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"INFLATION DYNAMICS AND THE CROSS-SECTIONAL DISTRIBUTION OF PRICES IN THE E.U. PERIPHERY"

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INFLATION DYNAMICS &
THE CROSS-SECTIONAL DISTRIBUTION OF PRICES
IN THE E.U. PERIPHERY[†]

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Abstract: We explore the connection between inflation and its higher-order moments for three economies in the periphery of the European Union (E.U.), Greece, Portugal and Spain. Motivated by a micro-founded model of inflation determination, along the lines of the hybrid New Keynesian Phillips curve, we examine whether and how much does the cross-sectional skewness in producer prices affect the path of inflation. We develop our analysis with the perspective of economic integration/inflation harmonization (in the E.U.) and discuss the peculiarities of these three economies. We find evidence of a strong positive relation between aggregate inflation and the distribution of relative-price changes for all three countries. A potentially important implication of our results is that, if the cross-sectional skewness of prices is directly related to aggregate inflation, not only the direction but also the magnitude of a nominal shock would influence output and inflation dynamics. Moreover, the effect of such a shock could be received asymmetrically, even when countries share a common currency.

Keywords: Inflation; Cross-sectional distribution of prices; Greece, Portugal, Spain; European Union; Harmonization.

JEL Classification: E31

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1. 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's and Elwetowski's (1976) seminal paper, different lines of research have examined both the existence of this relationship and its origins[1].

Attention has been concentrated 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 relationship. 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.)

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, concluding 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

[1] For an early extensive literature review see Marquez and Vining (1984) and, more recently, see Golob (1993).
[2] Fischer (1981) and Fischer (1982,) for example, are frequently cited 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.

the use of monthly data will sidestep the small-sample bias. Since our analysis employs monthly PPI data Verbugge's (1999) caveat will fit our research. We will not attempt to settle this issue as Bryan and Cecchetti (1999b) continue to discuss it; however, our results are rather robust for the three countries we analyze and continue to hold both at the individual country level and at a panel level respectively.

The question of the origin of this correlation between inflation and its higher order moments 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 been questioned by Balke and Wynne (2000.) These authors argue that technology shocks are, instead of menu costs, responsible for this empirical regularity.

Our analysis is motivated by the argument that adjustments to a firm's price schedule can be costly[4]. Borrowing from the large body of existing literature, see for example Driffill *et al.* (1990,) we can describe the price-adjustment process as follows. Firms face a cost when adjusting their nominal prices to changes in relative prices. Therefore, a monopolistically competitive firm would change its nominal prices infrequently and only when the magnitude of the required price adjustment equals or exceeds the menu cost. Heterogeneity in menu costs across industries[5] facing a common price shocks or, instead, industry-specific price shocks will promote adjustments of disparate magnitude in relative price levels. In other words, an indicator of the asymmetry of nominal price shocks is likely to contain valuable information regarding the magnitude of the change in the mean value of inflation.

Under a framework of analysis that follows that of Ball and Mankiw (1995), our main contention in this paper is that we expect positive relative price shocks to be positively related to contemporary inflation, while negatively related to future inflation. We examine this claim by studying the relationship between aggregate inflation and the cross-sectional distribution of relative-price changes in the context of three economies in the periphery of the European Union (E.U.), Greece, Spain and Portugal. Their similar

[4] 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.

economic traits and their parallel historical process of admission into the E.U. make a comparative analysis of these three countries particularly interesting. We offer a brief overview in section 2 of the paper. In our analysis we employ an expanded hybrid New Keynesian Phillips curve model, augmented by the presence of contemporaneous and lagged skewness as suggested by the theory. To preview our results, we find strong empirical evidence in all three countries, supporting the hypothesis connecting aggregate inflation with its third cross-sectional moment. Our results are very robust across a variety of different specifications and offer additional material for discussing issues such as inflation convergence in the context of the E.U. In addition, they have important implications for the new member states of the E.U. as well as for prospective members, such as Bulgaria, Romania and Turkey. To the extent that nominal rigidities, perhaps in the form of menu costs, are prevalent in these countries the impact of a common monetary policy on their price-adjustment processes is likely to be significantly different from that experienced by older E.U. members[6].

The remainder of the paper is organized as follows. In section 2 we offer a brief overview of the historical economic developments in Greece, Portugal and Spain. In section 3 we give an outline of the theoretical motivation and empirical model that we use for our analysis. In section 4 we summarize and discuss our data and offer a first glimpse into the skewness-inflation relationship. In section 5 we discuss our estimation results. Section 6 has some concluding remarks and suggestions for further research. Figures and Tables are to be found at the end of the paper.

2. GREECE, PORTUGAL & SPAIN: A BRIEF OVERVIEW

Greece, Portugal and Spain were late additions (in this order) to the pre-Euro European Union. Both the Spanish and Portuguese applications were finally accepted in 1977, within four months of one another; the Greek application was reactivated in 1974. In all these countries, slow and difficult political transitions from authoritarian regimes to full-fledged democracy hampered their admission to the European common market. At the

[5] Dhyne *et al.* (2006) found that in the E.U. the most frequently adjusted retail prices are those of energy and unprocessed food items; processed food, non-energy industrial goods, and services change less frequently.

time of admission (Greece in 1981, Spain and Portugal in 1986) their per capita GDP was at the bottom of the income distribution in Europe: in 1986 Spain and Greece barely exceeded 70 percent of the average of the 12 members and Portugal trailed with 60 percent[7]. In addition, they displayed significant positive inflation differentials (with respect to inflation in the core E.U. countries[8]): these were 10 percent for Greece, 6 percent for Portugal and Spain, based on the consumer price index.

During the following decade, much faster economic growth than the E.U. core helped to close part of the income gap for Spain and Portugal, while Greece receded. Remarkably, this growth was achieved while adhering strictly to the Maastricht treaty, which set stringent conditions for participation in the European Monetary Union[9]. Between 1994 and 1997, Portugal and Spain drove their budget deficits and debt levels to parity with the rest of the Union. Austerity came late to the Greek economy and the country was not included in the first wave of Euro members. With regards to price behavior, the convergence criteria reduced the inflation differentials between the periphery and the core: in Spain and Portugal to less than 1 percent, in Greece to 5 percent. These would be all-time minimums and encompass the first years of our sample period.

At the time of the launching of the Euro, January of 1999, the periphery had closed in on the core's income levels: Portuguese per capita income was 74 percent of the E.U. average; Spain's was 83 percent, and Greece's 66 percent - regaining some of the ground lost in the last decade. Paradoxically, this stage of the process of European monetary unification marked the end of the low inflation period for Spain and Portugal. When, after two years of stringent economic reforms, Greece joined the euro with the second wave of E.U. countries in 2001, its inflation rate exceeded the Euro zone's average by only 1.2 percent. Simultaneously, Spanish and Portuguese inflation

[6] See Dhyne *et al.* (2006) for an extended discussion of the potential reasons behind nominal rigidities across selected European countries.
[7] All figures of per capita income are from the European Commission.
[8] The core countries in the Exchange Rate Mechanism (ERM) and later the European Monetary Union (EMU) are widely considered to be Germany, France and Austria. Despite faster growth between 1996 and 2000 in the periphery, per capita GDP per hour of work in 2001 was only 63 percent of the core's value.
[9] Detailed accounts of the convergence process, towards achieving the Maastricht criteria, abound; see for example Detragiache and Hamann (1999)

differentials were on the rise, above the core's average by 1.5 and 3 percentage points, respectively.

The physical introduction of the Euro accompanied the process of real income per capita convergence for Greece and Spain, while during this time Portugal receded. By 2005 Spain had reached 90 percent of the EU-15 average, while Greece exceeded 76 percent; Portugal shrank to 66 percent. The Greek fast expansion widened the inflation gap to 2.7 percentage points, while the Spanish economy registered inflation rates 1.2 percentage points above the Euro-zone average. Only the 2001-2002 contraction in the Portuguese economy brought inflation rates to full parity.

The fact that during the last two decades inflation rates in these three countries of the EU periphery have remained consistently above the Euro-zone average presents a much debated question[10]. By studying the dynamic characteristics of inflationary processes in Portugal, Spain and Greece[11], employing a hybrid New Keynesian Phillips curve and incorporating a measure of price dispersion as a potential signal of asymmetric nominal shocks we believe we contribute to this ongoing discussion.

3. THEORETICAL AND EMPIRICAL FRAMEWORK

In this section we present in brief a standard formulation of the hybrid New Keynesian Phillips curve, sometimes referred to as the structural inflation equation. Then we integrate the measure of inflation skewness as one of its building blocks.

As in Calvo (1983) we assume that nominal individual prices are not subject to continuous revisions. The price-setting monopolistically competitive firms face adjustment costs that make these frequent price changes unfeasible[12]. As a result, only a fraction χ of all firms would revise their nominal prices at time t. The process of price adjustment will then depend on (a) the difference between the current and desired price

[10] While this paper doe not aim to compare the inflation dynamics in the E.U. periphery to those of the E.U. core per se, Busetti *et al.* (2006) document periods of inflation convergence and divergence between these two regions during 1980 and 2005.

[11] Garganas and Tavlas (2001) provide a detailed analysis on inflation performance during the post 1975 period for Greece.

[12] Even if individually small, these adjustment costs can generate significant aggregate nominal price rigidities. See Akerlof and Yellen (1985) and Mankiw (1985).

level[13] and (b) the gap between the actual and desired firm markup. This yields the familiar expression[14]:

$$p_t^* = p_t + \alpha y_t \tag{1}$$

where p_t^* is the desired price level, p_t is the actual price level, and by normalizing potential output to zero, y_t is the output gap at time *t*. Unless the magnitude of the price revisions exceeds the adjustment cost nominal prices are left unchanged. In terms of inflation rates, $\pi_t = \Delta p_t$, (1) can be expressed as:

$$\pi_t = \beta y_t + E(\pi_{t+1} \mid F_t) \tag{2}$$

where $\beta > 0$ and is determined[15] by both χ and α and $E(\pi_{t+1} \mid F_t)$ is the expectation conditional on time-*t* information of inflation.

In the above formulation inflation expectations $E(\pi_{t+1} \mid F_t)$ play a significant role in the determination of inflation. However, there is no independent role for lagged inflation. Multiple authors argue for the inclusion in the structural inflation equation of either past values of inflation or a combination of forward-looking and backward-looking elements. The choice of proxy variable which adequately captures the inflationary pressures of the output gap is also subject to extensive debate. Traditional proxies of the output gap employ de-trended computations of GDP and the unemployment rate, as well as multiple incarnations of the non-accelerating inflation rate of unemployment (NAIRU). More recently, an aggregate measure of the real marginal cost, also referred to as the labor income share, have gained momentum in the literature[16].

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[13] All references to levels are for the natural logarithms of the corresponding variables.

[14] Since marginal cost rises with increased demand, monopolistically competitive firms would like to increase their prices when the economy expands. In other words, the gap between the actual and desired firm markups is usually expressed as the output gap, the difference between actual output and its natural rate.

[15] See Calvo (1983) or Mankiw (2002) for the algebraic derivation of this coefficient.

[16] Rudd and Whelan (2007) offer a critical, and exceedingly accessible, review of recent contributions to this topic, as well as several others regarding the use of rational expectations sticky-price models to capture inflation dynamics.

For example, Gali *et al.* (2001) consider the following variation of the traditional staggered contract model for the European Union:

$$\pi_t = \gamma_b \pi_{t-1} + \lambda mc_t + \gamma_f E(\pi_{t+1} \mid F_t) \qquad (3)$$

where mc_t is the real marginal cost, computed as the labor income share, and λ is the discounted fraction of firms which in any given period can reset their prices and choose them optimally (i.e. on the basis of expected future marginal costs)[17]. Common labor and business practices, such as wage and price indexation, represent examples of backward-looking behavior in price setting. We argue that in all three cases examined past inflation, π_{t-1}, is a candidate proxy for inflation expectations[18]. At the same time, and in order to check the robustness of this assumption, we estimate, along the lines of Gali *et al*. (2001), a hybrid new Phillips curve incorporating both past inflation and forward looking inflation expectations. Gali *et al.* have shown that in the Euro area inflation dynamics display a strong forward-looking component. Although both sample periods coincide comparisons between works are limited due to the fact that we employ different measures of inflation and to the lack of individual Euro area country-level estimates.

As mentioned above, the parameterization of the output gap has recently added new candidate measures[19]. Although the fact that the countercyclical behavior of the labor share of income complicates the theoretical argument justifying its use as a measure of marginal cost, Gali *et al.* (2005) argue that its empirical contribution is robust[20]. At the

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[17] The fraction of firms able to reset prices but following a suboptimal rule of thumb consider the average of newly adjusted prices last period plus an adjustment for expected inflation, based on lagged inflation. Gali *et al.* (2001) estimate a closed form of this hybrid new Phillips curve for the European Union (1970-1998) and find that forward-looking behavior is dominant.

[18] For Spain: see Alvarez and Hernando (2005) for an extensive discussion of the pricing behavior of retailing firms, and Dolado *et al.* (2000) and Sobczak (1998) for applications of this approach. For Portugal: see Martins (2006) for a survey-based analysis of firms' pricing behavior, and Angeloni *et al.* (2003) for a discussion of the implications of this behavior on monetary transmission. Unfortunately no such study is available for Greece However, studies that discuss inflation persistence, monetary policy and exchange rate regimes belong to Hondroyannis and Lazaretou (2004); Hall and Zonzilos (2000); Alogoskoufis, Lee and Philippopoulos, (1998); Lazaretou (1995).

[19] See Rudd and Whelan (2007) for a discussion of the use of the labor income share as a proxy of the output gap. See Orphanides *et al.* (1999) for a discussion of the traditional operational definitions of the output gap: the difference between the current unemployment rate and the NAIRU and the difference between actual GDP and an estimation of potential GDP.

[20] Among other issues, these authors parry the claims of (a) Linde (2005) regarding how estimating a hybrid New Keynesian Phillips curve with full information maximum likelihood (FIML) yields results

same time, traditional measures of the output gap perform very well in conventional econometric estimation. We don't attempt to settle this issue here and choose to employ an instrument uniformly defined across the three economies object of our study.

The gap between the actual unemployment rate and the NAIRU is a widely used robust approximation to the output gap and therefore a good candidate for this study. Camba and Rodriguez (2003) show that in the E.U. case such a measure performs well relative to other candidate measures. At the same time Estrada *et al*. (2000) find that the usefulness of the NAIRU when discussing Spanish macroeconomic policy is very limited and that the NAIRU is indeed closely matched by actual unemployment[21]. Based on their findings we will proxy the output gap, y_t, by measures of the unemployment rate, u_t: an expansion (an incipient positive output gap) will be associated with a fall in the overall unemployment rate; a contraction (an incipient negative output gap) will be associated with a rise in the overall unemployment rate.

Putting them all together, a stylized, compact econometric representation of such a hybrid New Keynesian Phillips curve widely present in the literature could then be given as:

$$\pi_t = c^* + \gamma_b \pi_{t-1} + \gamma_g u_t + \gamma_f E\left(\pi_{t+1} \mid F_t\right) + \varepsilon_t^* \tag{4}$$

with ε_t^* an appropriately defined exogenous price shock.

Our contribution starts with equation (4) and adds to it various measures of the cross-sectional distribution of relative price changes and a couple of controla variables. The inclusion of additional variables in such 'structural' inflation equation is not explictly warranted by the theory; it is, however, guided by the arguments in Ball and Mankiw (1995) and can be seen as a theoreticaly-motivated test for the marginal predictive ability of the cross-sectional distribution of relative prices. We thus explore the association between inflation and its higher moments by focusing on the relationship between aggregate PPI inflation and the skewness in relative-price changes, which we denote

superior to the Generalized Method of Moments (GMM) and (b) Sbordone (2005) who proposes a two-step minimum distance estimation procedure to test whether *expected* future marginal costs drive inflation.

by s_t^π and define in the next secition. Note that the existing literature indistinctively employs weighted and un-weighted measures and obtains almost identical results. In all our empirical models the dependent variable is the (appropriately defined) PPI-based monthly or annual inflation rate π_t. As noted, on the right-hand side we have variables describing the distribution of relative-price changes. In order to capture the dynamic features of the price adjustment process, and in accordance with the model of equation (4), we include lagged inflation and lagged skewness terms. The most generic equation that we estimate then takes the following form:

$$\gamma_b\left(L\right)\pi_t = c + \gamma_g u_t + \gamma_f E\left(\pi_{t+1} \mid F_t\right) + \beta_s(L)s_t^\pi + \beta_v v_t + x_t\beta_x + \varepsilon_t + \theta\varepsilon_{t-1} \tag{5}$$

where $\gamma_b(L) = 1 - \sum_{j=1}^{p}\gamma_{bj}L^j$ is a polynomial in the lag operator L for inflation; $\beta(L) = \beta_0 + \beta_1 L$ takes into account current and lagged skewness; v_t is the cross-sectional standard deviation of inflation, also defined in the next section. Finally, x_t is an oil-inflation based control that we add for robustness. The model error ε_t can take the form of a moving average when forward looking expectations are included in the model.

We perform our analysis with the following specifications (and corresponding parameter restrictions):

- Without including inflation expectations $E\left(\pi_{t+1} \mid F_t\right)$ and unemployment; we thus have $\gamma_f = \gamma_g = \theta = 0$; we estimate the corresponding model twice, once by least squares (LS) and once by two stage least squares (2SLS) to account for possible endogeneity problems from the presence of the contemporaneous skewness and standard deviation variables.
- With inflation expectations but without unemployment; we thus have $\gamma_g = 0$; we estimate the model by 2SLS.
- With inflation expectations and unemployment (full model); we estimate the model again by 2SLS.

[21] Estrada *et al.* (2000) compute and compare several empirical definitions of the NAIRU and conclude (p. 26) that all but one of the NAIRU estimates shows a time-pattern in five-year sub-periods that is quite similar to the observed unemployment.

2SLS estimation is performed using lags of the dependent and explanatory variables as instruments. We follow estimation with a test for the long-run effect of skewness, namely a test for the hypothesis $H_0 : \beta_0 + \beta_1 = 0$. Finally, we perform a variety of residual diagnostic tests, including tests for residual autocorrelation, conditional heteroscedasticity, normality and functional (miss)specification.

4. DATA & DESCRIPTIVE STATISTICS

Our data were obtained from the respective statistical service agencies of the three countries we examine. We have data on the components of the Producer Price Index (PPI) as well as the index itself, and data on unemployment where they are consistently available.

The sample sizes are almost comparable, starting from 1995 and ending in 2003 for Greece and 2006 for Portugal and Spain. The data differ, however, in the classification digits for the components of the PPI: for Greece we have data for 4-digit classification, for Portugal we have data for 3-digit classification and for Spain we have data for 2-digit classification. Our choices were dictated by reasons of data consistency and availability. For all three countries there were frequent revisions, before 1995, both in base years and methods of aggregations and reporting on the components of the PPI. Unemployment was available on a quarterly frequency for Portugal and Spain but for Greece is was consistently available after 1998. The unemployment data for all three countries were interpolated to monthly frequency using a cubic spline method.[22]

The same procedure for all countries was followed in constructing the two measures of the cross-sectional distribution of relative prices, standard deviation and skewness. The industry-level data were arranged in a $T \times N$ matrix, of t = 1, 2, …, T months and j = 1, 2, …, N industries per month. The definition of the cross-sectional moments we use is given in the following equation:

$$v_t^2 = \frac{1}{N}\sum_{j=1}^{N}\left(\pi_{tj} - \overline{\pi}_t\right), \qquad s_t^{\pi} = \frac{1}{N}\sum_{j=1}^{N}\left(\frac{\pi_{tj} - \overline{\pi}_t}{v_t}\right)^3 \tag{6}$$

where π_{tj} is the (monthly or annual) inflation of the j^{th} sector for month t and $\bar{\pi}_t = N^{-1}\sum_{j=1}^{N}\pi_{tj}$ is the corresponding cross-sectional mean inflation of the j^{th} sector. Assuming that the industries are uncorrelated for every month in our sample, the above equation provides us with consistent estimators of the degree of dispersion and asymmetry in the distribution of producer prices. Finally, monthly and annual PPI inflation were computed using the standard formula $\pi_t = \log\left(P_t / P_{t-k}\right)\times 100$, for k = 1,12.

In Figures 1 through 3 we have a visual presentation of our data series for all three countries, in the order Greece, Portugal and Spain. The figures contain the monthly and annual inflation rates, the corresponding cross-sectional monthly and annual standard deviations and skewness of the PPI and, finally, unemployment. In Tables 1 and 2 we present some distributional and temporal descriptive statistics for all these series.

In Figure 1 we have the monthly and annual PPI inflation and the PPI skewness series for Greece. As usual, the path of monthly inflation exhibits less persistence than the path of annual inflation, something also presented in Table 2.[23] Annual inflation and the cross-sectional skewness move almost together, falling until the late 90's and then increasing before "stabilizing" after 2001. The average inflation rates for the whole period are 0.26 percent (monthly) and 3.41 percent (annual) respectively. It is noteworthy that skewness turned from positive to negative in the period (of about) 1996 to 1998 and then sharply increased to positive again. The evidence in the literature points out to the deflationary impact of the convergence criteria set by the Maastricht treaty. Afterwards, and even within a common currency framework, such stringent macroeconomic constraints were missing. The contemporaneous correlation between monthly and annual inflation and skewness is 43 percent and 63 percent respectively, suggesting that there is some substantial linear dependence between them. Estimating two simple linear regressions of monthly and annual inflation on the cross-sectional skewness we obtain the following results:

[22] The cubic spline matched the last monthly observation within a quarter to the corresponding actual quarterly observation. Further details about our data are available on request.
[23] We note that for all countries annual inflation exhibits strong persistence at low lags but it decays rapidly, in contrast to unit-root nonstationary behavior. Although we provide p-values for a standard unit root test that do not reject the null of nonstationarity we believe that a target variable like inflation cannot

Greece – Monthly PPI Inflation $\pi_t = 0.20 + 0.09 s_t^{\pi},\ R^2 = 18.58\%$ (7a)

Greece – Annual PPI Inflation $\pi_t = 2.70 + 0.90 s_t^{\pi},\ R^2 = 40.33\%$ (7b)

The estimated coefficients are significant at the 1 percent level, for both equations[24], lending some initial support to the potential relationship between inflation and its third moment. The finding that 20% (40%) of the variability of monthly (annual) inflation can be potentially be explained by its cross-sectional skewness is noteworthy and accords with the original predictions of Ball and Mankiw (1995).

In Figure 2 we have the monthly and annual PPI inflation and the PPI skewness series for Portugal. As before, the path of monthly inflation exhibits less persistence than the path of annual inflation. Similar comments to the case of Greece apply here, although there are differences in magnitudes. For example, the rise in annual inflation after 1999 is over two times that of Greece's (Greece peaks at about 7 percent while Portugal peaks at about 16 percent). The average inflation rates for the whole period are 0.22 percent (monthly) and 2.54 percent (annual) respectively. These are comparable to the corresponding averages for Greece. However, Portuguese inflation is more volatile during this period, the historical standard deviations being higher compared to Greece: 0.72 percent vs. 0.65 percent for monthly inflation and 4.26 percent vs. 2.23 percent for annual inflation. The contemporaneous correlation between monthly and annual inflation and their corresponding skewness measures is 62 percent and 82 percent respectively; this is higher than the corresponding numbers for Greece which are 43 percent and 63 percent respectively. Estimating, as before, two simple linear regressions of monthly and annual inflation on the corresponding skewness measures we obtain the following results:

Portugal – Monthly PPI Inflation $\pi_t = 0.11 + 0.14 s_t^{\pi},\ R^2 = 38.36\%$ (8a)

Portugal – Annual PPI Inflation $\pi_t = 1.31 + 1.54 s_t^{\pi},\ R^2 = 67.64\%$ (8b)

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be characterized as a behaving like a random walk. In addition, the low power of unit root tests combined with policy interventions suggests that we should treat the results of such tests with some caution.

[24] For all simple regressions reported in this section significance is based on either heteroskedasticity or heteroskedasticity-autocorrelation consistent standard errors. Results from 2SLS estimation are similar and available on request.

The estimated coefficients are again significant at the 1 percent level and the R-squared values are higher than in the case of Greece, suggesting an even stronger link between inflation and its third moment. The fit for the annual inflation equation is the strongest between the three countries.

Finally, in Figure 3 we have the visuals for the Spanish series. As before, the path of monthly inflation exhibits less persistence than the path of annual inflation. Similar comments to the case of Greece and Portugal also apply here, although there are differences in magnitude. After the sharp increase in 1999 and 2000 annual inflation then drops and starts increasing again, instead of "stabilizing" as in the case of Greece and Portugal. It is interesting to note that unemployment falls with a definite downward trend from 1995 – contrast this to the rise of Portugal's unemployment after 2000. However, the average unemployment rate for Spain is three times that of Portugal[25]. The average inflation rates for the whole period are 0.21 percent (monthly) and 2.21 percent (annual) respectively, almost identical to those of Portugal. In contrast to Portugal though, the historical standard deviation of inflation for Spain is closer to that of Greece with 0.39 percent (monthly) and 2.25 percent (annual) respectively. The contemporaneous correlation between monthly and annual inflation and skewness is 67 percent and 72 percent respectively. Estimating, as before, two simple linear regressions of monthly and annual inflation on the cross-sectional skewness we obtain the following results:

Spain – Monthly PPI Inflation $\pi_t = 0.16 + 0.10 s_t^{\pi},\ R^2 = 44.84\%$ (9a)

Spain – Annual PPI Inflation $\pi_t = 1.62 + 0.69 s_t^{\pi},\ R^2 = 52.11\%$ (9b)

The estimated coefficients are significant at the 1 percent level respectively and the fit for the monthly inflation equation is the highest among the three countries.

In the next section we present our estimation results for the generic model of equation (5) and discuss the implications of our findings, which are strongly supportive of the simple regressions presented above.

[25] As large as this disparity in unemployment rates may seem, particularly for neighboring countries, Castillo *et al.* (1998) found that both countries experienced similar shocks to their unemployment rates

5. ESTIMATION RESULTS AND DISCUSSION

The core estimation results, based on the model of equation (5), appear in Tables 3 through 9. We present the coefficient estimates along with their significance, the fit of each model and a number of residual diagnostics. Also, the long-run effect of skewness on inflation is also tested and reported. When using 2SLS our instrument lists included past values of (monthly or annual) inflation, the relevant standard deviation and skewness measures and unemployment. The structure of Tables 3 to 8 is the same: they have four column panels that correspond to the four different model combinations we considered and are, in order: the model without forward looking expectations and unemployment, estimated by LS, in column one; the same model estimated by 2SLS in column two; the model with forward looking expectations but without unemployment, estimated by 2SLS, in column three; and, finally, the hybrid Phillips curve model with both forward looking expectations and unemployment, estimated by 2SLS, in column four.

Our results for Greece are presented in Tables 3 and 4. Table 3 has the results for monthly inflation and Table 4 for annual inflation. Starting from the monthly inflation, the signs of the estimated coefficients of contemporaneous and lagged skewness are as anticipated by the theory suggested by Ball and Mankiw (1995): positive for contemporaneous skewness and negative for lagged skewness. The coefficients of contemporaneous skewness in three out of four cases is significant at the 1% level, while it is interesting to note that the coefficients of lagged skewness are significant only when forward looking inflation is included in the model (columns 3 and 4). Since lagged skewness is not significant in the first two models we do not consider the results on the long-run effects for these models. For the models 3 and 4 the results on the long-run effect of the third moment are mixed: when unemployment is not present we reject the null hypothesis of zero long-run effect (column 3) while when unemployment is present we do not reject it.

Lagged inflation is always highly significant (in most cases at the 1% level) and there is ample evidence of strong persistence as well as mean reversion, with lagged inflation alternating signs between lag one and lag two. Forward inflation also enters

from the 1980s onwards. We will follow Blanchard and Jimeno (1995) in approaching the treatment of this question with "humility" and follow their example by leaving this intellectual challenge to others.

significantly but the magnitude of the respective coefficient is smaller than that of the coefficient of the first lag of inflation. Finally, unemployment in the hybrid model in column 4 does not enter significantly.

The results for annual inflation for Greece, in Table 4, are much stronger and robust as far as the effects of skewness are concerned. Now both contemporaneous and lagged skewness enter with the expected signs and are strongly significant. The long-run effect of skewness is present only when lagged inflation is used; it disappears when forward looking expectations are included in the models. Similar comments to the monthly models about lagged and forward inflation apply here and, finally, unemployment now enters with the expected (negative) sign and is significant.[26]

The results for Portugal appear in Tables 5 and 6 and are quite similar to the results for Greece. The results on annual inflation are, again, stronger and more robust on the effect of cross-sectional skewness on relative prices. However, unemployment does not enter significantly in the monthly or in the annual inflation models. Similar comments to the case of Greece apply for the long-run effects of skewness.

The results for Spain appear in Tables 7 and 8 and are closer to the results for Portugal than those of Greece. Past inflation values influence current inflation levels in a very similar order of magnitude in both Spain and Portugal[27]. Unemployment does not enter significantly in any of the models for Spain as well. A result worth noting here is about the magnitudes of the estimated coefficients of contemporaneous and lagged skewness among the three countries: for annual inflation Greece has the highest estimated effect of skewness on inflation while Portugal's and Spain's estimates have almost half the magnitude of those of Greece. This can be a sign of an `idiosyncratic' response of the

[26] Gali and Gertler (1999) showed that the sum of the backward and forward-looking parameters could be smaller than or equal to unity depending on how strongly or weakly price-setting firms discount future prices. In this paper, for simplicity, we consider an implicit discount factor of one. The results for Greece indicate that the sum of these estimates is equal to unity cannot be rejected for both annual and monthly models. The same result applied to Portugal and Spain as well.

[27] The literature indicates that producer prices are revised more frequently than consumer prices; see Alvarez and Hernando (2005) for a discussion of the Spanish case. Adjustments of consumer prices take twice as long in the EU than in the USA: both Gali *et al.* (2000) and Dhyne *et al.* (2006) estimate an average adjustment period in Europe between four and six quarters. Nevertheless there is great variability within the EMU: according to Cecchetti and Debelle (2006) CPI inflation persistence is higher in Italy than in Finland by an order of magnitude of almost 30.

Greek economy to changes in the cross-sectional distribution of prices and, possibly, a sign of common shocks that underlie relative price changes in Portugal and Spain.[28]

An important overall result that comes out of all the models we considered so far is that our skewness estimates appear to be "dynamically consistent", in the sense that we strongly register a negative relationship between current inflation and lagged skewness. This expected relationship was pointed out by Ball and Mankiw (1995) but could not be identified in their data set for the U.S. economy from 1949 to 1989. Its robust presence here, regardless of whether the economy in question is experiencing continuous inflation (i.e. Spain) or a protracted deflation (i.e. Portugal)[29], is suggestive of its potential validity as an "inflation regularity" – at least in the three-country context we consider here.

We consider two extensions to the previously considered models and report them collectively in Table 9. First, we include the change in annual oil inflation as a control variable in the annual inflation hybrid models. The results on lagged and forward inflation and contemporaneous and lagged skewness are unaffected by the inclusion of this additional variable. Second, we pool our data and form a three-country panel equation similar to the hybrid model we had before (but without the oil variable and the moving average). We estimate this panel equation using the Generalized Method of Moments (GMM) with dynamic instruments (that include also the instruments used in the single country equations). The results are surprisingly robust: contemporaneous and lagged skewness enter with the anticipated, dynamically consistent, signs and are strongly significant. In addition, unemployment now enters with the expected negative sign and is significant as well.

All in all, our estimation results are strongly supportive of the theory expounded in Ball and Mankiw (1995) about the effects of the cross-sectional third moment of inflation on relative price changes. We find a significant presence of current and lagged skewness as a determinant of current inflation, even after controlling for a number of variables that enter in more traditional inflation equations (lagged and forward inflation,

[28] This finding may signal an idiosyncratic feature of the price-setting process in Greece, where firms may consider more relevant the nominal price adjustments across industries than the past (or future) realizations of aggregate inflation. In that light, Greek firms would not be as backward, or forward, looking as their Portuguese and Spanish counterparts but rather "lateral-looking". Unfortunately there is no study, to our knowledge, that discusses the price setting behavior of Greek firms.

the cross-sectional standard deviation, unemployment and oil inflation). Our results continue to hold when we pool our data together and consider the effect of skewness on the three countries simultaneously.

6. CONCLUSIONS

In this paper we explore the connection between inflation and its higher-order moments (skewness) for three peripheral E.U. countries Greece, Portugal and Spain along the lines of the menu cost theory of price rigidities in product markets. Our work is among the relatively few studies that appeared in previous literature that explore this topic using a similar approach. We contribute to this line of research by examining three economies that may be emulated by the new E.U. member states or candidate member states, such as Bulgaria, Romania or Turkey, with regard to the effects of monetary and anti-inflationary policies they pursue within the context of the E.U.

Employing an augmented version of the hybrid New Keynesian Phillips curve and monthly data on producer prices, we find a robust short-run impact of the skewness of observed relative prices on aggregate inflation in line with the Ball and Mankiw (1995). Our results are in accordance with the predictions of the menu cost models and importantly, they go beyond the contemporaneous mean-skewness correlation and are "dynamically consistent".

The main policy implications of our analysis stem from the impact that an asymmetric distribution of prices would have on the transmission of monetary policy. When the cross-sectional skewness of prices is directly related to aggregate inflation not only the direction but also the magnitude of a nominal shock would influence output and inflation dynamics. Peersman (2004) has found some evidence that the same monetary policy shocks have different effects across E.M.U. countries (e.g. a stronger price response in Spain and Italy than in Austria and the Netherlands). His "puzzling result" may be the result of an incomplete characterization of the process of inflation dynamics. We propose the inclusion of the cross-sectional skewness of relative prices in the hybrid

[29] Aucremanne *et al.* (2002) also identify this dynamic feature of the skewness-inflation relationship but in the case of Belgium they find that the long-run effect is negative and effectively zero.

New Keynesian Phillips curve as a more complete characterization of the process of price adjustments in the periphery of the E.M.U.

While our results are quite robust for the three countries examined here, it is important to look to a larger group of countries if we want to check whether the effects of skewness on inflation are truly an “empirical regularity”. Enlarging our sample to include countries from the E.U.-15 and E.U.-25 groups as well as using OECD countries and data of a lower (e.g. quarterly) frequency appears to be a fruitful extension of our research.

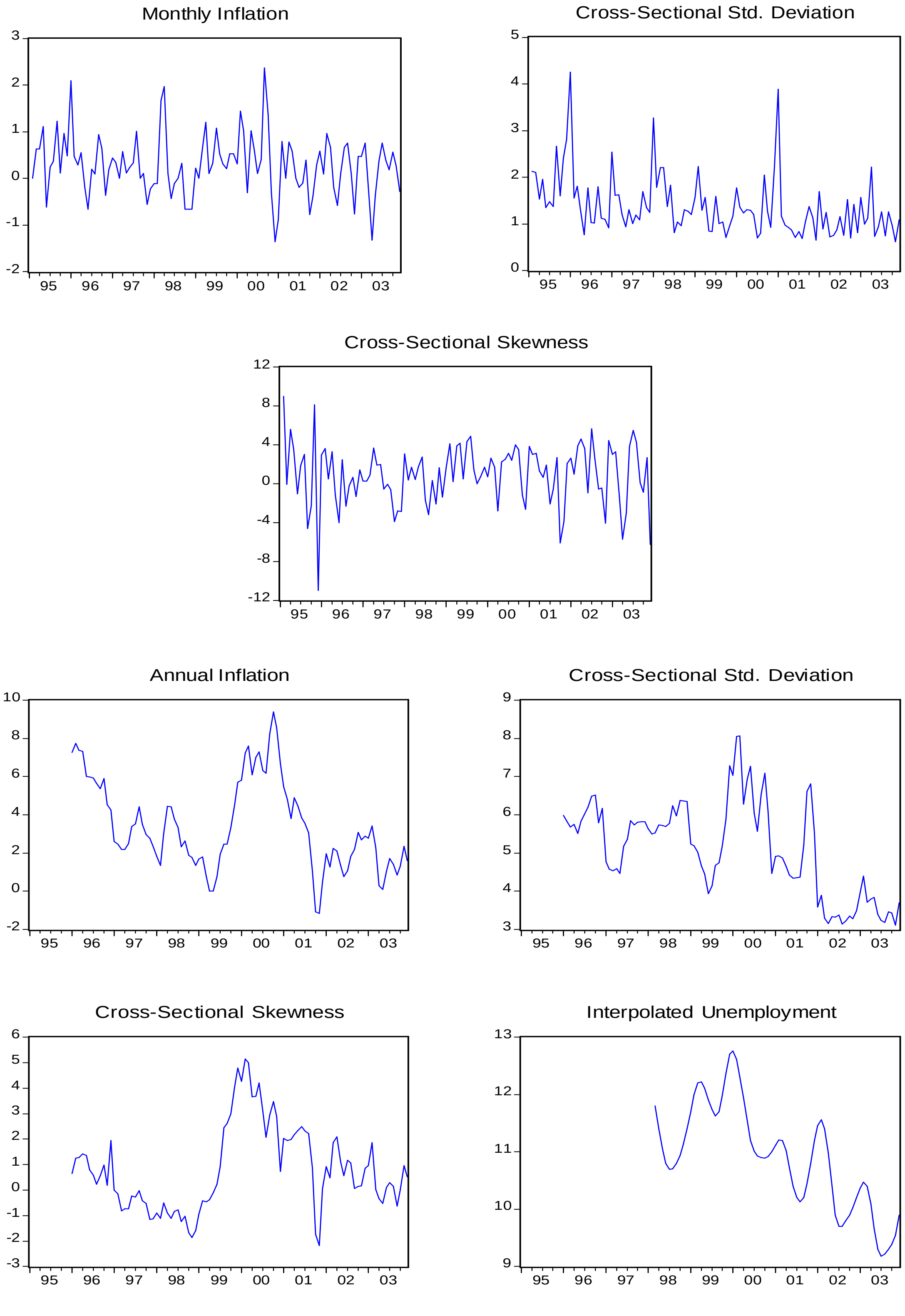

Figure 1. Monthly and Annual Inflation, Cross-Sectional Std. Deviation and Skewness and Unemployment for Greece

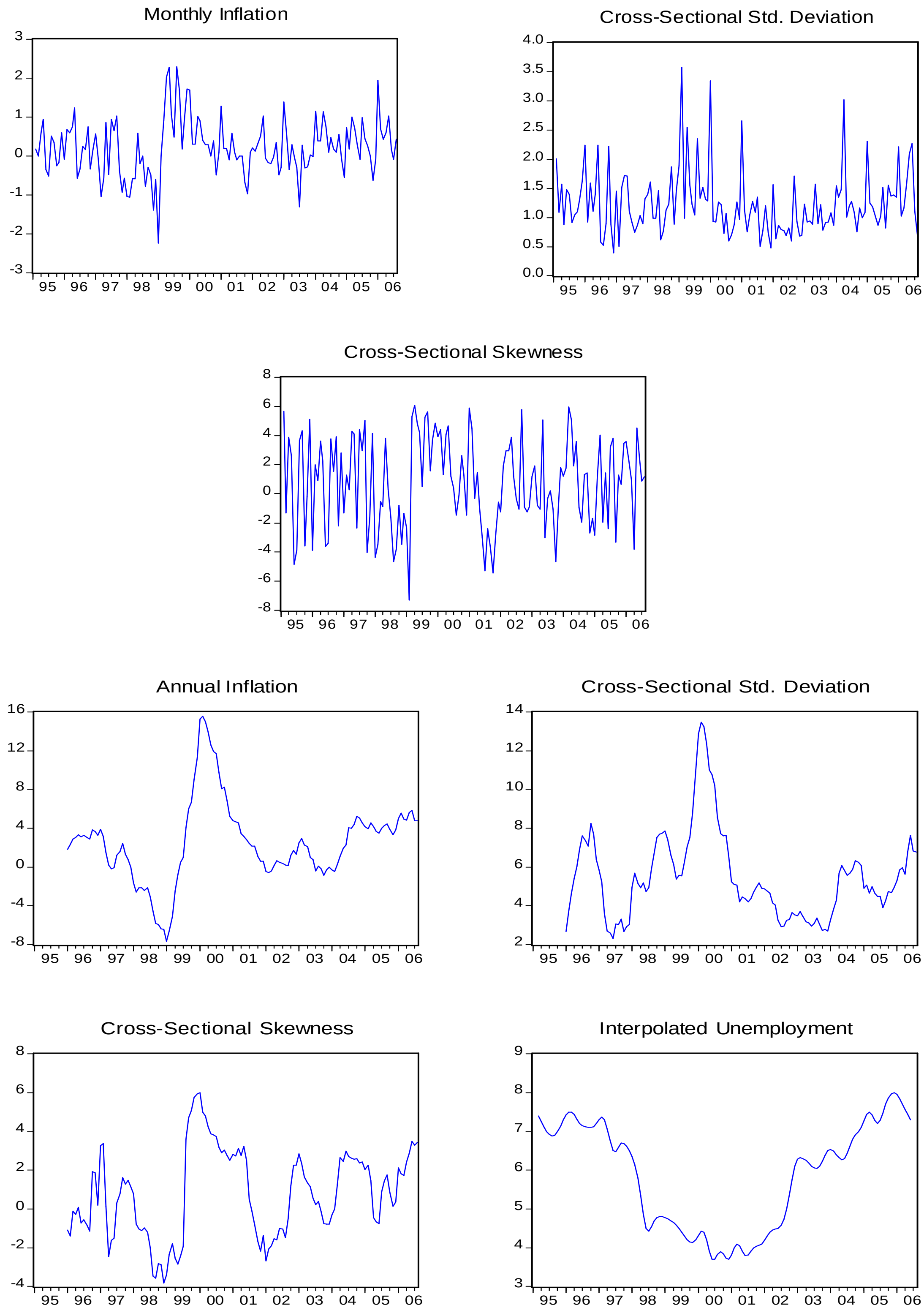

Figure 2. Monthly and Annual Inflation, Cross-Sectional Std. Deviation and Skewness and Unemployment for Portugal

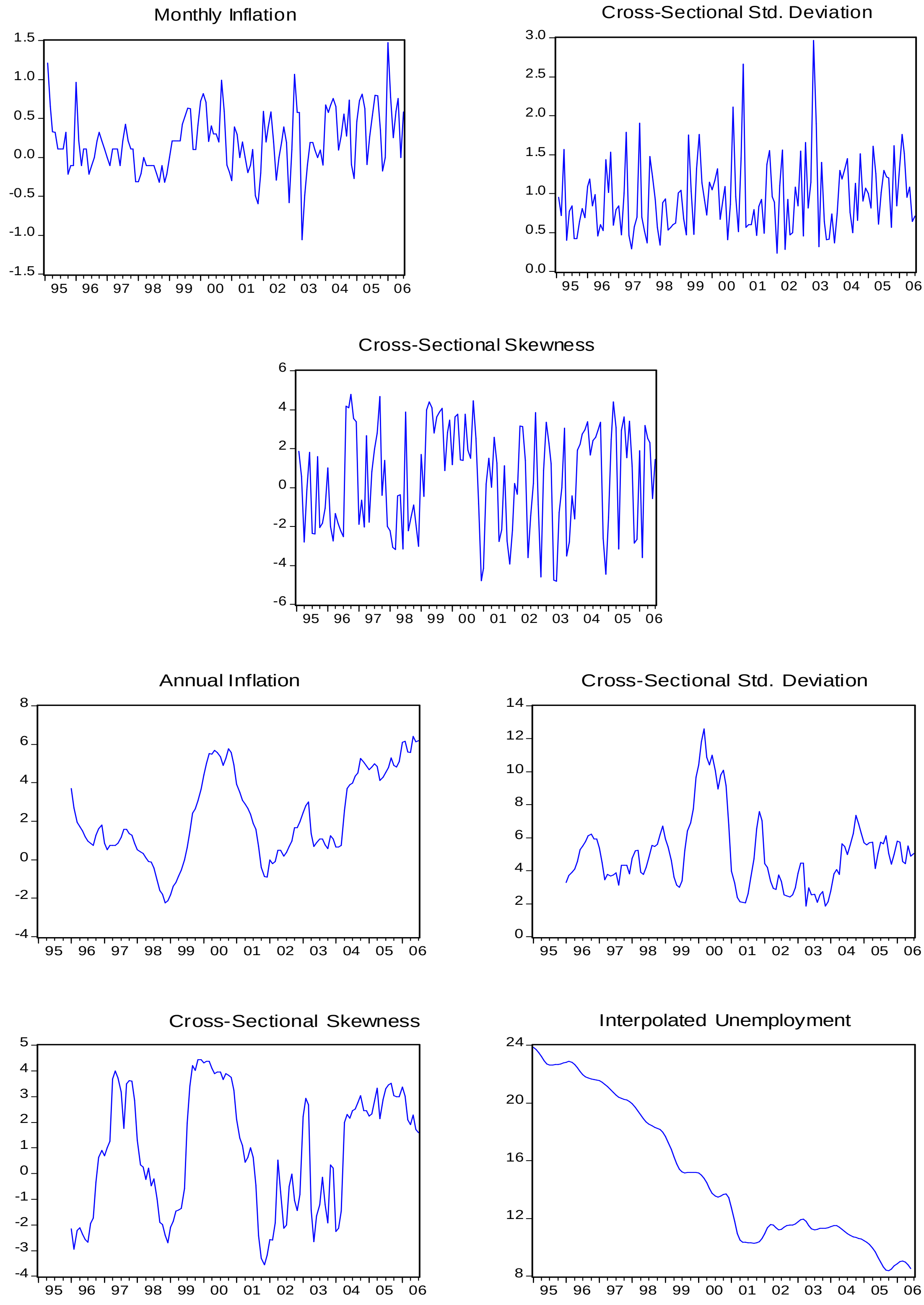

Figure 3. Monthly and Annual Inflation, Cross-Sectional Std. Deviation and Skewness and Unemployment for Spain

GREECE

	Monthly Inflation	Monthly Cross-Sectional Std. Deviation	Monthly Cross-Sectional Skewness	Annual Inflation	Annual Cross-Sectional Std. Deviation	Annual Cross-Sectional Skewness	Unemployment
Mean	0.26	1.31	0.88	3.41	5.10	0.80	10.91
Maximum	2.37	4.26	5.63	9.38	8.06	5.14	12.76
Minimum	-1.36	0.62	-6.28	-1.17	3.11	-2.17	9.18
Std. Dev.	0.65	0.62	2.69	2.33	1.23	1.66	0.91
Skewness	0.45	2.33	-0.60	0.50	0.08	0.67	-0.00
Kurtosis	4.19	10.25	2.81	2.52	2.25	2.87	2.27
Normality	0.01	0.00	0.05	0.08	0.31	0.03	0.46
Observations	96	96	96	96	96	96	70

PORTUGAL

	Monthly Inflation	Monthly Cross-Sectional Std. Deviation	Monthly Cross-Sectional Skewness	Annual Inflation	Annual Cross-Sectional Std. Deviation	Annual Cross-Sectional Skewness	Unemployment
Mean	0.22	1.24	0.74	2.54	5.50	0.80	5.86
Maximum	2.29	3.58	6.07	15.57	13.46	6.00	8.00
Minimum	-2.24	0.40	-7.29	-7.64	2.32	-3.82	3.70
Std. Dev.	0.72	0.55	3.10	4.26	2.28	2.28	1.37
Skewness	0.17	1.63	-0.19	0.59	1.34	0.07	-0.21
Kurtosis	4.07	6.56	2.10	4.52	5.18	2.24	1.49
Normality	0.03	0.00	0.06	0.00	0.00	0.20	0.00
Observations	139	139	139	128	128	128	136

SPAIN

	Monthly Inflation	Monthly Cross-Sectional Std. Deviation	Monthly Cross-Sectional Skewness	Annual Inflation	Annual Cross-Sectional Std. Deviation	Annual Cross-Sectional Skewness	Unemployment
Mean	0.21	0.95	0.47	2.21	5.04	0.87	15.05
Maximum	1.47	2.97	4.80	6.41	12.61	4.43	23.85
Minimum	-1.05	0.24	-4.81	-2.22	1.84	-3.56	8.39
Std. Dev.	0.39	0.47	2.64	2.25	2.23	2.37	4.94
Skewness	0.24	1.19	-0.22	0.18	1.21	-0.15	0.41
Kurtosis	3.59	5.27	1.82	1.93	4.45	1.63	1.64
Normality	0.19	0.00	0.01	0.03	0.00	0.01	0.00
Observations	138	138	138	127	127	127	138

Table 1. Distributional Descriptive Statistics

Notes:

1. The row “Normality” gives the p-value of the Jarque-Bera test for normality in the marginal distribution of the series.

GREECE

	Monthly Inflation	Monthly Cross-Sectional Std. Deviation	Monthly Cross-Sectional Skewness	Annual Inflation	Annual Cross-Sectional Std. Deviation	Annual Cross-Sectional Skewness
r(1)	0.29	0.31	0.03	0.91	0.86	0.89
r(12)	0.07	0.26	0.05	-0.24	0.00	0.12
r(24)	0.00	0.20	0.04	-0.14	0.20	-0.37
ADF	0.00	0.00	0.00	0.13	0.24	0.19

PORTUGAL

	Monthly Inflation	Monthly Cross-Sectional Std. Deviation	Monthly Cross-Sectional Skewness	Annual Inflation	Annual Cross-Sectional Std. Deviation	Annual Cross-Sectional Skewness
r(1)	0.43	0.08	0.22	0.97	0.95	0.89
r(12)	0.10	0.19	0.03	-0.11	0.11	-0.21
r(24)	-0.03	0.02	0.10	-0.42	-0.14	-0.28
ADF	0.00	0.00	0.00	0.18	0.06	0.03

SPAIN

	Monthly Inflation	Monthly Cross-Sectional Std. Deviation	Monthly Cross-Sectional Skewness	Annual Inflation	Annual Cross-Sectional Std. Deviation	Annual Cross-Sectional Skewness
r(1)	0.44	0.09	0.41	0.94	0.92	0.91
r(12)	0.07	0.01	0.02	-0.03	-0.02	-0.26
r(24)	0.01	0.07	-0.07	-0.19	-0.08	-0.14
ADF	0.00	0.00	0.00	0.48	0.04	0.09

Table 2. Temporal Descriptive Statistics

Notes:

1. The rows labeled "$r(h)$" give the sample h^{th} order autocorrelation of the series.
2. The rows labeled *ADF* give the p-value of the Augmented Dickey-Fuller test for a unit root.
3. The 5% standard error of the sample autocorrelations is 0.19 for Greece and 0.17 for Portugal and Spain.

Table 3. Estimation Results for Greece – Monthly Inflation

Explanatory Variables	Model 1 *OLS*	Model 2 *2SLS*	Model 3 *2SLS w/ FI*	Model 4 *Phillips Curve*
Forward Inflation (+1)			0.63***	0.30***
			(0.09)	(0.08)
Lagged Inflation (-1)	0.33***	0.32***	0.71***	0.73***
	(0.10)	(0.10)	(0.08)	(0.12)
Lagged Inflation (-2)	-0.17**	-0.18**	-0.21***	-0.27***
	(0.09)	(0.09)	(0.07)	(0.10)
Current Skewness	0.09***	0.11***	0.03	0.08***
	(0.02)	(0.04)	(0.03)	(0.03)
Lagged Skewness (-1)	-0.005	-0.004	-0.05*	-0.05*
	(0.02)	(0.03)	(0.03)	(0.03)
Current Std. Dev.	0.21**	0.09	-0.10**	-0.06
	(0.09)	(0.17)	(0.05)	(0.08
Unemployment				0.07***
				(0.02)
Moving Average			-0.88***	-0.88***
			(0.11)	(0.13)
R-squared	0.35	0.38	0.64	0.68
F-test	0.00	0.00	0.00	0.00
s.e.e.	0.54	0.51	0.39	0.40
Long-run effect of skewness	0.00	0.00	0.03	0.20
Ljung-Box(24) *test*	0.97	0.96	0.84	0.85
Normality test	0.32	0.29	0.00	0.84
ARCH (24)	0.24	0.46	0.56	0.39
RESET(1) *test*	0.66	0.51	0.00	0.03

Notes:

1. Std. errors in parentheses below the estimates; *, ** and *** denote significance at the 10%, 5% and 1% level respectively.
2. *F-test* gives the p-value for the joint significance of the explanatory variables
3. *s.e.e.* is the standard error of estimation.
4. *Long-run effect of skewness* gives the p-value of the Wald test for the null hypothesis that the sum of the coefficients of current and lagged skewness is zero.
5. *Ljung-Box*(24) *test* gives the p-value of the corresponding test for residual autocorrelation using 24 lags.
6. *Normality test* gives the p-value of the Jarque-Bera test for residual normality.
7. *ARCH*(24) *test* gives the p-value of the corresponding test for residual conditional autoregressive heteroskedasticity using 12 lags.
8. *RESET*(1) *test* gives the p-value of Ramsey's test for misspecification using one fitted term.

Table 4. Estimation Results for Greece – Annual Inflation

Explanatory Variables	Model 1 *OLS*	Model 2 *2SLS*	Model 3 *2SLS w/ FI*	Model 4 *Phillips Curve*
Forward Inflation (+1)			0.37***	0.34***
			(0.05)	(0.29)
Lagged Inflation (-1)	1.07***	1.06***	0.83***	0.94***
	(0.08)	(0.09)	(0.07)	(0.07)
Lagged Inflation (-2)	-0.22***	-0.19**	-0.17***	-0.28***
	(0.07)	(0.08)	(0.06)	(0.05)
Current Skewness	0.71***	0.89***	0.58***	0.42***
	(0.08)	(0.12)	(0.11)	(0.08)
Lagged Skewness (-1)	-0.59***	-0.76***	-0.58***	-0.39***
	(0.09)	(0.13)	(0.10)	(0.07)
Current Std. Dev.	0.14**	0.12*	0.01	0.096***
	(0.06)	(0.07)	(0.04)	(0.03)
Unemployment				-0.19***
				(0.04)
Moving Average			-0.24	-0.79***
			(0.15)	(0.14)
R-squared	0.93	0.93	0.96	0.98
F-test	0.00	0.00	0.00	0.00
s.e.e.	0.6	0.61	0.44	0.35
Long-run effect of skewness	0.02	0.02	0.91	0.48
Ljung-Box(24) *test*	0.61	0.58	0.81	0.89
Normality test	0.64	0.74	0.82	0.93
ARCH (24)	0.61	0.70	0.92	0.27
RESET(1) *test*	0.64	0.46	1.00	0.03

Notes: see Table 3.

Table 5. Estimation Results for Portugal – Monthly Inflation

Explanatory Variables	**Model 1** *OLS*	**Model 2** *2SLS*	**Model 3** *2SLS w/ FI*	**Model 4** *Phillips Curve*
Forward Inflation (+1)			0.37***	0.40***
			(0.08)	(0.06)
Lagged Inflation (-1)	0.24**	0.17*	0.58***	0.52***
	(0.11)	(0.1)	(0.10)	(0.08)
Lagged Inflation (-2)				
Current Skewness	0.12***	0.15***	0.12***	0.10***
	(0.015)	(0.03)	(0.03)	(0.02)
Lagged Skewness (-1)	0.02	0.02	-0.10***	-0.07***
	(0.02)	(0.02)	(0.04)	(0.03)
Current Std. Dev.	0.26***	0.29*	0.07	0.06
	(0.08)	(0.15)	(0.06)	(0.06)
Unemployment				-0.003
				(0.007)
Moving Average			-0.87***	-0.89***
			(0.18)	(0.12)
R-squared	0.49	0.47	0.64	0.68
F-test	0.00	0.00	0.00	0.00
s.e.e.	0.52	0.55	0.45	0.44
Long-run effect of skewness	0.00	0.00	0.52	0.21
Ljung-Box(24) *test*	0.54	0.70	0.93	0.79
Normality test	0.00	0.00	0.15	0.07
ARCH (24)	0.92	0.96	0.87	0.92
RESET(1) *test*	0.09	0.09	1.00	0.00

Notes: see Table 3.

Table 6. Estimation Results for Portugal – Annual Inflation

Explanatory Variables	Model 1 *OLS*	Model 2 *2SLS*	Model 3 *2SLS w/ FI*	Model 4 *Phillips Curve*
Forward Inflation (+1)			0.44***	0.42***
			(0.06)	(0.05)
Lagged Inflation (-1)	1.33***	1.32***	0.68***	0.73***
	(0.08)	(0.08)	(0.15)	(0.14)
Lagged Inflation (-2)	-0.43***	-0.42***	-0.12	-0.15*
	(0.07)	(0.07)	(0.10)	(0.09)
Current Skewness	0.48***	0.55***	0.17*	0.16**
	(0.06)	(0.12)	(0.09)	(0.08)
Lagged Skewness (-1)	-0.35***	-0.41***	-0.15**	-0.15**
	(0.07)	(0.11)	(0.07)	(0.06)
Current Std. Dev.	0.07**	0.06*	0.004	0.01
	(0.03)	(0.04)	(0.01)	(0.02)
Unemployment				0.01
				(0.02)
Moving Average			-0.38	-0.42
			(0.30)	(0.28)
R-squared	0.98	0.98	0.99	0.99
F-test	0.00	0.00	0.00	0.00
s.e.e.	0.67	0.69	0.43	0.42
Long-run effect of skewness	0.01	0.01	0.74	0.77
Ljung-Box(24) *test*	0.22	0.30	0.09	0.08
Normality test	0.00	0.00	0.67	0.59
ARCH (24)	0.03	0.04	0.02	0.03
RESET(1) *test*	0.27	0.28	1.00	0.30

Notes: see Table 3.

Table 7. Estimation Results for Spain – Monthly Inflation

Explanatory Variables	**Model 1** *OLS*	**Model 2** *2SLS*	**Model 3** *2SLS w/ FI*	**Model 4** *Phillips Curve*
Forward Inflation (+1)			0.33***	0.34***
			(0.07)	(0.08)
Lagged Inflation (-1)	0.32***	0.29***	0.64***	0.58***
	(0.05)	(0.09)	(0.09)	(0.11)
Lagged Inflation (-2)				
Current Skewness	0.09***	0.11***	0.1***	0.09***
	(0.01)	(0.02)	(0.02)	(0.01)
Lagged Skewness (-1)	-0.02*	-0.02*	-0.08***	-0.07***
	(0.01)	(0.01)	(0.02)	(0.02)
Current Std. Dev.	0.02	0.04	0.02	0.002
	(0.04)	(0.11)	(0.05)	(0.002)
Unemployment				-0.003
				(0.002)
Moving Average			-0.72***	-0.83***
			(0.16)	(0.19)
R-squared	0.52	0.52	0.66	0.71
F-test	0.00	0.00	0.00	0.00
s.e.e.	0.27	0.27	0.23	0.21
Long-run effect of skewness	0.00	0.00	0.34	0.30
Ljung-Box(24) *test*	0.96	0.96	0.99	0.99
Normality test	0.00	0.00	0.42	0.07
ARCH (24)	0.96	0.80	0.01	0.41
RESET(1) *test*	0.28	0.03	1.00	0.00

Notes: see Table 3.

Table 8. Estimation Results for Spain – Annual Inflation

Explanatory Variables	Model 1 *OLS*	Model 2 *2SLS*	Model 3 *2SLS w/ FI*	Model 4 *Phillips Curve*
Forward Inflation (+1)			0.38***	0.35***
			(0.04)	(0.03)
Lagged Inflation (-1)	1.28***	1.27***	0.84***	0.91***
	(0.07)	(0.08)	(0.08)	(0.08)
Lagged Inflation (-2)	-0.32***	-0.30***	-0.22***	-0.25***
	(0.07)	(0.07)	(0.05)	(0.05)
Current Skewness	0.29***	0.32***	0.08**	0.07**
	(0.03)	(0.05)	(0.03)	(0.03)
Lagged Skewness (-1)	-0.25***	-0.29***	-0.08**	-0.08***
	(0.03)	(0.05)	(0.03)	(0.03)
Current Std. Dev.	0.02	0.02	0.003	0.001
	(0.02)	(0.02)	(0.004)	(0.003)
Unemployment				0.002
				(0.005)
Moving Average			-0.67***	-0.85***
			(0.20)	(0.19)
R-squared	0.98	0.98	0.99	0.99
F-test	0.00	0.00	0.00	0.00
s.e.e.	0.33	0.34	0.19	0.18
Long-run effect of skewness	0.09	0.22	0.65	0.67
Ljung-Box(24) *test*	0.559	0.809	0.00	0.00
Normality test	0.12	0.34	0.04	0.16
ARCH (24)	0.38	0.59	0.03	0.05
RESET(1) *test*	0.5401	0.0889	0.00	0.00

Notes: see Table 3.

Table 9. Estimation Results with Oil Inflation + Panel – Annual Inflation

Explanatory Variables	**Greece** *Phillips Curve*	**Portugal** *Phillips Curve*	**Spain** *Phillips Curve*	**Panel** **Phillips Curve**
Forward Inflation (+1)	0.30***	0.47***	0.37***	0.31***
	(0.03)	(0.05)	(0.04)	(0.05)
Lagged Inflation (-1)	0.98***	0.68***	0.86***	0.86***
	(0.07)	(0.11)	(0.09)	(0.10)
Lagged Inflation (-2)	-0.28***	-0.14**	-0.23***	-0.19***
	(0.04)	(0.07)	(0.05)	(0.06)
Current Skewness	0.35***	0.13**	0.08***	0.44***
	(0.07)	(0.06)	(0.03)	(0.07)
Lagged Skewness (-1)	-0.32***	-0.13**	-0.08***	-0.21***
	(0.07)	(0.05)	(0.03)	(0.05)
Current Std. Dev.	0.13***	0.01	0.00	0.07**
	(0.03)	(0.01)	(0.00)	(0.04)
Unemployment	-0.23***	0.01	-0.00	-0.12**
	(0.04)	(0.02)	(0.00)	(0.05)
Change in Oil Inflation	0.01***	-0.01***	-0.00	
	(0.00)	(0.00)	(0.00)	
Moving Average	-0.76***	-0.47**	-0.77***	
	(0.15)	(0.20)	(0.16)	
R-squared	0.99	0.99	0.99	0.34
F-test	0.00	0.00	0.00	
s.e.e.	0.32	0.40	0.18	0.64
Long-run effect of skewness	0.30	0.82	0.99	0.00
Ljung-Box(24) *test*	0.69	0.04	0.00	
Normality test	0.59	0.60	0.28	0.00
ARCH (24)	0.54	0.09	0.06	
RESET(1) *test*	0.00	1.00	0.00	

Notes: see Table 3.

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