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New Disaggregate Evidence on Spanish Inflation Persistence

Diego Romero-Ávila & **Carlos Usabiaga**

Universidad Pablo de Olavide (Seville, Spain)

Corresponding Author: Carlos Usabiaga, Universidad Pablo de Olavide, Departamento de Economía, Carretera de Utrera, Km. 1, 41013 Sevilla. E-mail: cusaiba@upo.es. Phone: (+34) 954348553. Fax: (+34) 954349338. http://www.upo.es/econ/usabiaga.

Abstract

This article investigates the degree of persistence of different inflation rates for the Spanish economy using the Consumer Price Index (CPI) for the aggregate as well as for the regions, provinces and eight groups of goods and services, in addition to the Producer Price Index (PPI) for the aggregate and 24 industrial sectors. For that purpose, we employ: 1) the unit root tests with good size and power of Ng and Perron (2001) with the small-sample bias correction developed by Perron and Qu (2007); 2) the nonlinear ESTAR unit root test proposed by Kapetanios *et al.* (2003); 3) median-unbiased estimations of the persistence parameter and the respective confidence intervals through the grid-bootstrap method proposed by Hansen (1999); and 4) median-unbiased estimations of the half-life of a shock in addition to the associated confidence intervals through the method based on impulse-response functions proposed by Gospodinov (2004). The results from the application of these techniques indicate that most of the CPI-based inflation rate series clearly contain a unit root. As regards the results for the PPI-based inflation rate series, we have provided evidence that the aggregate series appears to contain a unit root, while at the industry level the inflation rate series are found to be nonlinear stationary in 13 sectors. On the basis of this robust evidence of high persistence in inflation, policymakers should pay more attention to any shock hitting inflation, since the effects are expected to be long-lasting, particularly for consumer prices. Along these lines, it is essential to implement correcting reforms with the aim of raising price adjustment flexibility if one wants to avoid having to intervene actively in the markets to reach the inflation target.

Key words: Inflation, Persistence, Consumer Prices, Producer Prices, Univariate econometric methods.

JEL Classification: C22, E31, E52, E58.

1. Introduction

In the Economic and Monetary Union (EMU) framework, three main issues concerning inflation have been analysed in the last decade by academic macroeconomists and researchers linked to central banks. Moreover, these three topics are much related among them: whether there exists a convergence process among country-specific inflation rates, the explanation of inflation differentials, and the measurement of inflation persistence. As a background for the Spanish economy, which is the subject of analysis, we offer some preliminary information regarding the evolution of inflation.[1] Spanish inflation reached its peak in 1977. Afterwards, inflation gradually decreased until the end of the 1990s. Since then and until the current crisis, Spanish inflation has usually been moderately above the target set by the European Central Bank (ECB). In addition to other problems mentioned throughout the paper, which also affect the Spanish inflation differential, we must stress the strength of domestic demand during this period mainly driven by high indebtedness, slow productivity improvements, convergence in the price level from a position below average and disorientation of consumers towards the single currency. In the 1970s, supply shocks dominated the economic landscape, while demand shocks prevailed during the 1980s and 1990s. More recently, supply shocks have been dominant, until the emergence of the current crisis. Generally, these shocks have been mostly driven by external rather than domestic forces. Within the EMU context, there has been convergence in inflation rates until the end of the 1990s. This process of convergence has stopped during the 2000s. While inflation rates across the Spanish regions and provinces exhibit convergence, there is an opposite pattern for the inflation rates of the groups of goods and services based on the Consumer Price Index (CPI).

Our article mainly focuses on inflation persistence. In our opinion, it is not surprising the great attention paid by monetary policy practicioners to the measurement of inflation persistence. This is because in a scenario with high persistence, the occurrence of a shock could keep the inflation rate far from its equilibrium or target value for a long period of time. For instance, in the case of the ECB, inflation targeting implies a 2% inter-annual inflation objective. Hence, high persistence might force monetary policy intervention to try to return inflation to its target value in a reasonable time span.[2] In this sense, this field is very important from an economic policy perspective. A good sample of the widespread interest in the analysis of inflation persistence can be found in the

[1] On inflation differentials at the European level see European Central Bank (2003); from a Spanish perspective see Estrada and López-Salido (2004) and Restoy *et al.* (2005). Concerning inflation convergence see Caraballo and Usabiaga (2009a), which approaches this topic from a wider perspective –ranging from an aggregated European focus to a high geographical and sectoral disaggregation for Spain, also paying attention to inflation differentials. For a detailed account of the shocks hitting Spanish (and other OECD countries) inflation, see Romero-Ávila and Usabiaga (2009) and Caraballo and Usabiaga (2009c).

[2] Angeloni *et al.* (2006) and Gaspar *et al.* (2007) provide an analysis of the implications of inflation persistence on macroeconomic modelling and economic policy, respectively.

important workgroup created for the ECB to work on these issues. This is called the IPN ("Inflation Persistence Network"), which has generated a large number of research forums and publications. As we will have the occasion to show below, the basic indicator of the IPN to measure persistence – which is the speed at which the rate of inflation reverts to its long-term objective after a shock hits the economy- has guided the approach of our work.

There are different methodological approaches to estimating inflation persistence –see Álvarez (2008) and Gordon (2011). A very important line of research that has rapidly developed during the last decade, with theoretical and empirical contributions, has been linked to the Neokeynesian Phillips curve that extends the framework initiated by Galí and Gertler (1999).[3] In essence, according to this framework in its "hybrid" version, the rate of inflation responds, by means of several parameters that depend on the structural characteristics of the economy, to three types of variables: lagged inflation, indicators of the state of the cycle and future inflation expectations. Thus, the persistence observed in the inflation rate might respond to one or more of these elements. "Intrinsic" persistence, linked to lagged inflation, is the best known component, and captures simply the inertia of the inflation rate due to for example the existence of widespread indexation in the economy**.** "Extrinsic" persistence, linked to indicators measuring the state of the cycle, is associated with the rigidities in the mechanisms of price and wage setting, which also considers important aspects like the frequency of price changes, the quantity of the changes, the change rules, etc.[4] Finally, inflation persistence can also be related to the mechanism of expectations formation by economic agents, which especially confronts the "backward-looking" and the "forward-looking" approaches. In sum, a country characterised by widespread indexation –for example in wages–, rigidities in the mechanism of price determination like long-term "explicit" or "implicit" contracts, coordination failures and menu costs as well as rigidities in wage formation –for instance, due to the model of collective bargaining or unionization–[5], and where "backward-looking" expectations have an important weight, should present a high degree of inflation persistence. In this regard, we must remember a large existing economic literature emphasising the importance of all these factors for

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[3] Galí and López-Salido (2001) apply this methodology to the Spanish economy, whereas Álvarez (2008) relates the existing evidence on the microeconomic behaviour of prices to alternative Phillips curve models. An important alternative framework is that initiated by Mankiw and Reis (2002).

[4] In this respect, see the work by Álvarez *et al.* (2006), Dhyne *et al.* (2006), Fabiani *et al.* (2006) and Vermeulen *et al.* (2007) for the euro-zone. It is usual to compare these results with those for the US, country characterised by lower inflation persistence. See also Álvarez and Hernando (2006) and Álvarez *et al.* (2010) for the Spanish case. Altogether, these papers render insightful findings on the differences in the setting of consumer and producer prices across both economic areas (euro-zone and Spain). This distinction between producer and consumer prices will be also studied in detail in this paper.

[5] On this respect, several papers have concluded that the Spanish labour market follows the implications of the insider-outsider model (Blanchard and Summers, 1986) very closely. Some factors contributing to this fact include, among others, the low geographical mobility of the labour force, the high temporary employment rate and the high and persistent unemployment rate (Romero-Ávila and Usabiaga, 2008).

the Spanish economy.[6] Consequently, we should expect from our analysis an outcome of high persistence in the inflation rate, feature that will be analysed from many perspectives. From another viewpoint, it is often argued that prevalence of the labour input within the cost structure of the firm generates high inflation persistence –particularly for the services sector–, while occurring the opposite with the participation of energy. Finally, there is widespread consensus on the idea that a high degree of intra-sectoral competition generates high price flexibility.

In this paper we focus essentially on what has been called "intrinsic" persistence, which measures the degree of inflation inertia. However, we study this aspect independently from a specific theoretical model by using several empirical approaches based on univariate econometric methods, some of them inspired in the basic definition of persistence handled by the IPN. We would like to emphasise that, unlike other papers –see the survey by Altissimo *et al.* (2006)[7]–, we will pay special attention to the disaggregation of the Spanish inflation rate figures, from a geographical and sectoral perspective, working both with consumer and producer prices provided by the National Statistical Institute (Instituto Nacional de Estadística, INE). We think that this kind of analysis can provide a very precise diagnosis of the problem. Romero-Ávila and Usabiaga (2009) study a similar topic, but for the aggregate inflation figures of a set of OECD countries, using panel data techniques and allowing for structural changes.[8] Nevertheless, for comparison purposes, they also apply univariate tests of the KPSS type without structural change, obtaining evidence of a unit root for Spain, which coincides with the results in this paper for aggregate CPI inflation. In essence, our work tries to contribute to the empirical evidence in this important field, applying a wide battery of econometric

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[6] For instance, Galí and López-Salido (2001) highlight the relevance of the backward-looking component as well as wage frictions, whereas Restoy *et al.* (2005) emphasise the role of wage indexation clauses as well as the dual inflation problems. Caraballo and Usabiaga (2009b, c) find robust evidence for the presence of nominal rigidities in the determination of Spanish consumer and producer prices. Finally, Fabiani *et al.* (2006) also point out that the backward-looking component receives more attention in Spain than in the euro-zone average, possibly due to the higher weight of SMEs in the Spanish aggregate.

[7] We have to point out that the estimates of the persistence parameter that we provide in this work for aggregate CPI and PPI inflation are higher than those reported in the survey of Altissimo *et al.* (2006). However, we have to note that the estimation of this parameter following our univariate methodology is different from the estimation conducted within a wider Phillips curve framework, where more variables and interrelations are involved –see Gordon (2011).

[8] In this work we opt for using univariate tests without structural change to maintain the consistency of the different areas of our empirical analysis and because we are interested in providing information with the highest possible level of disaggregation. This is because the alternative measures of persistence employed (median-unbiased estimates of the persistence parameter and half-life) do not explicitly allow for breaks. In addition, it is important to note that the use of panel methods that pool the data would not allow us to provide such detailed information for each individual series. Another reason for not allowing for breaks in unit root testing here is that, unlike our previous work for aggregate inflation series of 13 OECD countries (Romero-Ávila and Usabiaga, 2009), the focus of the present paper is on providing broad patterns of persistence of inflation rates disaggregated geographically and at the sectoral level. With such a high degree of disaggregation it would be almost impossible to characterise all breaks as well as to find consistent clusters of breaks across different inflation datasets. From the existing evidence for aggregate inflation rate series (see, among others, Romero-Ávila and Usabiaga (2009)), we are also aware that the degree of persistence decreases once breaks are considered in the analysis. However, for the reasons outlined we do not pursue this avenue in this article.

techniques to several databases. Also, we would like to point out that the main aim of our paper is to try to identify general patterns in the data rather than paying excessive attention to the specific results for each particular geographical unit, group of goods and services or sectors.[9]

Previous studies have analysed the persistence of inflation rates by means of unit root tests with the aim of determining whether the null hypothesis –that the sum of autoregressive (AR) coefficients constituting the persistence parameter is equal to unity– can be rejected in favour of the alternative hypothesis of stationarity –where the persistence parameter is lower than one. To assess whether there is a unit root in the various inflation rates we are dealing with –CPI for the aggregate, regions, provinces and eight groups of goods and services; and PPI (Producer Price Index) for the aggregate and 24 sectors–, we employ the unit root tests proposed by Ng and Perron (2001) with the finite-sample bias correction of Perron and Qu (2007). These statistics modify conventional unit root tests to render statistics with good size and power. By using these tests we can be more confident that non-rejections of the null of a unit root are not caused by the low power of conventional unit root tests such as the augmented Dickey-Fuller (1979, ADF hereafter) or Phillips-Perron tests (see De Jong *et al.*, 1992), while rejections are not due to size distortions caused by the presence of large negative moving average roots in the data (see Perron and Ng, 1996). In addition, we employ the nonlinear unit root test of Kapetanios *et al.* (2003) that allows us to detect the presence of nonstationarity in inflation against nonlinear but globally stationary exponential smooth transition autoregressive (ESTAR) processes.[10,11]

However, the information obtained from unit root testing may provide an incomplete picture of the degree of persistence of the inflation rate series, as they focus on testing whether the sum of the AR coefficients is equal or less than unity. Therefore, in order to complement that analysis we provide median-unbiased estimates of the persistence parameter. We also compute point estimates of the half-life of a shock following Cheung and Lai (2000) that recommend computing them

[9] We can find articles in the IPN framework that approach exhaustively (for countries, for sectors, for firm sizes, etc.) this kind of analysis for the euro-zone and Spain, but with empirical methodologies different from ours. In Caraballo and Usabiaga (2009a) it is possible to find information for the Spanish case, at different levels of disaggregation.

[10] As noted by Kapetanios *et al.* (2003), failure to allow for nonlinearities in the data generating process of a time series may lead traditional linear unit root tests to exhibit important power losses (see Pippenger and Goering, 1993). The test of Kapetanios *et al.* (2003) allows us to characterise inflation as a two regime process, with the regime change being smooth. The use of this test allows the inflation series to revert faster to its fundamental equilibrium when the deviation from equilibrium is larger. Theoretically, as pointed out by Nobay *et al.* (2007), adjustment towards an implicit or explicit inflation target can be nonlinear. As a matter of fact, the opportunistic approach to disinflation proposed by Orphanides and Wilcox (2002) and Aksoy *et al.* (2006) implies that a central bank controls inflation aggressively when inflation is far from its target, while focusing on stabilising output when inflation is close to its target.

[11] Interestingly, Nobay *et al.* (2007) model U.S. inflation as a nonlinear ESTAR process. Along similar lines, Byers and Peel (2000) provide evidence that a nonlinear ESTAR process best describes the required characteristics for analysing inflation in high inflation economies.

directly from the impulse-response function.[12] For that purpose, we employ the procedure proposed by Gospodinov (2004). In addition, and with the aim of providing information on the degree of precision in the point estimates, we also present confidence interval estimates for both the half-life and persistence parameter. The 95% confidence intervals for the persistence parameter are computed through the procedure proposed by Hansen (1999), while those for the half-life through the procedure explained in Gospodinov (2004). In both cases, the use of the grid-bootstrap method allows us to overcome the bias associated with the confidence intervals derived from the application of Ordinary Least Squares (OLS) or those obtained via conventional bootstrap for the case of near unit-root processes (Basawa *et al.*, 1991).

In sum, by using this whole set of persistence measures we should be able to provide a clear conclusion about the degree of persistence of the inflation rates for the Spanish economy, taking into account several degrees of disaggregation in the data. We also think as pivotal the comparison between consumer prices and producer prices, the latter covering the manufacturing and energy sectors. The rest of the article is structured as follows: Section 2 presents the methodology and results obtained from the linear and nonlinear unit root tests employed. Section 3 deals with the persistence analysis through the computation of median unbiased point estimates of the persistence parameter and half-life of a unit shock, along with their confidence intervals. Finally, Section 4 summarises the main findings and concludes.

2. Unit Root Analysis of Inflation Rates

2.1 Methodology

In this section we briefly describe the methodology behind the class of modified tests (M-tests), originally proposed by Stock (1999), and extended by Perron and Ng (1996) and Ng and Perron (2001). The latter apply local-to-unity Generalised Least Squares (GLS) detrending instead of OLS when estimating the deterministic components of an ADF regression (see Elliot *et al.*, 1996, ERS hereafter) so that important gains in statistical power can be achieved. Assuming a specification like:

$$\Delta y_t = d_t + \beta_0 y_{t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_{tk} \tag{1}$$

where $d_t = \sum_{i=0}^{p} \psi_i t^i$ with *p=0,1* for the case of a constant and a linear trend. The ADF test originally proposed by Dickey and Fuller (1979) is the t-statistic for $\beta_0 = 0$ in (1). To gain power,

[12] The half-life of a shock measures the number of years for a unit impulse to dissipate by one half.

the original series is detrended using local-to-unity GLS detrending rendering $\tilde{y}_t \equiv y_t - \hat{\psi}' z_t$, where z_t is a set of deterministic components and $\hat{\psi}$ minimizes $S(\bar{\alpha}, \psi) = (y^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})'(y^{\bar{\alpha}} - \psi' z^{\bar{\alpha}})$, for $(y_0^{\bar{\alpha}}, y_t^{\bar{\alpha}}) \equiv (y_0, (1-\bar{\alpha}L)y_t)$, *t=1,…,T*, for some chosen $\bar{\alpha} = 1 + \bar{c}/T$. The value of $\bar{c}$ is selected in a way that the asymptotic local power function of the test is tangent to the power envelope at the 50% power level. For *p=0,* $\bar{c} = -7.0$ and for *p=1,* $\bar{c} = -13.5$.[13]

Ng and Perron (2001) recommend using the ADF^{GLS} statistic, originally proposed by ERS, which is the t-statistic for testing $\beta_0 = 0$ in an ADF specification where the series is detrended with GLS prior to the estimation by OLS such that:

$$\Delta\tilde{y}_t = \beta_0 \tilde{y}_{t-1} + \sum_{j=1}^{k} \beta_j \Delta\tilde{y}_{t-j} + e_{tk} \tag{2}$$

Despite the substantially higher power of the ADF^{GLS} test, Ng and Perron (2001) find the test to exhibit size distortions in the presence of a large negative moving average root. To circumvent that shortcoming, they propose the M-class of unit root tests exploiting GLS detrending (M^{GLS} tests) for which size distortions are less important. Ng and Perron (2001) develop the MZ_α^{GLS} and MZ_t^{GLS} tests that are modified versions of the Phillips-Perron (1988) Z_α and Z_t tests and the MSB^{GLS} test which is a modified version of the test proposed by Sargan and Bhargava (1983) and is computed as $MSB^{GLS} = MZ_t^{GLS} / MZ_\alpha^{GLS}$. In addition, we also employ the feasible point optimal test (P_T^{GLS}) and the modified feasible point optimal test (MP_T^{GLS}).[14] All these statistics test the null hypothesis of a unit root versus the alternative of stationarity.

The computation of these tests requires determining the appropriate lag length (*k*) for the ADF regression using GLS-detrended data given by (2). Ng and Perron (2001) propose the modified Akaike information criterion (MAIC), which is given by:

$$MAIC(k) = \ln(\hat{\sigma}_{k_m}^2) + \frac{2(\tau_T(k) + k)}{T - k_{\max}} \tag{3}$$

where $\tau_T(k) = (\hat{\sigma}_{k_m}^2)^{-1} \hat{\beta}_0^2 \sum_{t=k_{maz}+1}^{T} \tilde{y}_{t-1}^2$, $\hat{\sigma}_{k_m}^2 = (T - k_{\max})^{-1} \sum_{t=k_{\max}+1}^{T} \hat{e}_{tk}^2$ and $k_{\max}$ is the maximum value of *k* considered ($k_{\max} = 20$ in our analysis). The MAIC aims at selecting a relatively long lag length in the presence of a large negative moving average root (thus preventing size distortions) and a short lag length when that root is not present (thus avoiding unnecessary loss of power). However,

[13] In our case, given the absence of a linear trend in the inflation rate series, we will only include a constant in the specification.
[14] The former was proposed by ERS, while the latter was developed by Ng and Perron (2001).

as noted by Perron and Qu (2007), a drawback of these tests is that the power of the tests falls significantly for non-local alternatives. To cope with this problem, we follow Perron and Qu (2007) by selecting the optimal AR order k using the MAIC constructed with OLS detrended data (rather than GLS data). In our application, we take a maximum lag truncation equal to $k_{\max} = \text{int}(12(T/100)^{1/4})$, where int() denotes the integer part.[15]

Regarding the nonlinear ESTAR unit root test of Kapetanios *et al.* (2003), after applying a first-order Taylor series approximation to a ESTAR specification, they obtain the following specification:

$$\Delta y_t = \delta y_{t-1}^3 + \sum_{j=1}^{p} \rho_j \Delta y_{t-j} + \upsilon_t \tag{4}$$

where the p augmentations to correct for serially correlated errors are calculated through the step-down procedure based on the significance of the coefficients with a maximum p equal to 20. The null hypothesis of nonstationarity $(H_0 : \delta = 0)$ versus the alternative of nonlinear ESTAR stationarity $(H_1 : \delta < 0)$ is tested with the respective t-statistic denoted as t_{NL}.

2.2 Results

In this section we present the results from the application of the unit root tests with good size and power of Ng and Perron (2001) as well as the nonlinear unit root test of Kapetanios *et al.* (2003). Prior to that, we should point out that our datasets consist of inter-annual inflation rates presented at a monthly frequency. We employ this particular definition of inflation because this is the specific inflation rate targeted by the ECB, to be slightly below 2%. Hence, the use of this series allows us to shed some light on the duration of the effect of adverse shocks that make the Spanish economy be far from the ECB inflation target.[16] Given the important data requirements of our analysis, we have employed homogeneous series from the National Institute of Statistics (INE) for both CPI and PPI, of the greatest possible length for the different disaggregation levels. As a result, the sample period generally differs for the different series. We shall also point out that we have conducted the analysis with data spanning until April 2008. This tries to prevent the recent cyclical fluctuations caused by the economic crisis –that are irregular from a historical perspective– from affecting the structural characteristics of Spanish inflation that our analysis tries to capture.

Table 1 presents the results for the inflation rates based on the CPI for the 17 Spanish regions and the aggregate over the period ranging between January 1979 and April 2008. Table 2 presents the

[15] All computations were conducted with the GAUSS software.

[16] Interestingly, the Spanish economy has been well above the 2% target until the current crisis. During the last months the Spanish economy is also exhibiting inflation rates above 2%.

results for the inflation rates based on CPI data for the 50 provinces plus the two autonomous municipalities of Ceuta and Melilla covering the period between January 1971 and April 2008. Table 3 shows the results for the CPI-based inflation rates for the main groups of goods and services over the period from January 1978 to December 2000. Finally, Table 4 reports the results for the inflation rates based on PPI data for 24 sectors[17] and the aggregate over the period between January 1976 and December 2002.

As shown in Tables 1, 2 and 3, all inflation rate series based on CPI (aggregate, region-level, province-level and good or service group-based) clearly exhibit a unit root, since we are unable to reject the nonstationarity null even at the 10% significance level. These results point to the existence of a very high degree of persistence in the inflation rate series based on CPI data. The unit root outcome indicates that the effect of a shock on CPI inflation will last for a long time, with the negative effect that can bring about on the economy. Let us for instance think about the external factors (political, institutional, natural, etc.) causing sharp peaks in the prices of oil, alternative energy sources and food, which can lead to highly persistent effects on consumer prices. When we allow for a nonlinear but globally stationary exponential smooth transition autoregressive process, the Kapetanios *et al.* (2003) test is able to reject the null of a unit root for five regions (Aragón, Asturias, Canarias, Cantabria and Madrid), one province (Ávila) and two groups of goods and services (Footwear and clothing and Cookware and home services).

[Insert Tables 1, 2 and 3 about here]

With respect to the results for the inflation rate series based on PPI data, Table 4 shows that the aggregate figure exhibits a unit root, since we fail to reject the nonstationarity null with all the statistics, even at the 10% significance level. However, at the sectoral level, we are able to reject the unit root null, at least at the 10% level, for the following sectors: 13 (Petroleum refineries), 21 (Iron and steel), 22 (Basic metals and fabricated metal products), 25 (Chemicals and chemicals products), 41 (Food products, beverages and tobacco), 44 (Leather), 46 (Wood, products of wood and cork), 47 (Pulp, paper, paper products, printing and publishing) and 49 (Other manufacturing industries). In addition, for Sector 31 (Metal products –except machinery and transportation equipment), there is also some evidence of stationarity, as we are able to reject the null hypothesis of a unit root with the ADF^{GLS} statistic at the 1% significance level. This indicates that for these sectors (10 out of 24), the effect of a shock on the inflation rate will tend to dissipate as time passes, thus leading the series to revert to the existing mean value prior to the occurrence of the shock. When we allow for a globally stationary ESTAR process under the alternative, the t_{NL} statistic is able to reject the unit

[17] In this respect, we must highlight that visual inspection shows remarkably different patterns in the following sectors: 13 (Petroleum refineries), 21 (Iron and steel) and 49 (Other manufacturing industries).

root null for eight sectors for which the linear unit root tests could reject the null, plus five additional sectors: 11 (Coke, petroleum and coal products), 24 (Nonmetal products), 36 (Motor vehicle and spare motor parts), 39 (Optical and precision equipment) and 45 (Textile products and footwear).

The comparison of our findings with CPI data to those obtained with PPI data for Spain highlights the existence of a higher degree of persistence for the former. Hence, the flexibility of Spanish consumer prices –that exhibit *hysteresis* or extreme persistence– is lower than that of producer prices.[18] This may underline the existence of problems in the distribution system, which serves as a complex bridge between producers and consumers, in line with the arguments put forward by Blanchard (1987) and Frey and Manera (2007), among others.[19]

[Insert Table 4 about here]

3. Persistence Analysis of Inflation Rates

3.1 Methodology

Once analysed the possible presence of a unit root in the inflation rate through a large battery of unit root tests, we now shift the focus to the application of more informative measures of persistence like the median unbiased estimates of the persistence parameter and the half-life of a unit shock, along with their associated confidence intervals. Unlike unit root tests that provide an incomplete picture of the persistence of a macroeconomic series as they focus on testing whether the sum of the AR coefficients –i.e. the persistence parameter– is equal or less than unity, these alternative persistence indicators will provide us with the exact measure of persistence of the inflation rate series.

The median-unbiased method, originally proposed by Andrews (1993) for AR(1) processes and later extended by Andrews and Chen (1994) for AR processes of order greater than one, can account for the downward bias associated with least squares estimates of the persistence parameter. This bias is caused by the skewness to the left in the distribution of the estimators of the persistence parameter in autoregressions, thus making the median exceed the mean. So under these circumstances the median constitutes a better measure of central tendency than the mean. The provision of a more accurate estimate of the persistence of shocks to a series through median-unbiased half-lives along with confidence intervals is expected to be crucial in addressing the low

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[18] Álvarez (2008) and Álvarez *et al.* (2010), among others, using different price indicators, also conclude the existence of greater flexibility in producer prices than in consumer prices for the Spanish economy.

[19] See also Álvarez and Burriel (2010) for Spain. Due to the features of the Spanish economy, it could be interesting to pay attention to the recent methodology on the price transmission mechanism applied to some agricultural commodities by Fernández-Amador *et al.* (2010). Another line of research could be, for instance, the use of oil price pass-through models.

power of univariate unit root tests. This is because they can be very informative about whether the failure to reject the unit root null is caused by low power or is due to the existence of a high degree of uncertainty about the true value of the persistence parameter.

Traditionally, half-lives are calculated with the formula *ln(0,5)/ln(*α*)* –rather than through impulse-response functions–, where the parameter α is the sum of the coefficients of the underlying AR process. Even though this is appropriate for an AR(1) process that decays monotonically, it is inappropriate for higher order AR processes for which shocks do not generally decay at a constant rate. To cope with this shortcoming, Cheung and Lai (2000) recommend computing estimates of the half-life of a shock –leading to deviations from the equilibrium inflation rate– directly from the impulse-response function. For that purpose, we follow the method proposed by Gospodinov (2004) for constructing confidence intervals for impulse-response functions and half-lives of near non-stationary processes. More specifically, Gospodinov (2004) proceeds with an ADF representation of an AR(p) process, which takes the form: [20]

$$\pi_t = \alpha\pi_{t-1} + \Psi(L)\Delta\pi_{t-1} + \varepsilon_t \tag{5}$$

where π_t denotes the inflation rate at time *t*, α is the sum of the AR coefficients –that according to Andrews and Chen (1994) represents a measure of persistence more informative than the larger root of an AR(p) process[21]– and $\Psi(L) = \sum_{j=1}^{p-1}\psi_j L^j$, where *L* represents the lag operator. If the process is a near unit root, it is useful to reparameterise it as local-to-unity ($\alpha_T \approx 1 + c/T$) for a fixed constant $c < 0$ (near unit root process) versus $c = 0$ (unit root process). This removes the discontinuity of the distribution function in the neighbourhood of one. In sum, this method is based on the inversion of the acceptance region of the likelihood ratio statistic under a sequence of null hypotheses that restrict the values of the half-life or impulse-response to a shock. By reparameterising the leading time of the impulse-response function as a function of the sample size, the order of the polynomial constraint by the null hypothesis is not constant but rises linearly with the sample size. This allows for the identification and consistent estimation of the persistence parameter, thereby facilitating the assessment of the asymptotic distribution of the maximum likelihood statistic. In addition, this statistic can be employed to construct median unbiased point estimates of the half-life and

[20] The parameter *p* refers to the number of lags necessary to correct for autocorrelation, and is determined through the MAIC procedure proposed by Ng and Perron (2001), with the small-sample bias correction of Perron and Qu (2007).

[21] This is due to the fact that two models AR(*p*) characterised by the same larger root can exhibit totally different persistence properties. The persistence parameter α can be considered as the cumulative impulse-response (CIR); that is, the sum of the impulse-response through the whole time horizon over which there is an effect, such that CIR=1/1-α.

associated confidence intervals, measuring it as the required time for a unit shock to dissipate by half.[22]

With the aim of presenting additional persistence measures, we also calculate the 95% confidence intervals for the persistence parameter. However, it is important to point out some problems associated with the construction of these confidence intervals for the parameter α, since the asymptotic distribution of the OLS estimator (as well as its convergence rate) is different in the case of a stationary process from the one of a unit root. In sum, if α<1, the confidence intervals can be calculated through asymptotic methods based on the standard normal distribution. Nevertheless, this method is not valid for finite samples, particularly when α approaches unity. Indeed, the formalisation of a near unit root process following the local-to-unity framework, where $\alpha_T \approx 1 + c/T$ with c as a constant when $T \to \infty$, leads the conventional t-statistic to be characterised by a non-standard distribution. This is due to the fact that the parameter c, which in turn is a function of α, makes the conventional confidence intervals suffer from severe Type I errors. In contrast, the method proposed by Hansen (1999) generates confidence intervals for near unit-root processes with first-order asymptotic coverage rates for the sum of the AR coefficients. As a result, this method constitutes a good alternative for obtaining correct estimates of the confidence intervals for the persistence parameter. Unlike conventional bootstrap procedures, the grid-bootstrap method of Hansen (1999) does not only compute the empirical percentiles forming the t-statistic distribution for the estimates of the persistence parameter ($\hat{\alpha}$), but also for a complete grid of α values. By conducting Monte Carlo simulations, Hansen (1999) finds his method to generate confidence intervals with good coverage rates even for finite samples.

3.2 Results

In Section 2 we presented the results from unit root testing, which in general supported the existence of a high degree of persistence, since we failed to reject the nonstationarity hypothesis for most inflation series based on CPI. The same occurred for the PPI-based aggregate inflation rate as well as for most of the individual sectors. As a complement to that analysis, we now proceed to present the results from the median-unbiased estimation of the persistence parameter and half-lives,

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[22] This procedure improves on conventional bootstrap methods –see, for instance, Inoue and Kilian (2002)– in that the confidence intervals are correctly computed for the sum of the AR coefficients rather than for individual slope coefficients of the AR(p) model. This method also provides correct coverage rates of the confidence intervals in large samples and good coverage rates in small samples comparing to conventional bootstrap methods when large AR roots are present. This is achieved by making the persistence parameter a function of the sample size. Finally, the inversion of the likelihood ratio statistic provides tighter confidence intervals by shifting them away from the upper boundary more often than other methods. In addition, Gospodinov (2004) shows that the confidence intervals computed through conventional bootstrap for processes AR(p) with p>1 exhibit very low coverage rates in finite samples for near unit-root processes.

along with 95% confidence intervals for both point estimates for all inflation rate series. More specifically, Table 5 shows the results for the CPI-based inflation rates for the Spanish regions and the aggregate. Table 6 presents the results for the CPI-based inflation series at the provincial level of disaggregation. Table 7 focuses on the results for inflation series based on consumer prices for the different groups of goods and services. Finally, Table 8 reports the findings from the analysis of aggregate and sectoral inflation series based on PPI data.

First of all, it is remarkable the fact that the persistence parameter (sum of the AR coefficients) estimated via OLS –which is characterised by a downward bias– is always lower than the median-unbiased estimate obtained by applying the procedure of Gospodinov (2004). In the case of the aggregate CPI, the median-unbiased point estimate of the persistence parameter is pretty close to unity, which would entail the existence of a unit root in the series. In fact, the median-unbiased upper bound of the 95% confidence interval is equal to one, corroborating the fact that aggregate inflation is likely to contain a unit root. The half-life of a unit shock on inflation based on aggregate CPI equals nine years, which implies that nine years are required for a unit shock on the series to dissipate by half. The yearly speed of adjustment for aggregate CPI-based inflation equals 7.4%, entailing that the effect of the shock vanishes very slowly. As with most disaggregate series, the upper bound associated with the 95% confidence interval for the half-life equals infinity.[23] This is compatible with the presence of a unit root in the series, which implies that the effect of the shock does not vanish no matter the time horizon considered.

Regarding regional CPI-based inflation, we find clear-cut evidence of a unit root, given that the persistence parameter is equal to unity for all the series. Half-life estimation appears to uncover a lower degree of persistence in regional inflation than in the aggregate, since the region with the highest persistence (País Vasco) presents a half-life below 5 years, and for most of the regions is lower than 4 years. The regions with the lowest persistence are Baleares, Cantabria and Navarra, with half-lives below 2.5 years; closely followed by Aragón, Asturias, Castilla-León, Castilla-La Mancha, Madrid and La Rioja, with half-lives near 3 years. In addition, Andalucía, Canarias, Cataluña, Comunidad Valenciana, Extremadura, Galicia and Murcia exhibit half-lives between 3.5 and 4.5 years. The median half-life for the 17 regions is equal to 3.24 years, with a convergence rate of 19.3% per year. Likewise, the average half-life point estimate equals 3.65 years, which implies a yearly convergence speed of 17.3%. As with aggregate inflation, the upper bound of the 95% confidence interval for the half-life of a unit shock on inflation equals infinity for all the regions; which is again compatible with the existence of a unit root in regional CPI inflation.

[23] This means that it surpasses the maximum value considered, which in this case is 25 years.

[Insert Table 5 about here]

As far as the results for provincial inflation are concerned, Table 6 shows that the median unbiased estimate of the persistence parameter is always above 0.99 and the upper bound of the associated confidence interval is consistent with the presence of a unit root in the series. Regarding the half-life estimates resulting from a unit shock on province-level inflation rates, we find only 8 provinces (Huelva, Lugo, Pontevedra, Santa Cruz de Tenerife, Teruel, Toledo, Ceuta and Melilla) with a half-life below 5 years, while the rest of the provinces exhibit a higher degree of persistence according to this measure. The province with the highest persistence is Asturias, with an estimated half-life of 11.2 years. The median of the half-life point estimates for all the provinces equals 6.24 years, while the average counterpart is equal to 6.4 years. The annual convergence rates towards the inflation levels prior to the occurrence of the shock are 10.5% for the median province and 10.3% for the average province. As with aggregate and regional CPI inflation, the upper bound of the 95% confidence interval for the half-life estimates is infinity for all the provinces.

[Insert Table 6 about here]

With regard to the results for the CPI-based inflation series for the different groups of goods and services, Table 7 shows that the median-unbiased estimate of the persistence parameter is equal to unity for all sectors, which supports the presence of a unit root in these series. This high degree of persistence is also reflected in the lower bound of the 95% confidence interval, which is always greater than 0.99. In addition, the half-life point estimates resulting from a unit shock on these series indicate that the groups with the lowest persistence are: Group 1 (Food, beverages and tobacco), Group 6 (Transport and communication) and Group 3 (Housing), with half-lives equal to 1.7, 2.6 and 2.97 years, respectively. These groups are followed by Group 7 (Leisure, education and culture), Group 8 (Other goods and services) and Group 5 (Medicine and healthcare), with half-lives of 4.3, 4.7 and 5.9 years, respectively. Moreover, the groups with the highest persistence are Group 2 (Footwear and clothing) and Group 4 (Cookware and home services) with half-lives of 7.5 and 8.2 years, respectively. The median of the half-life point estimates for the eight groups is 4.52 years, with an associated speed of adjustment of 14.2% per year; while the average counterparts are 4.73 years and 13.6%, respectively. As with previous findings, the upper bound of the 95% confidence interval associated with the half-life for these eight groups of goods and services equals infinity. This is again indicative of the existence of a unit root, thereby entailing that the effect of the shock does not vanish over time.[24]

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[24] Our analysis of persistence for CPI-based inflation series uncovers very high persistence from all perspectives. We must point out that when we apply a simpler methodology involving the use of the standard deviation of the inflation rate as an inverse indicator to the degree of persistence for five main groups of goods and services, we can observe that

[Insert Table 7 about here]

Finally, Table 8 presents the results for the PPI-based aggregate and sectoral inflation rate series. As far as the aggregate figure is concerned, we obtain a persistence parameter estimate very close to unity (0.994) and an upper bound of the 95% confidence interval equal to one, which is again consistent with the presence of a unit root in the series. The average half-life in this case is equal to 7.1 years, which is equivalent to an adjustment rate as a response to a unit shock on inflation of 9.3% per year. With regard to the sector-level inflation rate series, we find several sectors with a degree of persistence lower than that of a unit root, since even the upper bound of the 95% confidence interval of the median unbiased point estimate of the persistence parameter is lower than unity. This appears to be the case for the following sectors: 13 (Petroleum refineries), 21 (Iron and steel), 22 (Basic metals and fabricated metal products), 31 (Metal products –except machinery and transportation equipment), 44 (Leather), 46 (Wood, products of wood and cork) and 49 (Other manufacturing industries). It is remarkable that the results from unit root testing presented in Section 2 provided evidence of stationarity in the inflation rate series for these seven sectors, as we could reject the unit root null hypothesis. Hence, our detailed analysis of persistence largely confirms the evidence found with unit root tests with good size and power properties.[25] The median (average) of the median-unbiased persistence parameter equals 0.981 (0.976).

With respect to the half-life point estimates, the sectors with the lowest persistence, with a half-life lower than two years are: 11 (Coke, petroleum and coal products), 13 (Petroleum refineries), 15 (Electricity, gas and water supply), 21 (Iron and steel), 22 (Basic metals and fabricated metal products), 36 (Motor vehicle and spare motor parts) and 49 (Other manufacturing industries). These sectors are followed by those sectors with a half-life between two and five years: 23 (Nonferrous and nonfuel metals, peat), 24 (Nonmetal products), 33 (Office and computing –including installation), 34 (Electric machinery), 35 (Electronic equipment –except computers), 37 (Other transport equipment), 39 (Optical and precision equipment), 41 (Food products, beverages and tobacco), 44 (Leather), 45 (Textile products and footwear), 46 (Wood, products of wood and cork) and 47 (Pulp, paper, paper products, printing and publishing). Finally, the sectors with the highest persistence according to this measure, with half-lives greater than 5 years but lower than 8 years are: 25 (Chemicals and chemicals products), 31 (Metal products –except machinery and transportation equipment), 32 (Machinery and mechanical equipment), 43 (Textile) and 48 (Rubber

persistence follows the following decreasing order: services, nonfuel manufacturing goods, food products, unprocessed food and energy. However, when we further disaggregate by groups or sectors a higher degree of heterogeneity emerges.

[25] It is important to note that some previous work has concluded that heterogeneity in inflation patterns is more prevalent at the sectoral level than geographically, feature that appears clearly reflected in our results. We refer the reader to Álvarez (2008) and Caraballo and Usabiaga (2009a).

and plastic products). The median of the half-life point estimates for the 24 sectors equals 2.96 years, which entails a yearly adjustment speed of 20.9%; while the mean counterparts are 3.40 years and 18.7%, respectively. Finally, it is necessary to stress the fact that in all sectors, with the exception of sector 21 (Iron and steel) that is clearly stationary, the upper bound of the confidence interval associated with the half-life estimates is equal to infinity, which is again indicative of the high degree of persistence of PPI-based sectoral inflation.[26]

[Insert Table 8 about here]

If we compare the average half-life (4.73 years) and the median half-life (4.52 years) for the different CPI-based groups of goods and services, with the average half-life (3.40 years) and median half-life (2.96 years) for the PPI-based sectors, in addition to the comparison of the average half-life for aggregate CPI inflation (8.98 years) with respect to that associated with aggregate PPI inflation (7.09 years), we can observe that there are clear indications of higher flexibility in producer prices versus consumer prices for the Spanish economy.

4. Conclusions

This article has assessed the degree of persistence of different inflation rates for the Spanish economy computed using the CPI for the aggregate, regions, provinces and eight groups of goods and services, in addition to the PPI for the aggregate and 24 industrial sectors. For that purpose, we have employed: 1) the unit root tests with good size and power of Ng and Perron (2001) with the small-sample bias correction developed by Perron and Qu (2007); 2) the nonlinear ESTAR unit root test proposed by Kapetanios *et al.* (2003); 3) median-unbiased estimations of the persistence parameter and the respective confidence intervals through the grid-bootstrap method proposed by Hansen (1999); and 4) median-unbiased estimations of the half-life of a shock in addition to the associated confidence intervals through the method based on impulse-response functions proposed by Gospodinov (2004).

Our results from the unit root analysis clearly indicate that most inflation rate series based on CPI (aggregate, regional, provincial and group-based) contain a unit root, since we failed to reject the nonstationarity null even at the 10% significance level in most of the cases. These results thus support the presence of a very high degree of persistence in the CPI-based inflation series. Regarding the results for PPI-based inflation, there is evidence that the aggregate figure contains a unit root, while at the sectoral level there is clear-cut evidence of a unit root in only 11 of the 24 sectors analysed once we allow for a nonlinear but globally stationary SETAR process under the

[26] We must point out that we do not observe a remarkable difference in results between energy and manufacturing sectors, despite the fact that theoretical intuition would point to lower persistence in the former.

alternative. This highlights the existence of a different behaviour for consumer and producer prices, finding more persistence in the former.

Since the information obtained from unit root testing may be limited, as they focus on testing whether the persistence parameter is equal or less than unity, we have complemented that analysis with median-unbiased point estimates of the persistence parameter and half-lives. In order to provide information on the degree of precision in the point estimates, we have also computed 95% confidence intervals estimates. Overall, this set of measures has provided evidence of a very high degree of persistence, consistent with the presence of a unit root in: 1) the aggregate inflation rates for both CPI and PPI, 2) regional, provincial and group-based CPI inflation rates, and 3) PPI inflation rate series for most sectors –with the exception of seven sectors that exhibit a lower degree of persistence than that associated with a unit root. These results further confirm those derived from the application of unit root tests with good size and power, overall concluding that Spanish inflation rate series exhibit a very high degree of persistence, particularly from the consumer perspective. We have also noticed greater geographical than sectoral homogeneity. Future work could take the evidence of relatively high persistence in Spanish inflation and try to determine common causes of upward pressures affecting both producer and consumer inflation. One possibility would be the use of common factor models along the lines of Bai and Ng (2004a, 2004b) in an attempt to identify common forces that can explain the widespread tendency for inflation to exhibit high persistence across both the geographical and sectoral dimensions.

Given the existence of a large bulk of literature on the determinants of inflation persistence pointing to severe problems for the Spanish economy such as, among others, wide-spread indexation, prevalence of backward-looking expectations, low competition in the markets of goods and services, and important rigidities in the labour market mainly linked to the insider-outsider framework, our results should not be at all surprising. We should not disregard the fact that high inflation persistence makes external shocks (like those related to oil supply[27]) have a long-lasting effect on inflation. Hence, in order to prevent inflation from moving farther away from the targeted value, policymakers should take appropriate policy measures to make prices more flexible, particularly consumer prices. The greater flexibility associated with producer prices compared to consumer prices also highlights the need to pay more attention to the particularities of the distribution system in Spain. Among other policy measures, Spanish authorities should strengthen the defence of free competition in goods and services markets and conduct structural reforms in the

[27] The important effects of the oil shocks on Spanish inflation are shown in Caraballo and Usabiaga (2009c).

labour market with the aim of weakening the aforementioned sources of inflation persistence, as well as reducing the structural inflation differential of Spain in relation to the core euro-zone.

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References

Aksoy, Y., Orphanides, A., Small, D., Wieland, V. and Wilcox, D. (2006) A Quantitative Exploration of the Opportunistic Approach to Disinflation, *Journal of Monetary Economics*, **53** (8), 1877-1893.

Altissimo, F., Ehrmann, M. and Smets, F. (2006) Inflation Persistence and Price-Setting Behaviour in the Euro Area. A Summary of the Inflation Persistence Network Evidence", European Central Bank, Occasional Paper nº 46, June.

Álvarez, L.J. (2008) What Do Micro Price Data Tell Us on the Validity of the New Keynesian Phillips Curve, *Economics: The Open-Access, Open-Assessment E-Journal*, **2** (19).

Álvarez, L.J. and Burriel, P. (2010) Is Calvo Price Setting Model Consistent with Micro Price Data?, Banco de España, Documento de Trabajo nº 1010.

Álvarez, L.J., Burriel, P. and Hernando, I. (2010) Price Setting Behaviour in Spain: Evidence from Micro PPI Data, *Managerial and Decision Economics,* **31** (2-3), 105-121.

Álvarez, L.J., Dhyne, E., Hoeberichts, M.M., Kwapil, C., Le Bihan, H., Lünnemann, P., Martins, F., Sabbatini, R., Stahl, H., Vermeulen, P. and Vilmunen, J. (2006) Sticky Prices in the Euro Area:

A Summary of New Micro Evidence, *Journal of the European Economic Association*, **4** (2-3), 575-584.

Álvarez, L.J. and Hernando, I. (2006) Price Setting Behaviour in Spain. Evidence from Consumer Price Micro-Data, *Economic Modelling,* **23** (4), 699-716.

Andrews, D.W.K. (1993). Exactly Median-Unbiased Estimation of First Order Autoregressive/Unit Root Models, *Econometrica*, **61** (1), 139-165.

Andrews, D.W.K. and Chen, H.Y. (1994) Approximately Median-Unbiased Estimation of Autoregressive Models, *Journal of Business and Economic Statistics*, **12** (2), 187-204.

Angeloni, I., Aucremanne, L., Ehrmann, M., Galí, J., Levin, A. and Smets, F. (2006) New Evidence on Inflation Persistence and Price Stickiness in the Euro Area: Implications for Macro Modeling, *Journal of the European Economic Association*, **4** (2-3), 562-574.

Bai, J., and Ng, S. (2004a) A PANIC Attack on Unit Roots and Cointegration, *Econometrica*, **72** (4), 1127–77.

Bai, J., and Ng, S. (2004b) A New Look at Panel Testing of Stationarity and the PPP Hypothesis, in Andrews, D.W and Stock, J. (Eds.) *Identification and Inference in Econometric Models: Essays in Honor of Thomas J. Rothemberg*, Cambridge University Press, Cambridge, 426-450.

Basawa, I.V., Mallik, A.K., McCormick, W.P., Reeves, J.H. and Taylor, R.L. (1991) Bootstrapping Unstable First-order Autoregressive Processes, *Annals of Statistics*, **19** (2), 998-1101.

Blanchard, O.J. (1987) Aggregate and Individual Price Adjustment, *Brookings Papers on Economic Activity*, **1**, 57-109.

Blanchard, O.J. and Summers, L. (1986) Hysteresis and the European Unemployment Problem, *NBER Macroeconomics Annual 1986*, 15-78.

Byers, J.D. and Peel, D.A. (2000) Non-linear Dynamics of Inflation in High Inflation Economies, *Manchester School*, **68** (3), 23-37.

Caraballo, M.A. and Usabiaga, C. (2009a) *Análisis Desagregado de la Inflación Española y Andaluza*, Instituto de Estadística de Andalucía, Sevilla.

Caraballo, M.A. and Usabiaga, C. (2009b) Testing Nominal Rigidities in an Integrated Economy: An Application to Spain, in Marques, H., Souzakis, E. and Cerqueira, P. (Eds.) *Integration and Globalization. Challenges for Developed and Developing Countries*, Edward Elgar, Chentelham, 43-62.

Caraballo, M.A. and Usabiaga, C. (2009c) The Relevance of Supply Shocks for Inflation: The Spanish Case, *Applied Economics*, **41** (6), 753-764.

Cheung Y.W. and Lai K.S. (2000) On the Purchasing Power Parity Puzzle, *Journal of International Economics,* **52** (2), 321-330.

De Jong, J.Y., Nankervis, J.C., Savin, N.E. and Whiteman, C.H. (1992) The Power Problems of Unit Root Tests in Time Series with Autoregressive Errors, *Journal of Econometrics*, **53** (1-3), 323-343.

Dhyne, E., Álvarez, L.J., Le Bihan, H., Veronese, G., Dias, D., Hoffmann, J., Jonker, N., Lünnemann, P., Rumler, F. and Vilmunen, J. (2006) Price Setting in the Euro Area and the United States: Some Facts from Individual Consumer Price Data, *Journal of Economic Perspectives*, **20** (2), 171-192.

Dickey, D.A. and Fuller, W.A. (1979) Distribution of the Estimators for Autoregressive Time Series with a Unit Root, *Journal of the American Statistical Association*, **74** (366), 427-431.

Elliot, G., Rothenberg, T. and Stock, J.H. (1996) Efficient Tests of an Autoregressive Unit Root, *Econometrica*, **64** (4), 813-839.

Estrada, A. and López-Salido, J.D. (2004) Understanding Spanish Dual Inflation, *Investigaciones Económicas*, **28** (1), 123-140.

European Central Bank (2003) *Inflation Differentials in the Euro Area: Potential Causes and Policy Implications*, Monetary Policy Committee of the European System of Central Banks, Frankfurt.

Fabiani, S., Druant, M., Hernando, I., Kwapil, C., Landau, B., Loupias, C., Martins, F., Mathä, T., Sabbatini, R., Stahl, H. and Stokman, A. (2006) What Firms' Surveys Tell About Price Setting Behavior in the Euro Area, *International Journal of Central Banking*, **2** (3), 3-47.

Fernández-Amador, O., Baumgartner, J. and Crespo-Cuaresma, J. (2010) Milking the Prices: The Role of Asymmetries in the Price Transmission Mechanism for Milk Products in Austria, WIFO Working Papers, nº 378.

Frey, G. and Manera, M. (2007) Econometric Models of Asymmetric Price Transmission, *Journal of Economic Surveys*, **21** (2), 349-415.

Galí, J. and Gertler, M. (1999) Inflation Dynamics: A Structural Econometric Approach, *Journal of Monetary Economics*, **44** (2), 195-222.

Galí, J. and López-Salido, J.D. (2001) Una Nueva Curva de Phillips para España, *Moneda y Crédito*, **212**, 265-310.

Gaspar, V., Levin, A., Martins, F. and Smets, F. (2007) Evidence from Surveys of Price-Setting Managers: Policy Lessons and Directions for Ongoing Research, CEPR, Discussion Paper nº 6227, March.

Gordon, R.J. (2011) The History of the Phillips Curve: Consensus and Bifurcation, *Economica*, **78** (309), 10-50.

Gospodinov, N. (2004) Asymptotic Confidence Intervals for Impulse Responses of Near-Integrated Processes, *Econometrics Journal,* **7** (2), 505-527.

Hansen, B.E. (1999) The Grid Bootstrap and the Autoregressive Model, *Review of Economics and Statistics*, **81** (4), 594–607.

Inoue, A. and Kilian, L. (2002) Bootstrapping Autoregressive Processes with Possible Unit Roots, *Econometrica,* **70** (1), 377-391.

Kapetanios, G., Shin, Y. and Snell, A. (2003) Testing for a Unit Root in the Nonlinear STAR Framework, *Journal of Econometrics*, **112** (2), 359-379.

Mankiw, N.G. and Reis, R. (2002) Sticky Information Versus Sticky Prices: A Proposal To Replace The New Keynesian Phillips Curve, *The Quarterly Journal of Economics*, **117** (4), 1295-1328.

Ng, S. and Perron, P. (2001) Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power, *Econometrica*, **69** (6), 1519-1554.

Nobay, B., Paya, I. and Peel, D.A. (2007) Inflation Dynamics in the U.S. – A Nonlinear Perspective, Financial Markets Group, Discussion Paper nº 601.

Orphanides, A. and Wilcox, D. (2002) The Opportunistic Approach to Disinflation, *International Finance*, **5** (1), 47-71.

Perron, P. and Ng, S. (1996) Useful Modifications to Some Unit Root Tests with Dependent Errors and their Local Asymptotic Properties, *Review of Economic Studies*, **63** (3), 435-463.

Perron, P. and Qu, Z. (2007) A Simple Modification to Improve Finite-Sample Properties of Ng and Perron's Unit Root Tests, *Economics Letters*, **94** (1), 12-19.

Phillips, P.C.B. and Perron, P. (1988) Testing for Unit Roots in Time Series, *Biometrika*, **75** (2), 335-346.

Pippenger, M.K. and Goering, G. (1993) A Note on the Empirical Power of Unit Root Tests Under Threshold Processes, *Oxford Bulletin of Economics and Statistics*, **55** (4), 473-481.

Restoy, F., Vallés, J. and López-Salido, J.D. (2005) Inflation Differentials in EMU: The Spanish Case, *Moneda y Crédito*, **220**, 55-104.

Romero-Ávila, D. y Usabiaga, C. (2008) On the Persistence of Spanish Unemployment Rates, *Empirical Economics*, **35** (1), 77-99.

Romero-Ávila, D. and Usabiaga, C. (2009) The Hypothesis of a Unit Root in OECD Inflation Revisited, *Journal of Economics and Business*, **61** (2), 153-161.

Sargan, J.D. and Bhargava, A. (1983) Testing Residuals from Least Squares Regression for Being Generated by the Gaussian Random Walk, *Econometrica*, **51** (1), 153-174.

Stock, J. (1999) A Class of Tests for Integration and Cointegration, in McFadden, D.L. and Engle, R.F. (Eds.) *Cointegration, Causality and Forecasting: A Festschrift in Honor of Clive W.J. Granger*, Oxford University Press, Oxford, 137-167.

Vermeulen, P., Dias, D., Dossche, M., Gautier, E., Hernando, I., Sabbatini, R. and Stahl, H. (2007) Price Setting in the Euro Area. Some Stylised Facts from Individual Producer Price Data, European Central Bank, Working Paper Series, nº 727, February.

TABLES

Table 1: Unit Root Analysis of the Inflation Rate: CPI of the Spanish Regions and Spain. 1979M1-2008M4.

	k_{MAIC}[a]	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	ADF^{GLS}	P_{T}^{GLS}	MP_{T}^{GLS}	t_{NL}
Spain	13	-1.33	-0.80	0.60	-0.97	18.93	17.91	-1.47
1. Andalucía	12	0.32	0.42	1.33	0.44	113.87	101.41	-2.28
2. Aragón	15	0.42	0.66	1.59	0.65	165.99	146.17	-2.72*
3. Asturias	12	0.35	0.48	1.37	0.41	121.37	108.01	-3.25**
4. Baleares	13	0.51	1.04	2.03	1.11	271.49	238.38	-1.84
5. Canarias	12	0.06	0.06	1.02	0.08	66.01	59.34	-3.24**
6. Cantabria	12	0.42	0.80	1.91	0.81	235.53	206.17	-3.23**
7. Castilla-León	12	0.34	0.48	1.43	0.46	131.88	116.64	-2.02
8. Castilla - La Mancha	12	0.24	0.29	1.20	0.19	92.61	82.61	-2.50
9. Cataluña	16	0.37	0.47	1.26	0.45	105.36	93.28	-2.25
10. Comunidad Valenciana	16	0.35	0.46	1.32	0.50	113.83	101.28	-1.91
11. Extremadura	12	0.29	0.37	1.28	0.37	106.05	94.35	-1.85
12. Galicia	12	0.34	0.46	1.36	0.55	119.47	106.17	-2.05
13. Madrid	12	0.48	0.85	1.75	0.85	202.65	178.33	-3.31**
14. Murcia	16	0.16	0.16	0.99	0.20	64.15	57.46	-1.75
15. Navarra	13	0.52	0.94	1.82	0.98	219.05	192.99	-2.23
16. País Vasco	12	0.39	0.55	1.39	0.54	125.60	111.79	-1.50
17. La Rioja	13	0.37	0.49	1.30	0.51	112.23	99.03	-1.33
Critical Values								
1%		-13.8	-2.58	0.174	-2.58	1.78	1.78	-3.48
5%		-8.1	-1.98	0.233	-1.98	3.17	3.17	-2.93
10%		-5.7	-1.62	0.275	-1.62	4.45	4.45	-2.66

[a] Lag length *k* is determined according to MAIC proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007). ***,** and * imply rejection of the null hypothesis of a unit root at the 1%, 5% and 10% significance levels, respectively.

Table 2: Unit Root Analysis of the Inflation Rate: CPI of the Spanish Provinces. 1971M1-2008M4.

	$k_{MAIC}{}^{a}$	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	ADF^{GLS}	P_{T}^{GLS}	MP_{T}^{GLS}	t_{NL}
1. Álava	12	-3.09	-1.23	0.40	-1.21	7.94	7.91	-1.52
2. Albacete	14	-2.94	-1.15	0.39	-1.10	8.28	8.20	-2.07
3. Alicante	14	-3.12	-1.22	0.39	-1.16	7.78	7.82	-1.96
4. Almería	12	-3.10	-1.17	0.38	-1.14	7.95	7.80	-1.64
5. Asturias	15	-1.98	-0.90	0.46	-0.83	11.56	11.43	-1.73
6. Ávila	13	-5.33	-1.61	0.30	-1.52	4.65	4.67	-2.84*
7. Badajoz	16	-2.69	-1.14	0.42	-1.11	8.99	9.01	-1.51
8. Baleares	16	-2.51	-1.10	0.44	-1.05	9.61	9.63	-1.85
9. Barcelona	12	-2.17	-1.02	0.47	-1.01	11.21	11.15	-1.31
10. Burgos	15	-2.81	-1.12	0.40	-1.08	8.58	8.54	-1.85
11. Cáceres	17	-4.17	-1.42	0.34	-1.34	5.89	5.91	-2.42
12. Cádiz	13	-2.25	-0.94	0.42	-0.95	10.55	10.05	-1.98
13. Cantabria	13	-2.26	-1.00	0.44	-0.95	10.47	10.43	-2.11
14. Castellón	16	-2.08	-1.02	0.49	-0.98	11.84	11.72	-2.30
15. Ciudad Real	17	-3.30	-1.26	0.38	-1.19	7.37	7.40	-2.15
16. Córdoba	14	-4.26	-1.42	0.33	-1.35	5.79	5.81	-2.52
17. Coruña	14	-2.58	-1.13	0.44	-1.13	9.78	9.51	-1.47
18. Cuenca	12	-5.00	-1.56	0.31	-1.49	4.95	4.97	-1.94
19. Gerona	14	-4.40	-1.48	0.34	-1.43	5.60	5.58	-2.04
20. Granada	12	-2.66	-1.08	0.41	-1.04	8.95	8.94	-1.66
21. Guadalajara	17	-3.40	-1.30	0.38	-1.25	7.25	7.21	-2.06
22. Guipúzcoa	13	-3.18	-1.24	0.39	-1.20	7.69	7.69	-1.81
23. Huelva	12	-1.13	-0.61	0.54	-0.59	17.32	16.74	-1.39
24. Huesca	16	-3.64	-1.30	0.36	-1.23	6.75	6.74	-1.66
25. Jaén	13	-2.37	-1.01	0.42	-0.97	9.93	9.86	-1.06
26. León	13	-4.18	-1.41	0.34	-1.35	5.93	5.92	-1.66
27. Lérida	14	-3.55	-1.32	0.37	-1.29	6.89	6.90	-1.84
28. Lugo	15	-3.24	-1.26	0.39	-1.26	7.55	7.56	-2.12
29. Madrid	17	-3.64	-1.30	0.36	-1.21	6.72	6.74	-2.15
30. Málaga	12	-3.06	-1.17	0.38	-1.11	7.95	7.88	-1.08
31. Murcia	13	-2.45	-1.09	0.45	-1.10	10.02	9.95	-2.30
32. Navarra	12	-3.27	-1.24	0.38	-1.22	7.43	7.47	-1.77
33. Orense	17	-1.97	-0.96	0.49	-0.97	12.03	12.09	-0.98
34. Palencia	17	-2.14	-0.94	0.44	-0.88	10.85	10.70	-1.46
35. Las Palmas	14	-1.65	-0.82	0.50	-0.75	13.61	13.38	-1.70
36. Pontevedra	16	-3.56	-1.32	0.37	-1.35	6.90	6.89	-1.80
37. Rioja	12	-4.13	-1.42	0.34	-1.39	5.97	5.97	-1.74
38. Salamanca	15	-2.35	-1.06	0.45	-1.00	10.23	10.26	-2.09
39. S. Cruz Tenerife	17	-0.99	-0.55	0.56	-0.48	19.00	17.87	-1.23
40. Segovia	12	-1.61	-0.81	0.50	-0.77	13.78	13.62	-1.35
41. Sevilla	14	-3.31	-1.26	0.38	-1.22	7.36	7.39	-2.12
42. Soria	12	-2.21	-1.04	0.47	-1.03	10.97	10.98	-2.35
43. Tarragona	12	-2.32	-1.07	0.46	-1.07	10.66	10.53	-1.71
44. Teruel	17	-3.53	-1.31	0.37	-1.28	6.92	6.95	-1.14
45. Toledo	14	-1.25	-0.73	0.58	-0.68	17.70	17.75	-1.01
46. Valencia	15	-2.01	-0.95	0.47	-0.96	11.67	11.71	-1.76

47. Valladolid	12	-2.63	-1.09	0.41	-1.03	9.11	9.09	-1.35
48. Vizcaya	14	-4.79	-1.51	0.32	-1.43	5.17	5.20	-2.37
49. Zamora	12	-4.90	-1.54	0.32	-1.48	5.03	5.05	-2.00
50. Zaragoza	12	-2.93	-1.20	0.41	-1.16	8.35	8.33	-1.69
51. Ceuta	15	-3.06	-1.20	0.39	-1.16	7.93	7.96	-1.39
52. Melilla	14	-2.61	-1.05	0.40	-1.03	9.19	9.01	-1.61
Critical Values								
1%		-13.8	-2.58	0.174	-2.58	1.78	1.78	-3.48
5%		-8.1	-1.98	0.233	-1.98	3.17	3.17	-2.93
10%		-5.7	-1.62	0.275	-1.62	4.45	4.45	-2.66

[a] Lag length k is determined according to MAIC proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007). ***,** and * imply rejection of the null hypothesis of a unit root at the 1%, 5% and 10% significance levels, respectively.

Table 3: Unit Root Analysis of the Inflation Rate: CPI of the Main Groups of Goods and Services. 1978M1-2000M12.

	k_{MAIC}[a]	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	ADF^{GLS}	P_{T}^{GLS}	MP_{T}^{GLS}	t_{NL}
G1. Food, beverages and tobacco	12	0.28	0.30	1.07	0.21	79.63	67.54	-1.51
G2. Footwear and clothing	14	0.70	1.46	2.09	1.33	310.26	263.70	-6.87***
G3. Housing	12	0.22	0.22	1.01	0.11	67.52	59.98	-1.91
G4. Cookware and home services	13	0.57	1.07	1.87	0.93	244.31	206.93	-4.23***
G5. Medicine and healthcare	15	0.73	1.25	1.71	1.12	212.05	180.79	-1.55
G6. Transport and communication	13	0.29	0.37	1.25	0.27	106.52	90.50	-2.02
G7. Leisure, education and culture	12	0.61	1.06	1.73	1.01	212.15	179.90	-1.78
G8. Other goods and services	13	0.33	0.34	1.04	0.37	74.44	64.45	-2.49
Critical Values								
1%		-13.8	-2.58	0.174	-2.58	1.78	1.78	-3.48
5%		-8.1	-1.98	0.233	-1.98	3.17	3.17	-2.93
10%		-5.7	-1.62	0.275	-1.62	4.45	4.45	-2.66

[a] Lag length k is determined according to MAIC proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007). ***,** and * imply rejection of the null hypothesis of a unit root at the 1%, 5% and 10% significance levels, respectively.

Table 4: Unit Root Analysis of the Inflation Rate: Aggregate PPI and Main Manufacturing and Energy Sectors. 1976M1-2002M12.

	k_{MAIC}[a]	MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	ADF^{GLS}	P_{T}^{GLS}	MP_{T}^{GLS}	t_{NL}
Aggregate PPI	13	-3.41	-1.30	0.38	-1.31	7.32	7.18	-2.33
Sector 11	14	0.42	0.52	1.22	0.18	101.95	88.28	-4.22***
Sector 13	12	-9.68**	-2.19**	0.23	-2.08**	2.73**	2.58**	-3.61***
Sector 15	12	-0.33	-0.22	0.68	-0.41	31.66	27.61	-2.11
Sector 21	1	-24.81***	-3.51***	0.14***	-3.57***	1.03***	1.02***	-7.26***
Sector 22	13	-7.86*	-1.98*	0.25*	-2.16**	3.26*	3.13**	-3.52***
Sector 23	13	-1.56	-0.73	0.47	-0.77	13.93	13.02	-2.38
Sector 24	15	-4.48	-1.48	0.32	-1.56	5.76	5.63	-3.39**
Sector 25	13	-6.65*	-1.81*	0.27*	-1.72*	3.71*	3.73*	-2.54
Sector 31	16	-1.25	-0.77	0.62	-3.07***	19.31	18.96	-2.35
Sector 32	16	0.03	0.02	0.75	-0.09	36.96	34.51	-2.32
Sector 33	16	0.28	0.29	1.05	0.37	71.57	64.72	-0.61
Sector 34	12	-0.23	-0.15	0.64	-0.73	27.43	25.82	-2.38
Sector 35	14	0.35	0.38	1.09	0.32	79.68	70.63	-2.62
Sector 36	16	0.37	0.41	1.09	0.15	78.85	71.81	-3.28**
Sector 37	12	-3.30	-1.21	0.37	-1.39	7.51	7.38	-2.07
Sector 39	16	-0.55	-0.34	0.61	-0.29	24.77	22.28	-3.90***
Sector 41	13	-5.90*	-1.69*	0.29	-2.25**	4.23*	4.23*	-4.00***
Sector 43	13	-2.00	-0.86	0.43	-1.20	11.06	10.86	-1.28
Sector 44	15	-9.50**	-2.17**	0.23**	-2.62***	2.63**	2.64**	-3.40**
Sector 45	16	-1.37	-0.71	0.52	-1.31	15.65	15.07	-3.18**
Sector 46	12	-7.00*	-1.84*	0.26*	-2.66***	3.61*	3.63*	-2.81*
Sector 47	15	-7.45*	-1.92*	0.26*	-2.21**	3.35*	3.32	-3.24**
Sector 48	12	-2.44	-1.02	0.42	-1.37	9.61	9.59	-2.01
Sector 49	14	-8.85**	-2.10**	0.24*	-2.15**	2.88**	2.79**	-4.85***
Critical Values								
1%		-13.8	-2.58	0.174	-2.58	1.78	1.78	-3.48
5%		-8.1	-1.98	0.233	-1.98	3.17	3.17	-2.93
10%		-5.7	-1.62	0.275	-1.62	4.45	4.45	-2.66

[a] Lag length *k* is determined according to MAIC proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007). ***,** and * imply rejection of the null hypothesis of a unit root at the 1%, 5% and 10% significance levels, respectively.

Table 5: Persistence Analysis of the Inflation Rate: CPI of the Spanish Regions and Spain, 1979M1-2008M4.

	p_{MAIC}[a]	α-OLS[b]	α-MU[c]	MU [95% CI][d]	HL_{IRF}[e]	HL[95% CI][f]
Spain	13	0.994	0.997	[0.991, 1.000]	8.980	[4,923, ∞]
1. Andalucía	12	0.986	1.000	[0.994, 1.000]	3.683	[2.916, ∞]
2. Aragón	15	0.982	1.000	[0.995, 1.000]	3.006	[2.137, ∞]
3. Asturias	12	0.982	1.000	[0.994, 1.000]	2.965	[0.987, ∞]
4. Baleares	13	0.980	1.000	[0.996, 1.000]	2.243	[2.169, ∞]
5. Canarias	12	0.987	1.000	[0.991, 1.000]	3.728	[2.138, ∞]
6. Cantabria	12	0.977	1.000	[0.996, 1.000]	2.163	[0.958, ∞]
7. Castilla-León	12	0.983	1.000	[0.995, 1.000]	2.990	[1.589, ∞]
8. Castilla - La Mancha	12	0.981	1.000	[0.992, 1.000]	2.971	[1.547, ∞]
9. Cataluña	16	0.986	1.000	[0.994, 1.000]	4.224	[2.191, ∞]
10. Com. Valenciana	16	0.987	1.000	[0.995, 1.000]	4.283	[2.233, ∞]
11. Extremadura	12	0.986	1.000	[0.994, 1.000]	3.782	[2.204, ∞]
12. Galicia	12	0.986	1.000	[0.995, 1.000]	3.455	[2.932, ∞]
13. Madrid	12	0.980	1.000	[0.996, 1.000]	2.907	[0.968, ∞]
14. Murcia	16	0.986	1.000	[0.992, 1.000]	4.259	[2.844, ∞]
15. Navarra	13	0.981	1.000	[0.996, 1.000]	2.154	[0.964, ∞]
16. País Vasco	12	0.990	1.000	[0.995, 1.000]	4.983	[2.928, ∞]
17. La Rioja	13	0.984	1.000	[0.994, 1.000]	3.019	[1.009, ∞]

[a] Lag length p is determined according to the MAIC procedure proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007).

[b] OLS estimate of the sum of the AR coefficients.

[c] Median unbiased estimate of the sum of the AR coefficients computed following Gospodinov (2004).

[d] Median unbiased estimate of the confidence intervals for the persistence parameter computed at the 95% confidence level following the procedure of Hansen (1999).

[e] Estimate of the half-life measured in years based on the impulse-response function as in Gospodinov (2004).

[f] Median unbiased estimate of the confidence intervals for the half-life measured in years computed at the 95% confidence level following the procedure of Gospodinov (2004).

Table 6: Persistence Analysis of the Inflation Rate: CPI of the Spanish Provinces, 1971M1-2008M4.

	p_{MAIC}[a]	α-OLS[b]	α-MU[c]	MU [95% CI][d]	HL_{IRF}[e]	HL[95% CI][f]
1. Álava	12	0.993	0.995	[0.984, 1.000]	7.719	[3.760, ∞]
2. Albacete	14	0.991	0.994	[0.978, 1.000]	5.367	[2.969, ∞]
3. Alicante	14	0.992	0.995	[0.981, 1.000]	6.460	[3.357, ∞]
4. Almería	12	0.991	0.993	[0.979, 1.000]	6.121	[3.341, ∞]
5. Asturias	15	0.995	0.997	[0.986, 1.000]	11.186	[4.239, ∞]
6. Ávila	13	0.989	0.992	[0.977, 1.000]	5.383	[2.918, ∞]
7. Badajoz	16	0.994	0.996	[0.984, 1.000]	8.291	[3.708, ∞]
8. Baleares	16	0.992	0.996	[0.981, 1.000]	5.377	[2.891, ∞]
9. Barcelona	12	0.993	0.997	[0.984, 1.000]	6.933	[3.511, ∞]
10. Burgos	15	0.993	0.995	[0.982, 1.000]	7.296	[3.695, ∞]
11. Cáceres	17	0.991	0.994	[0.981, 1.000]	6.795	[3.589, ∞]
12. Cádiz	13	0.989	0.994	[0.979, 1.000]	5.186	[2.959, ∞]
13. Cantabria	13	0.994	0.997	[0.985, 1.000]	9.410	[4.259, ∞]
14. Castellón	16	0.992	0.996	[0.982, 1.000]	5.429	[2.863, ∞]
15. Ciudad Real	17	0.992	0.995	[0.982, 1.000]	7.225	[3.810, ∞]
16. Córdoba	14	0.990	0.993	[0.979, 1.000]	5.705	[3.042, ∞]
17. Coruña	14	0.991	0.995	[0.983, 1.000]	6.351	[3.633, ∞]
18. Cuenca	12	0.989	0.992	[0.977, 1.000]	5.049	[2.970, ∞]
19. Gerona	14	0.989	0.992	[0.978, 1.000]	5.047	[2.959, ∞]
20. Granada	12	0.993	0.995	[0.980, 1.000]	6.433	[3.442, ∞]
21. Guadalajara	17	0.991	0.994	[0.980, 1.000]	6.088	[3.436, ∞]
22. Guipúzcoa	13	0.993	0.995	[0.984, 1.000]	7.702	[4.195, ∞]
23. Huelva	12	0.993	0.997	[0.981, 1.000]	4.458	[2.453, ∞]
24. Huesca	16	0.990	0.993	[0.977, 1.000]	5.343	[2.981, ∞]
25. Jaén	13	0.992	0.995	[0.980, 1.000]	5.841	[3.088, ∞]
26. León	13	0.992	0.994	[0.982, 1.000]	7.289	[3.760, ∞]
27. Lérida	14	0.990	0.994	[0.979, 1.000]	5.131	[3.069, ∞]
28. Lugo	15	0.989	0.992	[0.975, 1.000]	4.117	[2.342, ∞]
29. Madrid	17	0.993	0.996	[0.984, 1.000]	9.111	[4.912, ∞]
30. Málaga	12	0.991	0.994	[0.977, 1.000]	5.659	[2.967, ∞]

31. Murcia	13	0.992	0.996	[0.982, 1.000]	5.703	[3.600, ∞]
32. Navarra	12	0.993	0.995	[0.984, 1.000]	8.573	[4.193, ∞]
33. Orense	17	0.993	0.997	[0.981, 1.000]	6.570	[3.760, ∞]
34. Palencia	17	0.994	0.996	[0.983, 1.000]	8.345	[3.734, ∞]
35. Las Palmas	14	0.994	0.997	[0.983, 1.000]	7.754	[3.603, ∞]
36. Pontevedra	16	0.988	0.991	[0.973, 1.000]	3.672	[2.297, ∞]
37. Rioja	12	0.990	0.993	[0.978, 1.000]	5.104	[2.930, ∞]
38. Salamanca	15	0.993	0.996	[0.982, 1.000]	5.463	[2.269, ∞]
39. Santa Cruz Tenerife	17	0.990	0.997	[0.983, 1.000]	4.923	[2.937, ∞]
40. Segovia	12	0.994	0.997	[0.983, 1.000]	7.270	[3.260, ∞]
41. Sevilla	14	0.993	0.995	[0.983, 1.000]	7.649	[3.689, ∞]
42. Soria	12	0.993	0.996	[0.982, 1.000]	5.393	[2.926, ∞]
43. Tarragona	12	0.993	0.996	[0.985, 1.000]	7.587	[3.628, ∞]
44. Teruel	17	0.987	0.991	[0.970, 1.000]	3.717	[1.448, ∞]
45. Toledo	14	0.994	0.998	[0.979, 1.000]	2.919	[2.135, ∞]
46. Valencia	15	0.994	0.997	[0.984, 1.000]	7.858	[4.920, ∞]
47. Valladolid	12	0.994	0.996	[0.983, 1.000]	8.503	[4.036, ∞]
48. Vizcaya	14	0.992	0.994	[0.984, 1.000]	7.917	[4.427, ∞]
49. Zamora	12	0.991	0.993	[0.982, 1.000]	6.662	[3.727, ∞]
50. Zaragoza	12	0.994	0.996	[0.986, 1.000]	9.447	[4.752, ∞]
51. Ceuta	15	0.991	0.994	[0.978, 1.000]	4.147	[3.387, ∞]
52. Melilla	14	0.990	0.992	[0.974, 1.000]	3.735	[2.842, ∞]

[a] Lag length p is determined according to the MAIC procedure proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007).
[b] OLS estimate of the sum of the AR coefficients.
[c] Median unbiased estimate of the sum of the AR coefficients computed following Gospodinov (2004).
[d] Median unbiased estimate of the confidence intervals for the persistence parameter computed at the 95% confidence level following the procedure of Hansen (1999).
[e] Estimate of the half-life measured in years based on the impulse-response function as in Gospodinov (2004).
[f] Median unbiased estimate of the confidence intervals for the half-life measured in years computed at the 95% confidence level following the procedure of Gospodinov (2004).

Table 7: Persistence Analysis of the Inflation Rate: CPI of the Main Groups of Goods and Services, 1978M1-2000M12.

	p_{MAIC}[a]	α-OLS[b]	α-MU[c]	MU [95% CI][d]	HL_{IRF}[e]	HL[95% CI][f]
G1. Food, beverages and tobacco	12	0.972	1.000	[0.992, 1.000]	1.679	[1.006, ∞]
G2. Footwear and clothing	14	0.992	1.000	[0.999, 1.000]	7.497	[3.724, ∞]
G3. Housing	12	0.983	1.000	[0.988, 1.000]	2.966	[0.982, ∞]
G4. Cookware and home services	13	0.993	1.000	[0.999, 1.000]	8.184	[3.129, ∞]
G5. Medicine and healthcare	15	0.992	1.000	[0.997, 1.000]	5.902	[2.202, ∞]
G6. Transport and communication	13	0.979	1.000	[0.990, 1.000]	2.601	[0.991, ∞]
G7. Leisure, education and culture	12	0.986	1.000	[0.997, 1.000]	4.284	[1.673, ∞]
G8. Other goods and services	13	0.987	1.000	[0.992, 1.000]	4.751	[2.769, ∞]

[a] Lag length p is determined according to the MAIC procedure proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007).

[b] OLS estimate of the sum of the AR coefficients.

[c] Median unbiased estimate of the sum of the AR coefficients computed following Gospodinov (2004).

[d] Median unbiased estimate of the confidence intervals for the persistence parameter computed at the 95% confidence level following the procedure of Hansen (1999).

[e] Estimate of the half-life measured in years based on the impulse-response function as in Gospodinov (2004).

[f] Median unbiased estimate of the confidence intervals for the half-life measured in years computed at the 95% confidence level following the procedure of Gospodinov (2004).

Table 8: Persistence Analysis of the Inflation Rate: Aggregate PPI and Main Manufacturing and Energy Sectors, 1976M1-2002M12.

	p_{MAIC}[a]	α-OLS[b]	α-MU[c]	MU [95% CI][d]	HL_{IRF}[e]	HL[95% CI][f]
Aggregate PPI	13	0.992	0.994	[0.983, 1.000]	7.091	[3.818, ∞]
Sector 11	14	0.974	1.000	[0.990, 1.000]	0.975	[0.952, ∞]
Sector 13	12	0.942	0.961	[0.927, 0.995]	1.202	[0.995, ∞]
Sector 15	12	0.959	0.996	[0.981, 1.000]	1.375	[0.958, ∞]
Sector 21	1	0.910	0.915	[0.867, 0.961]	0.596	[0.413, 2.032]
Sector 22	13	0.972	0.978	[0.960, 0.999]	1.632	[1.442, ∞]
Sector 23	13	0.984	0.992	[0.968, 1.000]	2.930	[2.025, ∞]
Sector 24	15	0.988	0.988	[0.972, 1.000]	4.841	[2.842, ∞]
Sector 25	13	0.988	0.991	[0.977, 1.000]	5.174	[3.200, ∞]
Sector 31	16	0.988	0.988	[0.979, 0.997]	5.042	[3.003, ∞]
Sector 32	16	0.993	0.999	[0.987, 1.000]	7.647	[3.546, ∞]
Sector 33	16	0.981	1.000	[0.973, 1.000]	2.363	[0.342, ∞]
Sector 34	12	0.988	0.994	[0.983, 1.000]	4.493	[2.925, ∞]
Sector 35	14	0.973	1.000	[0.988, 1.000]	2.418	[2.035, ∞]
Sector 36	16	0.979	1.000	[0.985, 1.000]	0.943	[0.928, ∞]
Sector 37	12	0.975	0.983	[0.953, 1.000]	2.363	[0.953, ∞]
Sector 39	16	0.976	0.996	[0.978, 1.000]	2.262	[0.974, ∞]
Sector 41	13	0.987	0.988	[0.976, 1.000]	4.237	[1.790, ∞]
Sector 43	13	0.991	0.994	[0.982, 1.000]	6.853	[3.643, ∞]
Sector 44	15	0.980	0.981	[0.966, 0.997]	2.990	[1.460, ∞]
Sector 45	16	0.986	0.990	[0.978, 1.000]	4.933	[3.052, ∞]
Sector 46	12	0.983	0.985	[0.971, 0.998]	3.633	[2.613, ∞]
Sector 47	15	0.982	0.985	[0.970, 1.000]	3.469	[2.977, ∞]
Sector 48	12	0.991	0.994	[0.980, 1.000]	6.671	[3.771, ∞]
Sector 49	14	0.960	0.967	[0.940, 0.997]	1.091	[0.997, ∞]

[a] Lag length p is determined according to the MAIC procedure proposed by Ng and Perron (2001) with the finite-sample modification by Perron and Qu (2007).

[b] OLS estimate of the sum of the AR coefficients.
[c] Median unbiased estimate of the sum of the AR coefficients computed following Gospodinov (2004).
[d] Median unbiased estimate of the confidence intervals for the persistence parameter computed at the 95% confidence level following the procedure of Hansen (1999).
[e] Estimate of the half-life measured in years based on the impulse-response function as in Gospodinov (2004).
[f] Median unbiased estimate of the confidence intervals for the half-life measured in years computed at the 95% confidence level following the procedure of Gospodinov (2004).

Appendix

Sectors (PPI):

11. Coke, petroleum and coal products.
13. Petroleum refineries.
15. Electricity, gas and water supply.
21. Iron and steel.
22. Basic metals and fabricated metal products.
23. Nonferrous and nonfuel metals, peat.
24. Nonmetal products.
25. Chemicals and chemicals products.
31. Metal products (except machinery and transportation equipment).
32. Machinery and mechanical equipment.
33. Office and computing (including installation).
34. Electric machinery.
35. Electronic equipment (except computers).
36. Motor vehicle and spare motor parts.
37. Other transport equipment.
39. Optical and precision equipment.
41. Food products, beverages and tobacco.
43. Textile.
44. Leather.
45. Textile products and footwear.
46. Wood, products of wood and cork.
47. Pulp, paper, paper products, printing and publishing.
48. Rubber and plastic products.
49. Other manufacturing industries.