5148	-0.3710	-0.1647
Kurtosis	2.2451	3.4907	1.8519
Skewness of PPI in Levels			
Mean	0.3993	1.6505	-0.4416
Sd. Dev.	1.3905	0.3720	1.1739
Skewness	-0.2351	-0.2568	0.8393
Kurtosis	1.6380	2.0387	3.1356
Unemployment (*)			
Mean	0.1711	0.1212	0.1994
Sd. Dev.	0.0532	0.0522	0.0262
Skewness	-0.9215	0.0662	-0.1567
Kurtosis	2.9406	1.6679	1.8451

(* sample period: Sept. 1976 - Dec. 1999)

Table 2. Baseline Regressions. January 1975 - January 2000.

Dependent variable: Monthly PPI Inflation

	(1)	(2)	(3)	(4)	(5)	(6)
Constant	0.0046 (12.574)***	0.0100 (13.291)***	0.0096 (14.979)***	0.0048 (11.950)***	0.0089 (12.827)***	0.0090 (15.144)***
PPI Skewness	0.0027 (10.698)***	0.0007 (2.207)**	0.0007 (2.292)**			
Skewness of Inflation				0.0011 (6.846)***	0.0015 (4.784)***	0.0011 (4.103)***
EEC Dummy Variable (85)		-0.0077 (-7.995)***	-0.0077 (-9.538)***		-0.0070 (-8.734)***	-0.0075 (-10.958)***
Beginning of Year Dummies			0.0173 (10.853)***			0.0164 (10.569)***
Interactive Dummy					-0.0011 (-3.018)***	-0.0007 (-2.296)***
R^2	0.2775	0.4055	0.5747	0.1359	0.4563	0.6057
D.W.	1.0606	1.3643	1.3654	0.9359	1.3864	1.3741
s.e.e.	0.0061	0.0056	0.0047	0.0067	0.0053	0.0045
F-test	114.466	101.286	133.343	46.873	82.808	113.267
Chow-test	32.582			87.215		

(t statistics in parentheses: (1%)***, (5%)**, (10%)*)

Residual Diagnostics

LM test (12 lags)	9.576 (0.0000)	3.603 (0.0000)	5.298 (0.0000)	16.305 (0.0000)	3.753 (0.0000)	5.853 (0.0000)
LM test for ARCH (1 lag)	5.363 (0.0212)	4.661 (0.0316)	5.039 (0.0255)	6.225 (0.0131)	3.934 (0.0482)	5.633 (0.0182)

(p-values in parentheses)

Table 3. Regressions with Lags. January 1975 - January 2000.

Dependent variable: Monthly PPI Inflation

	(7)	(8)	(9)	(10)	(11)	(12)
Constant	0.0067 (9.125)***	0.0063 (7.207)***	0.0063 (7.008)***	0.0051 (6.989)***	0.0047 (5.471)***	0.0086 (14.431)***
PPI Skewness	0.0026 (3.920)***	0.0027 (4.162)***	0.0027 (4.266)***			
Lagged PPI Skewness	-0.0024 (-3.559)***	-0.0024 (-3.740)***	-0.0024 (-3.835)***			
Skewness of Inflation				0.0012 (4.819)***	0.0014 (5.579)***	0.0010 (3.904)***
Lagged Skewness of Inflation						0.0002 (1.883)*
Lagged Inflation (-1)	0.3041 (6.823)***	0.2922 (6.398)***	0.2873 (6.337)***	0.3175 (7.424)***	0.2998 (6.971)***	
Lagged Inflation (-12)		0.0806 (1.838)*	0.0860 (1.966)**		0.0761 (1.830)*	
EEC Dummy Variable (85)	-0.0056 (-6.891)***	-0.0053 (-6.020)***	-0.0053 (-5.729)***	-0.0043 (-5.760)***	-0.0040 (-4.843)***	-0.0072 (-10.555)***
Beginning of Year Dummies	0.0163 (11.144)***	0.0157 (10.943)***	0.0159 (11.234)***	0.0161 (11.464)***	0.0154 (11.319)***	0.0162 (10.650)***
Interactive Dummy				-0.0008 (-2.799)***	-0.0009 (-3.439)***	-0.0006 (-2.184)**
Moving Average (3)			0.1147 (1.922)**			
R^2	0.6470	0.6729	0.6781	0.6710	0.7036	0.6138
D.W.	1.8646	1.9102	1.8688	1.9491	2.0481	1.3457
s.e. regression	0.0042	0.0041	0.0041	0.0041	0.0039	0.0044
F-test	107.418	96.336	84.263	119.515	111.147	93.127

(t statistics in parentheses: (1%)***, (5%)**, (10%)*)

Residual Diagnostics

LM test (12 lags)	3.176 (0.0000)	2.480 (0.0042)	3.0 (0.0005)	2.840 (0.0010)	2.229 (0.0108)	6.339 (0.0000)
LM test for ARCH (1 lag)	8.135 (0.0046)	12.051 (0.0005)	11.606 (0.0006)	6.979 (0.0086)	7.962 (0.0051)	6.176 (0.0134)

(p-values in parentheses)

Table 4. Phillips Curves. September 1976 - January 2000.

Dependent variable: Monthly PPI Inflation

	(13)	(14)	(15)	(16)	(17)	(18)	(19)
Constant	0.0125 (9.693)***	0.0125 (11.942)***	0.0096 (8.580)***	0.0086 (7.227)***	0.0124 (13.064)***	0.0087 (8.185)***	0.0076 (6.532)***
PPI Skewness	0.0008 (2.374)**	0.0007 (2.454)***	0.0027 (4.199)***	0.0027 (4.279)***			
Lagged PPI Skewness			-0.0024 (-3.771)***	-0.0025 (-3.953)***			
Skewness of Inflation					0.0005 (5.006)***	0.0006 (5.488)***	0.0006 (5.641)***
Lagged Inflation (-1)			0.2693 (5.925)***	0.2655 (5.874)***		0.2767 (6.292)***	0.2707 (6.194)***
Lagged Inflation (-12)				0.0946 (2.135)**			0.1012 (2.360)**
Unemployment	-0.0183 (-2.123)**	-0.0233 (-3.295)***	-0.0191 (-2.906)***	-0.0193 (-2.951)***	-0.0204 (-2.980)	-0.0162 (-2.503)***	-0.0163 (-2.540)***
EEC Dummy Variable (85)	-0.0065 (-5.626)***	-0.0061 (-6.494)***	-0.0046 (-4.989)***	-0.0039 (-3.925)***	-0.0070 (-9.143)***	-0.0047 (-5.823)***	-0.0037 (-4.166)***
Begn. of Year Dummies		0.0175 (11.762)***	0.0164 (11.733)***	0.0160 (11.431)***	0.0170 (11.768)***	0.0167 (12.324)***	0.0163 (12.031)***
R^2	0.4344	0.6237	0.6790	0.6843	0.6476	0.6921	0.6982
D.W.	1.4537	1.5249	1.9901	1.9702	1.5384	2.0640	2.0511
s.e. regression	0.0054	0.0044	0.0041	0.0040	0.0042	0.0040	0.0039
F-test	70.650	113.941	96.245	84.223	126.316	123.153	105.269

(t statistics in parentheses: (1%)***, (5%)**, (10%)*)

Residual Diagnostics

LM test (12 lags)	2.768 (0.0014)	4.289 (0.0000)	3.252 (0.0000)	2.675 (0.0020)	4.948 (0.0000)	3.391 (0.0000)	2.756 (0.0015)
LM test for ARCH (1 lag)	3.037 (0.0824)	2.276 (0.1324)	11.410 (0.0008)	14.192 (0.0002)	3.358 (0.0679)	11.414 (0.0001)	13.748 (0.0002)

(p-values in parentheses)

Table 5. Baseline Regressions. January 1975 - September 1992.

Dependent variable: Monthly PPI Inflation

	(20)	(21)	(22)	(23)	(24)	(25)	(26)
Constant	0.0004	0.0066	0.0076	0.0076	0.0056	0.0105	0.0099
	(0.501)	(5.338)***	(7.322)***	(6.908)***	(9.542)***	(14.187)***	(16.239)
PPI Skewness	0.0095***	0.0050	0.0031	0.0031			
	(8.904)	(4.160)***	(3.055)***	(2.871)***			
Skewness of Inflation					0.0010	0.0003	0.0002
					(5.074)***	(1.629)*	(1.485)
EEC Dummy Variable (85)		-0.0068	-0.0072	-0.0071		-0.0086	-0.0082
		(-6.559)***	(-8.352)***	(-4.200)***		(-8.957)***	(-10.444)***
Beginning of Year Dummies			0.0226	0.0226			0.0230
			(9.656)***	(9.592)***			(10.175)***
Interactive Dummy				-0.0001			
				(-0.026)			
R^2	0.2741	0.3980	0.5844	0.5844	0.1092	0.3563	0.5703
D.W.	1.1017	1.4382	1.4638	1.4639	1.1475	1.4412	1.4369
s.e.e.	0.0066	0.0061	0.0051	0.0051	0.0074	0.0063	0.0051
F-test	79.287	69.092	97.478	72.757	25.751	57.854	92.005
Chow-test	21.743				39.970		

(t statistics in parentheses: (1%)***, (5%)**, (10%)*)

Residual Diagnostics

LM test (12 lags)	6.650	2.229	3.740	3.730	7.358	2.001	3.7285
	(0.0000)	(0.0117)	(0.0000)	(0.0000)	(0.0000)	(0.0258)	(0.0000)
LM test for ARCH (1 lag)	4.829	2.490	0.0429	0.0421	2.056	2.335	1.118
	(0.0290)	(0.1160)	(0.8360)	(0.8374)	(0.1530)	(0.1279)	(0.2915)

(p-values in parentheses)

Table 6. Regressions with Lags. January 1975 - September 1992.

Dependent variable: Monthly PPI Inflation

	(27)	(28)	(29)	(30)	(31)	(32)	(33)
Constant	0.0064 (6.296)***	0.0056 (4.600)***	0.0054 (4.234)***	0.0068 (7.536)***	0.0067 (8.171)***	0.0059 (5.740)***	0.0058 (5.618)***
PPI Skewness	0.0123 (4.300)***	0.0116 (3.829)***	0.0115 (3.855)***				
Lagged PPI Skewness	-0.0110 (-3.920)***	-0.0108 (-3.615)***	-0.0106 (-3.603)***				
Skewness of Inflation				0.0002 (1.647)*	0.0002 (1.718)*	0.0002 (1.832)*	0.0003 (2.043)**
Lagged Skewness of Inflation				-0.0002 (-0.189)			
Lagged Inflation (-1)	0.2611 (4.958)***	0.2536 (4.719)***	0.2585 (4.839)***	0.2661 (5.043)***	0.2650 (5.063)***	0.2506 (4.672)***	0.2461 (4.669)***
Lagged Inflation (-12)		0.1392 (2.781)***	0.1395 (2.820)***			0.1177 (2.306)**	0.1242 (2.448)***
EEC Dummy Variable (85)	-0.0056 (-6.524)***	-0.0050 (-4.969)***	-0.0048 (-4.588)***	-0.0056 (-6.048)***	-0.0056 (-6.369)***	-0.0050 (-5.080)***	-0.0050 (-4.818)***
Beginning of Year Dummies	0.0177 (7.095)***	0.0177 (7.087)***	0.0183 (7.494)***	0.0234 (10.744)***	0.0233 (10.787)***	0.0230 (10.893)***	0.0233 (11.294)***
Moving Average (3)			0.1297 (1.807)***				0.1377 (1.927)**
R^2	0.6508	0.6767	0.6830	0.6245	0.6244	0.6581	0.6656
D.W.	1.9752	1.9799	1.9407	1.8823	1.8804	1.9261	1.8825
s.e. regression	0.0046	0.0045	0.0045	0.0048	0.0048	0.0046	0.0046
F-test	76.404	67.330	59.088	68.174	85.610	74.667	64.011

(t statistics in parentheses: (1%)***, (5%)**, (10%)*)

Residual Diagnostics

LM test (12 lags)	2.634 (0.0027)	1.740 (0.0615)	2.436 (0.0058)	2.685 (0.0023)	2.690 (0.0022)	1.630 (0.0864)	2.038 (0.0233)
LM test for ARCH (1 lag)	0.043 (0.8358)	0.287 (0.5925)	0.173 (0.6777)	0.0585 (0.8090)	0.0536 (0.8170)	0.9229 (0.3378)	0.759 (0.3846)

(p-values in parentheses)

Table 7. Phillips Curves. September 1975 - September 1992.

Dependent variable: Monthly PPI Inflation

	(34)	(35)	(36)	(37)	(38)	(39)
Constant	0.0106	0.0078	0.0142	0.0134	0.0103	0.0088
	(6.152)***	(5.094)***	(9.583)***	(11.560)***	(7.592)***	(5.990)***
PPI Skewness	0.0053	0.0116				
	(4.303)***	(3.920)***				
Lagged PPI Skewness		-0.0096				
		(-3.280)***				
Skewness of Inflation			0.0004	0.0003	0.0003	0.0003
			(2.236)**	(2.303)***	(2.448)***	(2.347)**
Lagged Inflation (-1)		0.1923			0.2134	0.2068
		(3.545)***			(3.971)***	(3.896)***
Lagged Inflation (-12)		0.1453				0.1276
		(2.901)***				(2.489)***
Unemployment	-0.0365	-0.0245	-0.0311	-0.0307	-0.0255	-0.0247
	(-3.428)***	(-2.989)***	(-2.848)***	(-3.602)***	(-3.064)***	(-3.014)***
EEC Dummy Variable (85)	-0.0040	-0.0029	-0.0065	-0.0060	-0.0043	-0.0033
	(-2.913)***	(-2.578)***	(-5.242)***	(-6.189)***	(-4.234)***	(-3.014)***
Beginning of Year Dummies		0.0181		0.0236	0.0232	0.0232
		(7.484)***		(11.024)***	(11.252)***	(11.404)***
R^2	0.4493	0.6989	0.4110	0.6422	0.6701	0.6807
D.W.	1.5781	2.0438	1.5966	1.6891	2.0607	2.0222
s.e. regression	0.0058	0.0043	0.0060	0.0047	0.0045	0.0045
F-test	51.4054	61.3501	43.953	84.373	75.958	66.091

(t statistics in parentheses: (1%)***, (5%)**, (10%)*)

Residual Diagnostics

LM test (12 lags)	1.473	2.129	1.440	3.232	3.216	2.231
	(0.1379)	(0.0173)	(0.1513)	(0.0003)	(0.0003)	(0.0121)
LM test for ARCH (1 lag)	1.379	1.037	1.075	0.4528	0.5552	1.568
	(0.2417)	(0.3097)	(0.3010)	(0.5018)	(0.4571)	(0.2119)

(p-values in parentheses)

Table 8. ARCH Regressions. January 1975 - January 2000.

Dependent variable: Monthly PPI Inflation

	(40)	(41)	(42)	(43)	(44)	(45)
Constant	0.0101 (18.182)***	0.0085 (15.835)***	0.0061 (8.620)***	0.0045 (7.021)***	0.0082 (9.742)***	0.0072 (9.376)***
PPI Skewness	0.0003 (1.065)***		0.0025 (3.574)***		0.0025 (3.503)***	
Lagged PPI Skewness			-0.0023 (-3.345)***		-0.0024 (-3.463)***	
Skewness of Inflation		0.0011 (6.011)***		0.0011 (6.604)***		0.0053 (5.840)***
Lagged Inflation (-1)			0.3435 (8.605)***	0.3624 (10.395)***	0.3488 (8.987)***	0.3523 (9.321)***
Lagged Inflation (-12)			0.0743 (2.190)**	0.0604 (1.985)**	0.0901 (2.766)***	0.0821 (2.797)***
Unemployment					-0.0181 (-3.546)***	-0.0158 (-3.350)***
EEC Dummy Variable (85)	-0.0081 (-10.78)***	-0.0070 (-11.49)***	-0.0054 (-6.531)***	-0.0036 (-5.728)***	-0.0038 (-4.376)***	-0.0034 (-4.725)***
Interactive Dummy		-0.0007 (-3.569)***		-0.0007 (-3.603)***		
Beginning of Year Dummies	0.0137 (18.94)***	0.0143 (21.92)***	0.0135 (16.56)***	0.0133 (19.26)***	0.0138 (17.76)***	0.0143 (21.13)***
R^2	0.5650	0.6023	0.6723	0.6964	0.6767	0.6911
D.W.	1.3239	1.3431	1.9727	2.1614	2.1116	2.1935
s.e. regression	0.0047	0.0045	0.0041	0.0040	0.0040	0.0039
F-test	76.385	73.984	63.376	79.999	62.817	75.822

(z statistics in parentheses: (1%)***, (5%)**, (10%)*)

Variance Equation

ARCH (1 lag)	0.5388 (4.9)	0.6040 (5.6)	0.3702 (3.5)	0.3953 (4.1)	0.3479 (3.5)	0.3959 (3.8)

(z statistics in parentheses at 1%)

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