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INFLATION PERSISTENCE DURING PERIODS OF STRUCTURAL CHANGE: *AN ASSESSMENT USING GREEK DATA*

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ABSTRACT

The paper estimates inflation persistence in Greece from 1975 to 2003, a period of high variation in inflation and changes in policy regimes. Three empirical methodologies, univariate autoregressive (AR) modelling, second-generation random coefficient (RC) modelling, and vector autoregressive (VAR) modelling, are employed to estimate inflation persistence. The empirical results from all the procedures suggest that inflation persistence was high during the inflationary period and the first six years of the disinflationary period, while it started to decline after 1997, when inflationary expectations seem to have been stabilised, and thus, monetary policy was effective at reducing inflation. Empirical findings also detect a sluggish response of inflation to changes in monetary policy. This observed delay seems to have changed little over time.

Keywords: CPI inflation; persistence; structural change.
JEL Classification: E31; E37.

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1. Introduction

Conceptually and empirically, considerable attention has focused in recent years on the idea that the degree of inflation persistence depends on the monetary policy regime. Empirical studies, mainly dealing with the US economy and other industrial economies and applying different empirical methodologies, have found abundant evidence that post-war inflation in these countries exhibits very high persistence. In this connection, Alogoskoufis and Smith (1991), looking at historical data for US and UK inflation, point out that the post-Bretton Woods regime of managed floating was associated with more persistent inflation. Bordo (1993) and Bordo and Schwartz (1999) find that inflation was more stable during the heyday of Bretton Woods. More recently, Benati (2002) argues that the degree of inflation persistence in the US and the UK was rather low during the gold standard, peaked in the 1970s, and fell again in the 1990s. Batini (2002), for the euro area countries, and Levin and Piger (2004), for twelve industrial countries, find little or no evidence of an upward shift in inflation persistence in post-1980 data, while there was a substantial downward shift in the average value of inflation. In general, the empirical literature suggests that inflation persistence may not be an intrinsic structural phenomenon of industrial countries, '…but rather varies with the stability and transparency of the monetary policy regime' (Levin and Piger, 2004, p.6).

This paper focuses on the estimation of the time profile of inflation persistence in Greece applying different empirical methodologies.[1] Greece is a country with a high variance in inflation which has experienced a variety of policy regimes post-1975 that might have considerably affected the behaviour of inflation process both in the sense of the steady state inflation rate and its autocorrelation properties. In particular, during this period, major inflation peaks occurred, in 1979-80, 1985-86 and 1990; on each occasion, inflation rose to at least 20 per cent. By 1995, however, the annual rate of inflation had fallen for the first time during the period to single-digit levels and, by the late 1990s, had fallen even further to the lower single digits.

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[1] To the best of our knowledge, this paper is the first study of inflation persistence in Greece. In an earlier paper, Lazaretou (1995) presents some historical perspective on the behaviour of inflation in Greece in the pre- and post-WWI periods. The historical data of Greek price indices suggest that there was a positive association between the time series properties of inflation and the country's choice of nominal exchange rate regime. By employing non-parametric measures of persistence, Kalyvitis and Lazaretou (1997) examine the persistence of Greek inflation under the Bretton Woods system and the subsequent floating rate system. Statistics show that inflation exhibits higher persistence under floating.

We present a model-free assessment of the time profile of inflation persistence for the Greek economy. Specifically, based on three different definitions of persistence proposed in the literature, we show purely statistical evidence of the autocorrelation properties of the inflation series as well as of the sluggish response of inflation to systematic and non-systematic monetary policy changes using quarterly macro data for CPI inflation in Greece. In order to capture possible shifts in inflation persistence, we apply both classical and Bayesian analysis (second generation random coefficients (RC) estimates) to study the dynamic behaviour of the inflationary process post-1975. In particular, we estimate a univariate autoregressive (AR) model for the inflation rate and consider the possibility of a structural break in the slope parameter and the intercept term of the AR equation. The results from the classical analysis are confirmed by the RC estimates and the rolling regression methods. The advantage of second generation RC modelling is that it estimates a time varying inflation persistence and relaxes key restrictions, such as a specific functional form, often imposed in the empirical literature.[2] Thus, the evolution of the degree of inflation persistence can be modelled during a period of structural breaks. Finally, we present evidence on the lag in the response of inflation to monetary policy changes from impulse response analysis from a VAR model.

The Greek experience may be illustrative for the future of other medium-sized economies, such as the Central and Eastern European countries, entering the EU. The fulfilment of the inflation convergence criterion in order to join the euro area is a significant driving factor in policies aimed at disinflation in acceding countries. In fact, acceding countries have made substantial efforts to bring inflation close to the levels observed in the euro area. Over the last two years, they experienced a slight increase in headline inflation. The main inflation strategies are inflation targeting, a fixed peg or a currency board arrangement. Hence, the conclusions drawn on the basis of the Greek data could be very useful.

The remainder of the paper is organised as follows. Section 2 briefly overviews inflationary developments and changes in the monetary policy framework in post-1975 Greece. Section 3 discusses various definitions of inflation persistence, since there are various views in the economic literature. Section 4 contains the empirical results from univariate autoregressive estimates for inflation persistence

[2] For a detailed analysis on second generation RC estimates, see Swamy and Tavlas (1995, 2001).

using OLS and rolling regressions. Section 5 estimates CPI inflation persistence following the RC methodology. Section 6 presents evidence on the lag in the response of inflation to changes in monetary policy. Section 7 concludes.

2. Inflationary Developments and Changes in the Monetary Policy Framework: Narrative Evidence

Figure 1 plots the four-quarter changes in the CPI and depicts the major phases of Greek inflation. As the figure suggests, the inflation process has behaved differently through the various phases of Greek inflation; namely, the post-Bretton Woods period of high inflation in the 1970s and the 1980s, and the disinflation of the 1990s. In particular, after averaging 3.5 per cent, with little volatility, in the 20 years prior to 1973, the behaviour of the inflationary process changed substantially after the collapse of the Bretton Woods fixed exchange rate regime.

The pegging of the drachma to the dollar after the collapse of Bretton Woods, in conjunction with the loose economic policy pursued, resulted in the drachma's significant depreciation during the two-year period of 1972-73. This depreciation, combined with a rise in the prices of raw materials (first oil crisis) and the overheating of the economy, led to a rise in the inflation rate double digits. The drachma's uncoupling from the dollar in 1975 and its float against a basket of currencies amounted to a policy of rapid sliding depreciation in order to counterbalance the effect of wage increases on the country's international competitiveness. The process of economic recovery was interrupted, while inflation remained around 20% in the 1970s and the 1980s.

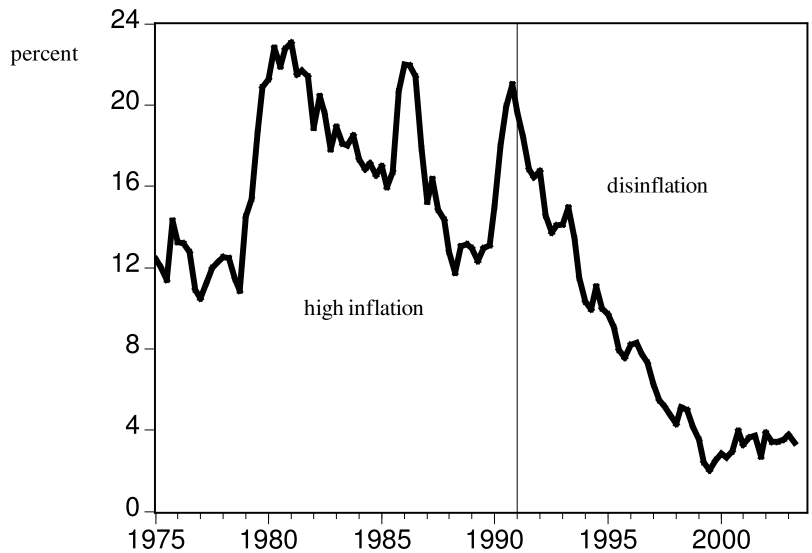
The drachma depreciated rapidly until 1987; thereafter, its sliding depreciation was limited. With the stabilisation programme of 1985 the government switched to an anti-inflationary policy. The limited depreciation of the drachma, the rigorous income policy and the tight monetary policy led to some inflation deceleration. However, the abandonment of the programme at the end of 1987 and the relaxation of wage controls caused a new inflation crisis at the end of the decade.

In the 1990s, economic policy was mainly anti-inflationary. Successive Greek governments had to check fiscal deficits and high inflation so as to meet the

Maastricht criteria, a necessary and sufficient condition for the convergence in nominal terms of the country to the other members of the EMU. The Bank of Greece changed its policy decisively towards fighting inflation, and its control became the major goal of monetary policy. The years 1991-94 marked a transition from a high inflation rate to a more moderate level (11%). Beginning in 1995, the Bank of Greece adopted, for the first time, a nominal exchange rate anchor as an important mechanism for disinflation. As a result, inflation fell to single digit levels for the first time since the early 1970s, reaching a trough of 2.0% in the third quarter of 1999. Thereafter, it increased and remained at a level around 3.5%.

The post-1975 period has also been characterised by changes in monetary policy regime.[3] Prior to 1990, monetary policy was not effective in reducing inflation. Subsequently, the effectiveness of monetary policy increased while deregulation and subsequent development of the Greek financial market was set in

Figure 1
Price Inflation in Greece, 1975-2003



Note: The inflation rate has been computed as 4-month percentage change based on quarterly CPI averages.

motion and, in a process that began in 1995, monetary policy was eventually successful in curbing price pressures and bringing inflation down to a level consistent with the Maasrticht criterion.

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[3] For a detailed analysis on inflation performance during this period, see Garganas and Tavlas (2001). For a brief chronology of the major monetary policy changes, see the Annex Box at the end of the paper.

After substantial efforts, Greece met the Maastricht criteria and the drachma successfully entered the euro area on 1^{st} January 2001. Despite strong growth increased fiscal imbalances, over the last two years, have contributed to price pressures. Consumer price inflation is forecast to remain high standing at 3.6% in 2003 and 3.7% in 2004, more than 1.5 percentage points above that of the euro area.

3. Definitions of Inflation Persistence

Following the typology proposed by Batini and Nelson (2001) and Batini (2002), measures of inflation persistence refer, first, to positive serial correlation in inflation series, second, to lags between systematic monetary policy actions and their peak effect on inflation, and third, to lagged responses of inflation to policy shocks. To quote Batini (2002), ‘the first type of persistence is a reduced-form property of inflation that manifests simultaneously the underlying pricing process, the conduct of monetary policy and the expectations’ formation process of price-setting agents. Changes of any of these three factors will influence the autocorrelation properties of inflation’. The second measure of persistence, i.e. the length of the delay between a monetary policy action and its maximum effect on inflation, ‘…determines the costs of disinflation’, while the third measure, i.e. the lagged inflation response to a monetary policy shock, ‘…is often the only one consulted by economic modellers when validating models vis-a-vis the dynamics of real-world data generating process’ (pp.11-12).

Willis (2003, p.7) defines inflation persistence as the ‘speed with which inflation returns to baseline after a shock’. This definition implies that the degree of inflation persistence shows the speed with which inflation responds to a shock. When the value is high, inflation responds quickly to a shock. On the contrary, when the value is small, the speed of adjustment is low, the response of inflation to a shock is slow and the overall variability of inflation is small.

However, concerning the normative aspects of persistence, inflation persistence does not have the same connotation in periods of high inflation as in periods of deflation or price stability. Persistence in high inflation periods has a negative connotation, whereas in low inflation periods it may not. Explicitly, in inflationary periods high inflation rates are accompanied by high inflation in the future (vicious circle), whereas in disinflation the inertia of inflationary expectations breaks down and inflation is steadily falling (virtuous circle). As measured in the empirical literature, persistence would be high even in the extreme case of complete price stability. This surely has to be distinguished from inflation persisting after a shock.

4. Univariate Autoregressive Estimates

We apply classical analysis, based on the approach by Batini (2002) and Levin and Piger (2004), to study the dynamic behaviour of quarterly CPI inflation data for Greece. We are interested in assessing whether both the steady state of inflation and its serial correlation have changed over time.

4.1 Classical Analysis

To study the autocorrelation properties of inflation, we consider a simple AR(k) process for the inflation series, i.e.

$$\Delta p_t = \text{constant} + \sum_{j=1}^{k} \alpha_j \Delta p_{t-j} + u_t \qquad (1)$$

where Δp_t is annualised quarterly inflation, i.e. $\Delta p_t=(p_t-p_{t-1})*4$, p_t is the logarithm of the CPI, α_j is the autoregressive coefficient and u_t is a serially uncorrelated, but possibly heteroskedastic, random error term.

As in Andrews and Chen (1994), a measure of the degree of inflation persistence is the sum of the estimated lagged-terms autoregressive coefficients, i.e. $\rho \equiv \sum\alpha_j$, where ρ indicates the parameter of persistence. To measure persistence in terms of the sum of AR(k) coefficients, we re-write equation (1) as follows,

$$\Delta p_t = \text{constant} + \rho\Delta p_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta\Delta p_{t-j} + u_t \qquad (2)$$

where ρ is the persistence parameter, while β_j parameters are transformations of AR coefficients in equation (1), $\beta_{k-1}= -\alpha_k$. The inflation process has a unit root if ρ takes a value close to unity. The parameter ρ provides a reduced-form measure of inflation persistence, namely it directly measures the sluggishness with which the inflation process responds to shocks.

Estimating equation (2), we obtain a measure of inflation persistence, while the regression on a constant produces estimates of the sample mean of inflation and its standard deviation. Equation (2) requires a AR(k) lag order to be chosen, with a maximum lag length of k=4, given the quarterly frequency of the data. Based on both Akaike Information Criterion and Schwarz Criterion, we choose a lag of 4 for the inflation series. There is some evidence of a change in monetary conditions in the very early 1990s. This change can be considered as a shift towards a more aggressive approach towards fighting inflation and hence a breakpoint date in the time-series properties of inflation. If we had *a priori* knowledge of the date of the structural break, then we could run split-sample autoregressions for the inflation series and then confirm the structural break by using a Chow test. In the absence of *a priori* knowledge, we experiment with a variety of possible dates using the Chow test. The results support the timing of the split in the first quarter of 1991.[4]

In addition, we test for the case of multiple breaks using the test developed by Bai and Perron (1998, 2003).[5] Bai and Peron (1998) suggested a method based on sequential testing of the null hypothesis of l breaks against the alternative of l+1 breaks employing the sup $F_T(l+1/l)$ type test. In addition, Bai and Perron (2003) showed that the performance of this procedure could be improved by estimating the double maximum tests, namely UD max $F_T(M,q)$ and WD max $F_T(M,q)$, where M is the maximum number of breaks and p is the number of parameters allowed to structural change, in order to detect at least one break point. UD and WD are used to

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[4] That is, the break point Chow test and recursive coefficient estimates. In running Chow break point tests for all potential break point dates, we allow for a change both in the intercept and in the autoregressive coefficients. By allowing different intercept and slope coefficients, equation (2) is fitted separately to different sub-samples and then we compare them to the restricted (full-sample) equation by means of a Chow test. However, the timing of a structural change in the inflation series need not have coincided precisely with the formal adoption of a new policy regime. By recursive coefficient estimates we are able to trace the evolution of the persistence coefficient as more and more of the sample data are added to the estimating regression. Recursive estimates suggest that although there is no movement outside the critical values of the 2 percent standard error bands, persistence increases till the mid-1990s, while the sample mean value of the inflation rate declines.

[5] Over the last decade the analysis of structural breaks in a time series has considerably developed. Different test statistics are proposed in the literature detecting single and multiple breaks. Properly accounting for them is of great importance in the analysis of the time series properties of a macroeconomic variable, such as shifts in the mean rate of inflation and the measure of its persistence. See Altissimo and Corradi (2003) and Altissimo (2003).

test the null hypothesis of no structural break against an unknown number of breaks given some upper bound M. In our empirical analysis, first, we estimate the sup $F_T(l)$ for l=1,2,3,4,5. The estimated sup $F_T(l)$ tests are all significant at the 5% level for all l, (sup $F_T(1)$=34.47, sup $F_T(2)$=76.25, etc.). Therefore, we conclude that the series appears to have at least one structural break. Next, the UD max and WD max are estimated, which are also significant at the 5% level, indicating the existence of at least one break point. Finally, the sequential procedure is employed. The estimated sup $F_T(2/1)$=0.0789 using 5% level of significance suggests one break point after the last quarter of 1990. On the basis of these statistics, we can detect one break point between the last quarter of 1990 and the first quarter of 1991. The regressions, therefore, are estimated over the two sub-samples 1975 Q1-1990 Q4 and 1991 Q1-2003 Q2.

Table 1 presents the estimates of persistence and the sample mean in the form of univariate representations of annualised quarterly inflation series. As can be seen in Table 1, Greece primarily experienced a sizeable shift in the average level of CPI inflation but only a small shift in persistence. ρ rises from just a little bit over 0.70 before 1991 to 0.78 thereafter. Importantly, the sample mean of the inflation rate falls significantly between the two periods, as illustrated by the regressions of the inflation rate on a constant alone. The results verify that there is a substantial drop in the sample mean from a yearly average level of 16.9% to 7.2% after 1991, reflecting the fact that inflation convergence with other euro area members took place. By contrast, the volatility of CPI inflation remains almost unchanged. Inflation's standard deviation decreases from 5.5% in the first period to 5.1% in the second period.[6]

Overall, according to the autoregressive estimates using classical analysis, there appears to have been a sizeable shift in the mean of Greek inflation over the past twenty-five years, but only a small shift in inflation persistence. As we have noted, inflation persistence does not have the same notion in an inflationary period as in

[6] The same behaviour is exhibited by the WPI inflation series. There is a structural break in the intercept, while the serial correlation of the inflationary process remains unchanged. The sample mean of the inflation rate falls to 5.8% in the second period compared with 15.8% pre-1991, whereas there is no variation in persistence over time; it goes from 0.59 to 0.58 after 1991. Furthermore, the variance before and after the break decreases; WPI inflation is 1.4 times less volatile in disinflation relative to the period of high inflation. The results for core CPI inflation reveal a substantial rise in the serial correlation of Greek inflation, with the autoregressive coefficient going from 0.53 to 0.88, before and after 1991. The mean of core inflation is 2.5 times lower after the break, while inflation has approximately the same variance between the two periods.

disinflation. For example, the persistence of the 1970s and the 1980s is more of the type we would associate with negative effects - a shock occurred which raised inflation and inflation only very slowly returned to its previous low level. By contrast, the persistence of the 1990s results from steadily falling inflation, which can hardly have a negative connotation.

Table 1. Estimates of AR representation of CPI Inflation

High inflation *sample period: 1975 Q1-1990 Q4*

$$\underset{(23.444)}{\Delta p_t = 16.878} \quad \text{adj-}R^2=0.000,\ se=5.530,\ DW=1.167$$

$$\Delta p_t = \underset{(1.846)}{5.216} + \underset{(4.239)}{0.706}\,\Delta p_{t-1} + \sum \beta_j\, \Delta\Delta p_{t-j}\,,\ j=1\ldots k-1,\ k=4$$

adj-R^2=0.247, *se*=4.797, DW=1.938, LM(1)=2.043, LM(2)=2.237, ARCH(1)=2.076, ARCH(2)=3.238

Disinflation *sample period: 1991 Q1-2003 Q2*

$$\underset{(9.926)}{\Delta p_t = 7.212} \quad \text{adj-}R^2=0.000,\ se=5.138,\ DW=0.566$$

$$\Delta p_t = \underset{(0.989)}{0.655} + \underset{(11.165)}{0.781}\,\Delta p_{t-1} + \sum \beta_j\, \Delta\Delta p_{t-j}\,,\ j=1\ldots k-1,\ k=4$$

adj-R^2=0.757, *se*=2.530, DW=2.174, LM(1)=1.270, LM(2)=3.343,

ARCH(1)=0.299, ARCH(2)=3.718

Notes: annualised quarterly inflation in CPI. s*e* is the standard deviation of price changes, LM(1) and LM(2) are the Breusch-Godfrey (1978) L-statistic for the presence of first- and second-order residual autocorrelation, and ARCH(1) and ARCH(2) are Engle LM test for autoregressive conditional heteroskedasticity.

These results are in line with recent evidence for the euro area and other industrial countries. For example, Batini (2002) finds that there has been very little change in the serial correlation of HICP euro area inflation after 1970, while there is a substantial drop in the mean value of the euro area inflation rate. Gali, Gertler and Lopez-Salido (2000) present evidence on the fit of the New Keynesian Phillips Curve to euro area-wide data and conclude that inflation in the euro area may be less inertial than in the US. Smets and Wouters (2002) and Coenen and Wieland (2000, 2002) reach similar conclusions, whereas O'Reilly and Whelan (2004) detect some evidence of a decline in the autocorrelation coefficients in the US and less evidence of declining autocorrelations for the euro area in the 1990s. Levin and Piger (2004) present results that imply that high inflation persistence is not an inherent characteristic of industrial economies.[7]

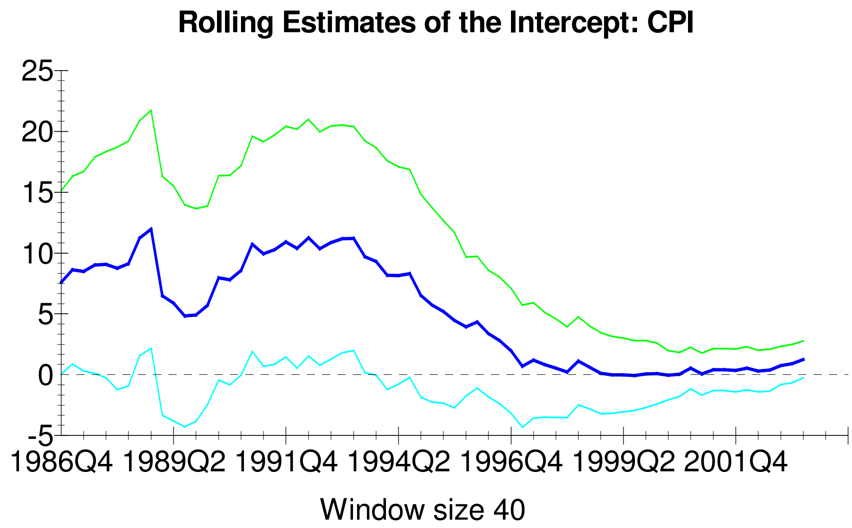
4.2 Rolling Regressions

In the preceding analysis, we allowed for a structural break in the regression coefficients by estimating regression (2) across two sub-periods following a Chow test and Bai and Perron multiple breaks tests for structural breaks. An alternative way of analysing the influence of structural breaks in the persistence parameters is to use rolling regressions to estimate the parameters of equation (2). We estimate the constant and the persistence parameter for every sub-sample of a fixed length (known as a data window) following Stock (2001), Pivetta and Reis (2001) and Levin and Piger (2004). By changing the length of the sub-sample, i.e. by shortening or extending the data window, we study whether the conclusions for the parameter estimates change considerably. We use alternative window lengths of 40, 48 and 56 quarters (10, 12 and 14 years).

Figures 2-4 present rolling estimates both of the persistence parameter and the intercept of equation (2). There is evidence of an increase in CPI inflation persistence to 0.8 in the early 1990s. However, with a rolling sample of 40 and 48 quarters, persistence starts roughly to decline by the end of the sample. The decline in inflation

[7] By contrast, empirical work for the US and the UK detects high inflation persistence, at least for the pre-1990 period. See, for example, Nelson and Plosser (1982), Fuhrer and Moore (1995), Gali and Gertler (1999), Taylor (2000), Cogley and Sargent (2001), Batini and Nelson (2001), Stock (2001) and Sbordone (2002).

Figure 2


Rolling Estimates of the Intercept: CPI
25
20
15
10
5
0
-5
1986Q4 1989Q2 1991Q4 1994Q2 1996Q4 1999Q2 2001Q4
Window size 40

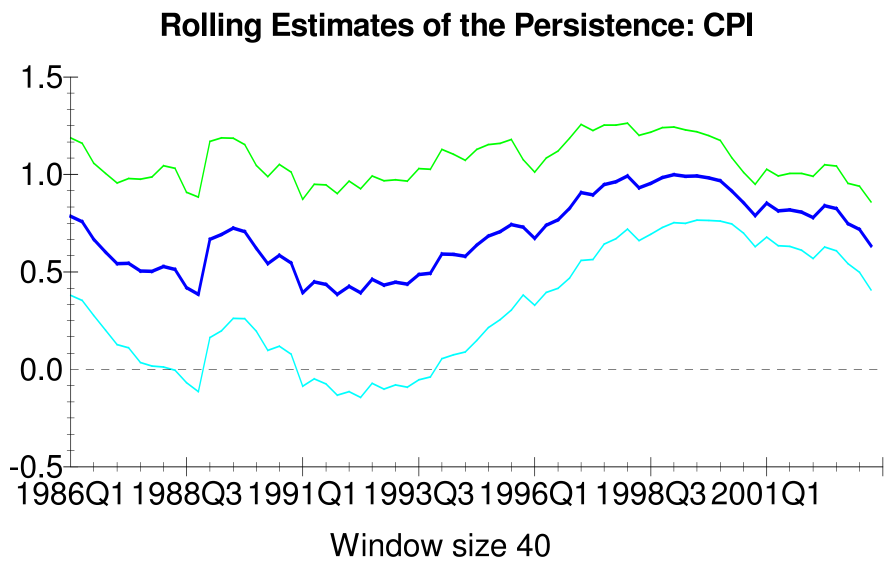
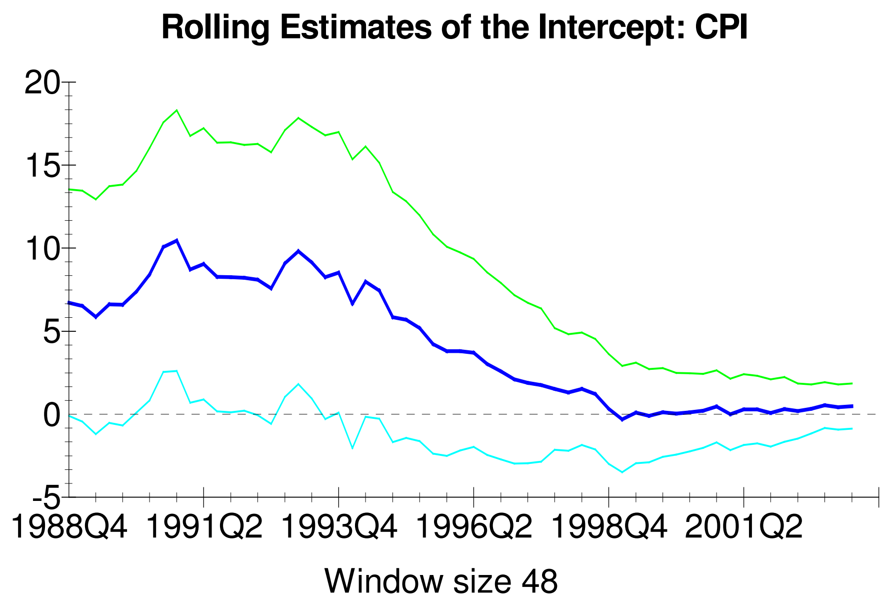

Rolling Estimates of the Persistence: CPI
1.5
1.0
0.5
0.0
-0.5
1986Q1 1988Q3 1991Q1 1993Q3 1996Q1 1998Q3 2001Q1
Window size 40

Figure 3


Rolling Estimates of the Intercept: CPI
20
15
10
5
0
-5
1988Q4
1991Q2
1993Q4
1996Q2
1998Q4
2001Q2
Window size 48

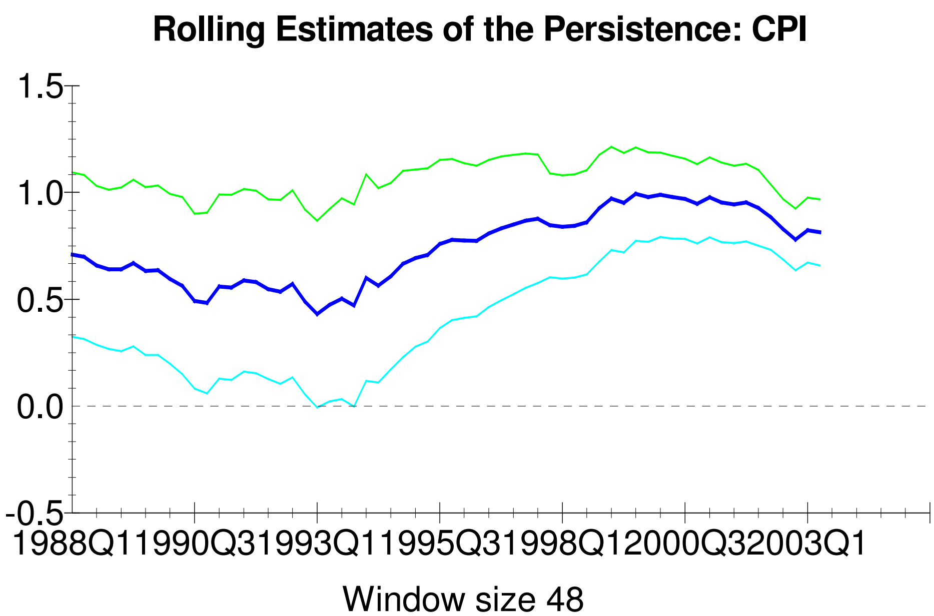
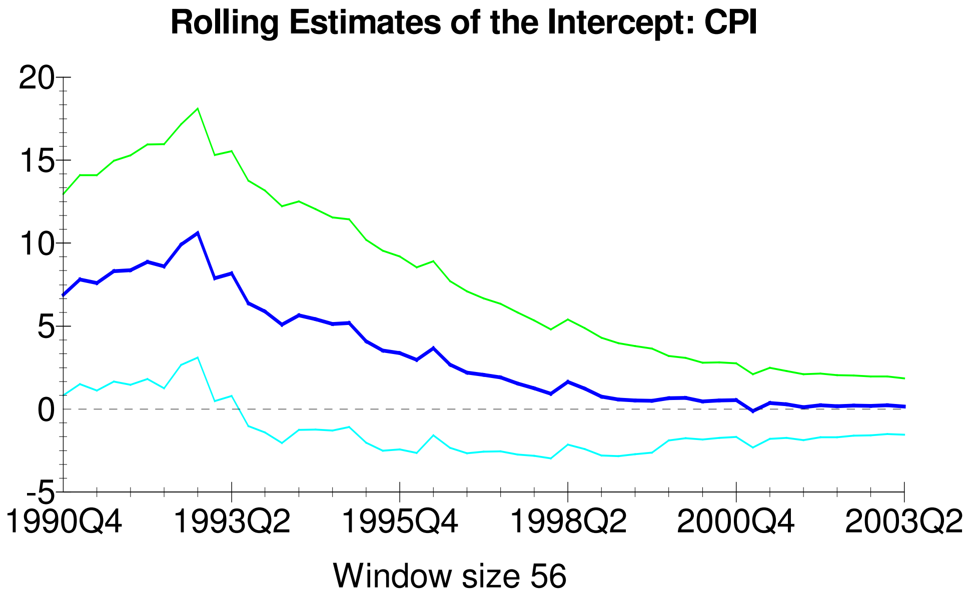
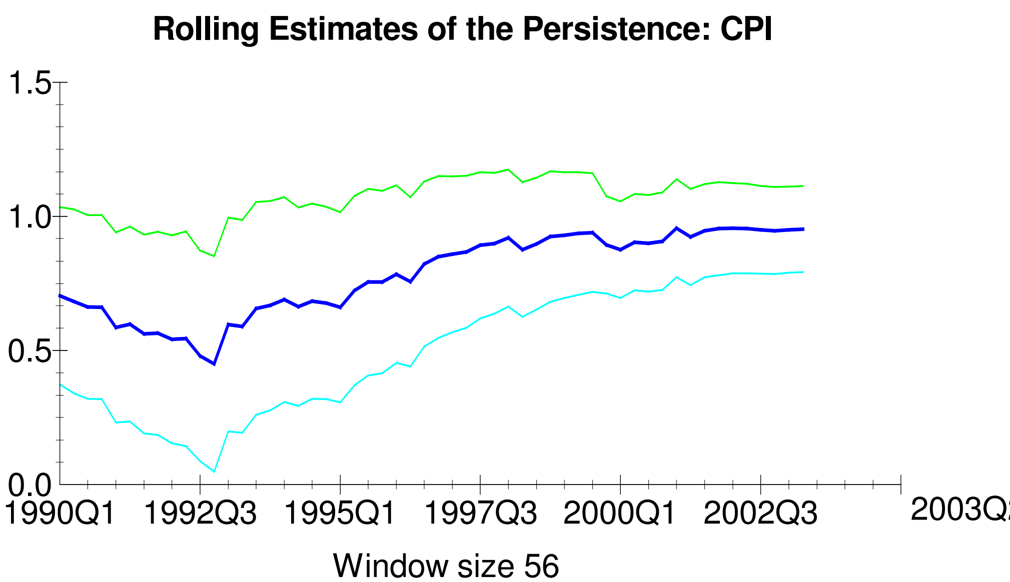

Rolling Estimates of the Persistence: CPI
1.5
1.0
0.5
0.0
-0.5
1988Q1
1990Q3
1993Q1
1995Q3
1998Q1
2000Q3
2003Q1
Window size 48

Figure 4



Rolling Estimates of the Intercept: CPI
20
15
10
5
0
-5
1990Q4
1993Q2
1995Q4
1998Q2
2000Q4
2003Q2
Window size 56



Rolling Estimates of the Persistence: CPI
1.5
1.0
0.5
0.0
1990Q1
1992Q3
1995Q1
1997Q3
2000Q1
2002Q3
Window size 56

persistence appears earlier when the window size is shortened to 40 quarters, while there is no evidence of a decrease when the data window size is extended to 56 quarters.[8] Concerning the constant, there is strong evidence, regardless of the data window size, that a substantial drop in the steady state inflation rate occurred in the early 1990s.

5. RC Estimates

5.1 RC Estimation Method

An alternative approach to measuring the degree of inflation persistence is based on RC methodology.[9] Consider the simple AR(k) process of the previous section, i.e. equation (1), which we reproduce for convenience here,

$$\Delta p_t = \text{constant} + \sum_{j=1}^{k} \alpha_j \Delta p_{t-j} + u_t \qquad (1)$$

where Δp_t is the annual quarterly inflation rate ($\Delta p_t = p_t - p_{t-4}$, p_t is the logarithm of the CPI), α_j is the autoregressive coefficient and u_t is a serially uncorrelated but possibly heteroskedastic random error term. Three lags are used in the estimation. The degree of persistence is equal to the sum of the autoregressive coefficients α_j.

In the RC estimation we employ the year-on-year inflation rate instead of the quarter-on-quarter rate. This choice is motivated by the underlying assumptions of RC methodology employed to eliminate the four restrictions, which are often imposed when classical estimation procedures are used. The estimated model, employing the RC estimation, can only coincide with the true underlying relationship if each of its estimated coefficients is interpreted as the sum of three terms: a "bias free effect", an "omitted variable bias" and a "measurement-error bias". The "bias free effect" is the estimated coefficient that would be observed if a) there were no measurement errors, b) there were no omitted variable biases and c) the specified functional form was true. The omitted-variables bias terms arise due to the fact that the true values of included explanatory variables affect those of excluded explanatory variables, which, in turn,

[8] EMU membership might explain this structural change appeared at the end of the sample. Levin and Piger (2004) run rolling regressions using different data window sizes for US GDP price deflator and also find that with a shorter window the drop in persistence appears earlier.

influence the true value of the dependent variable. The measurement-error bias effect is due to errors in the measurement of the included explanatory variables. From these interpretations it follows that equation (1) can coincide with the true inflation persistence equation only if the coefficients on its included explanatory variables are time-varying and not constant as in the classical model. The sum of the intercept of equation (1) and u_t is also time-varying, since it is equal to the sum of three time-varying terms. Therefore the quarter-on-quarter inflation rate is expected to provide an erratic time-varying persistence coefficient since according to the RC estimation procedure the estimated persistence coefficient constantly is changing at each point in time. In contrast, other methods such as the recursive OLS do not write off the past and hence the estimated persistence coefficient at each point of time is mainly affected by the past observations. Employing the year-on-year inflation rate which is a four quarter moving average of the quarter-on-quarter rate, the estimated time–varying persistence coefficient does not appear erratic.

Standard estimation procedures often impose a number of restrictions when applied to equations such as equation (1) above, including the following: (i) α_0, α_1, α_2, and α_3 are constant; (ii) excluded explanatory variables are proxied by an error term and, therefore, these excluded variables are assumed to have means equal to zero and to be mean independent of the included explanatory variables; (iii) the true functional form is known (whether linear or nonlinear); and (iv) the variables are not subject to measurement error.

Swamy and Tavlas (1995, 2001) and Chang, *et al.* (2000) define, first, any variable or value that is not mis-measured is true (definition I) and, second, any economic relationship with the correct functional form, without any omitted explanatory variable and without mis-measured variables is true (definition II). Using these definitions, we can specify a class of functions which is wide enough to cover the true inflation persistence function (in the sense of definition II) as a member. To rewrite this class in a form that has the same explanatory variables as equation (1), we assume that explanatory variables that are in the true inflation persistence function but excluded from equation (1) are related (linearly or nonlinearly) to the explanatory variables included in equation (1). This assumption is reasonable, given that economic

[9] For other recent applications of RC methodology on different topics, see Hondroyiannis, Swamy and Tavlas (2000, 2001a, b) and Brissimis, *et al.* (2003).

variables are rarely, if ever, uncorrelated and may not be linearly related to each other. To account for measurement errors, we assume that each variable in equation (1) is the sum of the underlying true value and the appropriate measurement error. These assumptions imply that equation (1) does not correspond to the true inflation persistence function unless it is changed to:

$$\Delta p_t = \gamma_{ot} + \gamma_{1t}\Delta p_{t-1} + \gamma_{2t}\Delta p_{t-2} + \gamma_{3t}\Delta p_{t-3},\ \text{t=1,2,..,T}, \tag{3}$$

where the real-world interpretations of the coefficients follow from the derivation of equation (3): γ_{0t} is the sum of three parts: (a) the intercept of the inflation persistence equation; (b) the joint effect on the true value of Δp_t of the portions of excluded variables remaining after the effects of the true values of the included explanatory variables have been removed; and (c) the measurement error in Δp_t. The coefficient γ_{1t} (γ_{2t} or γ_{3t}) is also the sum of three parts: (a) a bias-free effect of the true value of Δp_{t-1} (Δp_{t-2} or Δp_{t-3}) on the true value of Δp_t; (b) a term capturing omitted-variables bias; and (c) a mis-measurement effect due to mis-measuring Δp_{t-1} (Δp_{t-2} or Δp_{t-3}) (see Chang, *et al.* 2000). The bias-free effects provide economic explanations. An implication of these interpretations is that the explanatory variables of equation (3) are correlated with their coefficients. With these correlations, none of the explanatory variables is exogenous. The effects of such dynamic factors as technical change in economic activity and excluded lagged explanatory variables are captured in the omitted-variables bias component of each of the coefficients of equation (3). Consequently, equation (3) is a dynamic specification.

One question that needs to be answered before estimating equation (3) is that of parametrization: which features of equation (3) ought to be treated as constant parameters? Inconsistencies arise if this parametrization is not consistent with the real-world interpretations of γ's. To achieve consistency, the γ's are estimated using concomitants. A formal definition of concomitants is provided in Chang, *et al.* (2000) and Swamy and Tavlas (2001). Intuitively, these may be viewed as variables that are not included in the equation used to estimate inflation persistence, but help deal with the correlations between the γ's and the included explanatory variables (Δp_{t-1}, Δp_{t-2}, Δp_{t-3}).

<u>Assumption I.</u> The coefficients of equation (3) are linear functions of λ variables, called concomitants, including a constant term with added error terms

which may be contemporaneously and serially correlated. The error terms are mean independent of the concomitants.

Assumption II. The explanatory variables of equation (3) are independent of their coefficients' error terms, given any values of the concomitants.

Assumption II captures the idea that the explanatory variables of equation (3) can be independent of their coefficients conditional on the given values of concomitants even though they are not unconditionally independent of their coefficients. This property provides a useful procedure for consistently estimating the bias-free effects contained in the coefficients of equation (3).

Under Assumptions I and II, equation (3) can be written as

$$\Delta p_t = \pi_{00} z_{0t} + \sum_{j=1}^{\lambda-1} \pi_{0j} z_{jt} + \pi_{10} \Delta p_{t-1} z_{0t} + \sum_{j=1}^{\lambda-1} \pi_{1j} z_{jt} \Delta p_{t-1} + \pi_{20} \Delta p_{t-2} z_{0t} + \sum_{j=1}^{\lambda-1} \pi_{2j} z_{jt} \Delta p_{t-2} + \pi_{30} \Delta p_{t-3} z_{0t} + \sum_{j=1}^{\lambda-1} \pi_{3j} z_{jt} \Delta p_{t-3} + \varepsilon_{0t} + \varepsilon_{1t} \Delta p_{t-1} + \varepsilon_{2t} \Delta p_{t-2} + \varepsilon_{3t} \Delta p_{t-3} \tag{4}$$

where the z's denote concomitants and the ε's denote the error terms of the coefficients of equation (3).[10] In our empirical work we set λ=4, z_{0t}=1 for all t, z_{1t} is the short-term bank lending rate, z_{2t} is the annual growth rate of nominal GDP and z_{3t} is the quarterly change in nominal wages.[11] This means that we use four concomitants to estimate the γ's. We are attempting to capture the bias-free effect contained in γ_{1t} by using a linear function $(\pi_{10} + \pi_{11} z_{1t} + \pi_{12} z_{2t})$ of the short term interest rate and annual change of nominal GDP.[12] The biased and mis-measurement effects are captured by using a function $(\pi_{13} z_{3t} + \varepsilon_{1t})$ of the quarterly change of nominal wages and ε_{1t}. The measures of bias-free effects contained in γ_{2t} and γ_{3t} are $\pi_{20} + \pi_{21} z_{1t} + \pi_{22} z_{2t}$ and $\pi_{30} + \pi_{31} z_{1t} + \pi_{32} z_{2t}$, respectively, and those of omitted-variable and mis-measurement effects contained in γ_{2t} and γ_{3t} are $\pi_{23} z_{3t} + \varepsilon_{2t}$ and $\pi_{33} z_{3t} + \varepsilon_{3t}$,

[10] Adding concomitants successively should reduce the RMSE of the estimated regressions.

[11] Economic theory provides us with many categories of variables that may, indirectly or directly, influence the inflation process. In particular, interest rates reflect monetary conditions in the economy. Assuming a constant mark-up, wages might be expected to be transmitted to output prices and hence CPI inflation. Output is considered as a measure of demand pressure and thus affects inflation expectations.

[12] Growth rate of nominal GDP and short-term interest rate are broader measures compared to growth rates of wages. Hence, the first two concomitants may give a better measure of the bias-free effect than the latter.

respectively. The components of the coefficients of equation (3) can take different values in different phases of the business cycle. Inflation persistence is estimated as the summation of γ_{1t}, γ_{2t} and γ_{3t} to estimate the total effect (biased effect) and the summation of $(\pi_{10} + \pi_{11} z_{1t} + \pi_{12} z_{2t})$, $(\pi_{20} + \pi_{21} z_{1t} + \pi_{22} z_{2t})$ and $(\pi_{30} + \pi_{31} z_{1t} + \pi_{32} z_{2t})$ to estimate the bias-free effect. Inflation persistence may be lower in periods of low inflation compared to periods of high inflation. Consequently, changes in the values of the included explanatory variables that occur during the peak of a business cycle may exhibit very different effects on persistence than the same changes that occur during the trough of a business cycle. If so, more accurate results can be obtained by taking changing conditions into account.[13]

Note that equation (4) has four error terms, three of which are the products of ε's and the included explanatory variables of equation (1). The sum of these four terms is both heteroskedastic and serially correlated. Under Assumptions I and II, the right-hand side of equation (4) with the last four terms suppressed gives the conditional expectation of the left-hand side variable as a nonlinear function of the conditioning variables. This conditional expectation is different from the right-hand side of equation (1) with u_t suppressed. This result shows that the addition of a single error term to a equation and the exclusion of the interaction terms on the right-hand side of equation (4) introduce inconsistencies in the usual situations where measurement errors and omitted-variable biases are present and the true functional forms are unknown. A computer program developed by Chang, *et al.* (2000) is used to estimate equation (4).

It might be useful at this point to summarize intuitively the above theoretical discussion. Empirical models often contain problems, such as omitted variables and measurement errors, introducing biases in the estimated coefficients. Each of the estimated coefficients is comprised of three terms: a bias-free coefficient, an omitted variables-bias term and a measurement error. The RC methodology overcomes this problem by introducing two assumptions. According to the first assumption, concomitants are employed in the empirical estimation to explain the variations in each of the three terms of the estimated coefficient. In particular, the set of included concomitants is separated into two subsets. The first subset is used to estimate the

[13] A richer specification of concomitants might provide different results.

bias-free effect and the second one to estimate the combined omitted-variables-bias and the measurement-error-bias term. The second assumption is used to assure the independence between the explanatory variables and the remainders of the estimated coefficients, after removing the effects of the concomitants on the coefficients.

5.2 RC Empirical Results

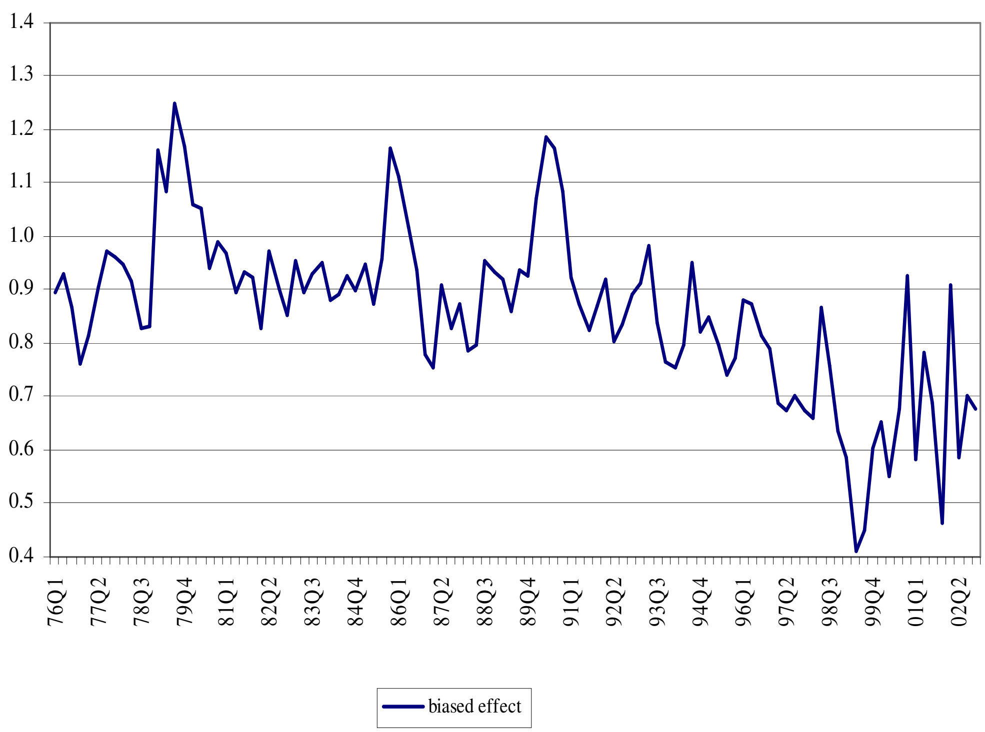
Figure 5 presents the results from RC estimation without concomitants. The estimated RMSE[14] is equal to 0.772. Next the RC model is estimated employing initially one concomitant that is the short-term bank lending rate (RMSE=0.646) and two concomitants that is the short-term bank lending rate and the annual growth rate of nominal GDP (RMSE=0.548). Finally, the RC model is estimated using three concomitants. Figures 6.a and 6.b present the results from RC estimation, for the intercept and the degree of persistence employing three concomitants: the short-term lending rate, the annual growth of nominal GDP and the quarterly change of nominal wage (RMSE=0.301).

In the post-1975 period, there appears to have been a sizeable shift in inflation persistence as well as in the mean value of inflation. CPI inflation increased on a yearly basis from 13.2% in the first quarter of 1976 to 21.0% in the last quarter of 1990. The RC results also show a sharp rise in the degree of inflation persistence, from 0.6 to 0.9 during this period. The behaviour of inflation changed after 1990. The mean value dropped from a double-digit level (16.9%) to a single-digit one (8.0%). The degree of inflation persistence declined from 0.84 in 1990 Q4 to 0.57 in 2002 Q4 (bias-free effect). Even though there is a considerable shift in the mean value in the post-1990 period and the degree of inflation persistence started to decline in the early 1990s, persistence is still quite high and increases when inflation is high.

Both total and bias-free effects show a tendency for the degree of inflation persistence to decline during the disinflation period. In particular, during the high inflation period (1976 Q1-1990 Q4) the total effect estimates indicate that the average degree of inflation persistence is equal to 0.75, while it decreases to 0.67 during the disinflation period. By contrast, the bias-free effect estimates indicate an increase in the average degree of inflation persistence from 0.70 during the inflation period to

0.76 during the disinflation period. This latter result is qualitatively the same as the result from OLS estimation for the two different periods.[15]

Figure 5
RC Estimation of Inflation Persistence in Greece Without Concomitants



ROLS and recursive OLS provide much higher estimates, which are due to omitted-variables and measurement-error biases. ROLS estimation for different size windows is a way of testing for stability if the structure is changing. However, ROLS estimation does not give consistent estimates of the changing parameters. Unlike RC estimation, an underlying assumption of rolling and recursive estimation is that the coefficients are constant. Also, recursive estimation does not write off the past. Rolling estimation writes off the past slowly as the estimation period changes. If a

[14] The estimation period for all RC models is 1976 Q1-2001 Q4 and the forecast period is 2001 Q1-2002 Q4. The estimation period for Figures 5 and 6 is 1976 Q1-2002 Q4.

[15] A true comparison of the empirical findings from the RC estimation with those of OLS requires us to re-estimate equation (1) using year-on-year inflation. The estimated inflation persistence is equal to 0.98 for the period 1976 Q1 to 2002 Q4, 0.86 for the period 1976 Q1 to 1990 Q4 and 0.94 for the period 1991 Q1 to 2002 Q4. As expected, the estimates of persistence are higher using year-on-year data since the latter is a 4 quarter moving average of the quarter-on-quarter rate which in itself introduces some persistence.

regime change has occurred, it averages the old regime with the new regime with changing weights; the weight of the new regime becomes larger as more and more observations are added. Using recursive estimation, the best estimate of a model is that produced by the entire sample. In contrast, RC estimation picks up a new regime quickly. By consistently estimating changing coefficients, it attempts to estimate how the true coefficient is changing at each point in time. Therefore, the RC estimates are a better measure of the true degree of persistence implying the true response of inflation to a shock at each point in time.

The total effect (biased effect) in the RC estimation is not purged of omitted-variables and measurement-error biases. The bias-free effect displays much less volatility indicating that a big impact of specification errors on the time profiles of the estimated coefficients of equation (1) is present. The bias-free effect clearly shows the increase in the degree of inflation persistence during the inflation period and the moderate decrease during the disinflation period. The bias-free effect in Figure 6 shows that the degree of inflation persistence remained high (average=0.84) during the first six years of the disinflation period (1991 Q1-1996 Q4) and started to decline (average=0.67) after 1997 Q1.

Figure 6.a
Inflation Persistence in Greece

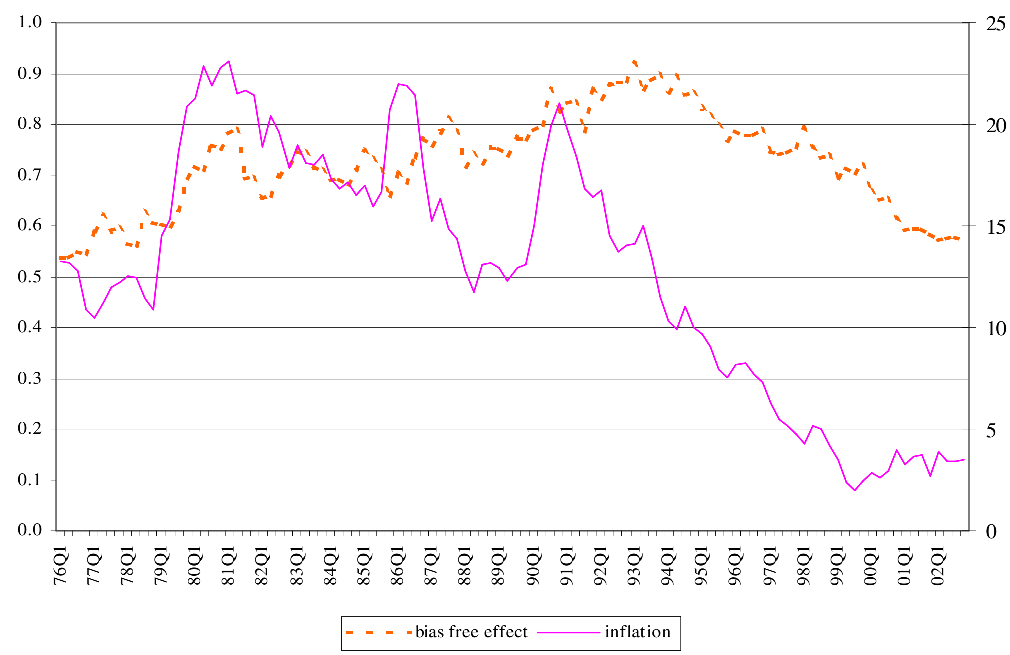
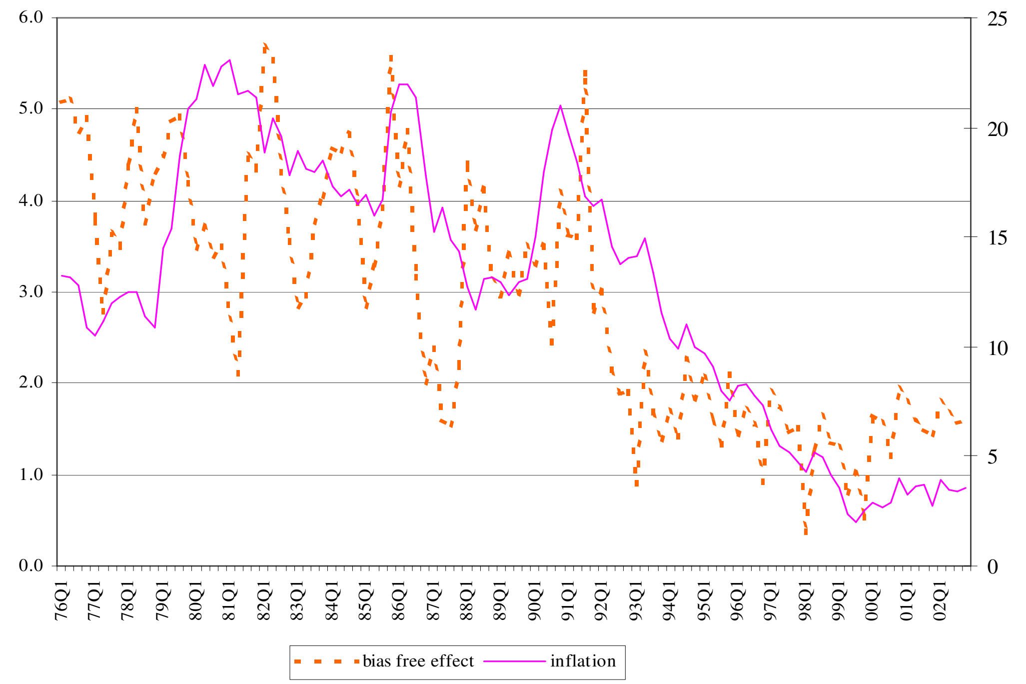


Figure 6.b
Intercept Time Profile (RC estimates)



6. Delayed Inflation Response and Monetary Policy Regime Switches

Having studied the serial correlation properties of Greek inflation, we are now interested in identifying whether inflation responds with a delay to monetary policy moves and whether this observed delay is dependent on changes in the monetary policy regime. In an earlier paper, Gibson and Lazaretou (2001), looking at Greek monthly data over the period 1958-1998, find that money (M0) growth leads CPI inflation at turning points by about 4 months. In this section, we distinguish between a change in monetary stance (systematic monetary policy action) and a monetary policy shock (non-systematic) and we try to determine whether and how much the response of inflation (i.e. the number of periods it takes for a change in monetary policy to have its peak effect on inflation) has changed following changes in the monetary policy framework.

6.1 Systematic Policy Actions

The response of inflation (as well as of output) to changes in monetary policy conditions determines the cost of disinflation and stabilisation. Knowledge of this is therefore essential for the monetary policymaking. Friedman (1972) and Friedman and Schwartz (1982) look at monthly data for the US and the UK and found that monetary changes affected output with a short lag. The lagged effect on prices was found to be longer. Batini and Nelson (2001) extend Friedman's evidence to cover the most recent period and find that it takes over a year before monetary actions have their maximum effect on inflation. More recently, Batini (2002) using the data for the euro area reaffirms Friedman's results.

A simple statistic of the relation between inflation and prior monetary changes is the correlation coefficient ρ(k) of quarterly changes of the CPI with money growth, k periods earlier, where k corresponds to the lead with the highest correlation. As a monetary policy stance measure, we consider quarterly changes in the broad money aggregate (M3). We also compute correlations between inflation and alternative monetary policy stance measures, such as narrow money growth (M1), the short-term *ex ante* real interest rate[16] and the yield spread.[17]

Table 2 presents the results for different indicators of systematic monetary policy changes across separate sub-samples. The sub-periods are selected according to the major monetary events occurring in the whole period which might distort the steady-state velocity of money. Thus, the complete sample period 1980 Q1-2003 Q1 is divided into the following sub-samples: 1980 Q1-1987 Q4 (rapidly sliding depreciation of the drachma); 1988 Q1-1994 Q4 (limited sliding depreciation of the drachma); 1995 Q1-2000 Q4 (hard drachma policy; pre-ERM and pre-EMU period); 2001 Q1-2003 Q1 (EMU).

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[16] The short-term real interest rate is computed as the difference between the nominal interest rate and the *ex ante* inflation rate. *Ex ante* inflation is approximated by the fitted values of an AR(4) scheme for CPI inflation. A H-P trend of the log of the real GDP is also included in the explanatory variables. We use the short-term bank lending rate (quarterly averages) and inflation is the quarter-on-quarter change of CPI.

[17] The yield spread is the spread of the long-term bond rate over the short-term interest rate. As the long-term interest rate we use the 12-month Treasury bill rate. Fixed income government securities at 5 years and over were issued for the first time in 1997. Initially, Greek markets were thin; it took some time to develop. Therefore, the 12-month Treasury bill rate is used as a proxy for the long-term interest rate.

Table 2. Correlations Between CPI Inflation and Systematic Monetary Policy Actions (maximun positive value)

Sample Period	Policy Stance Measure: M3 growth rate $\rho_{\Delta p}{}^{CPI}{}_{m3}$ (k)	Policy Stance Measure: M1 growth rate $\rho_{\Delta p}{}^{CPI}{}_{m1}$ (k)	Policy Stance Measure: *ex ante* real interest rate $\rho_{\Delta p}{}^{CPI}{}_{r}$ (k)	Policy Stance Measure: yield spread $\rho_{\Delta p}{}^{CPI}{}_{spread}$ (k)
1980 Q1-2003 Q1	0.717 (k=1) (9.589)	0.381 (k=3) (3.844)	-0.708 (k=18) (-9.561)	-0.431 (k=17) (-4.556)
1980 Q1- 1987 Q4	0.477 (k=3) (2.761)	0.467 (k=4) (2.693)	-0.557 (k=1) (-3.674)	-0.302 (k=16) (-1.734)*
1988 Q1-1994 Q4	0.597 (k=10) (3.791)	0.355 (k=1) (1.936)	-0.452 (k=6) (-2.584)	-0.597 (k=8) (-3.795)
1995 Q1-2000 Q4	0.459 (k=7) (2.423)	0.250 (k=2) (1.211)	-0.456 (k=1) (-2.403)	-0.439 (k=11) (-2.292)
2001 Q1-2003 Q1	*a*	*a*	*a*	*a*

Notes: (*) significant at 10% statistical level. Parentheses below the value of the statistic present the value of the t-test. *a.* Due to the small size of the sample, the correlation cannot be computed.

The expected relationship between money growth and inflation is positive. The results for the complete period suggest a very short lead of money growth of one quarter for M3 and three quarters for M1 over inflation. Excluding the 1980s, the lead time increases for M3 money growth at k=10 and k=7 for the periods of the policy of limited sliding depreciation of the drachma and the pre-ERM period. Looking at M1 growth, the correlations are smaller in size and statistically insignificant. For the period as a whole, the evidence from the interest-rate-based indicator of monetary policy stance reports a delay in the reaction of inflation to changes in policy. The correlation is negative (as expected) and significant; the peak response occurring after around 4 years. We check the robustness of this result using the yield spread as an alternative interest-rate-based-measure (see Gali, Gertler and Lopez-Salido, 2002 and Batini, 2002). As can be seen from the last column of Table 2, the correlation is (correctly) negative and statistically significant, although smaller in size, and the lead-lag relationship appears to last 17 quarters. Concerning the cross-period correlations, the lagged response of inflation to a change in the *ex ante* real interest rate, although significant, is much shorter, whereas the results display a longish lead of the yield spread over inflation.[18]

To sum up, the results reveal that there exists a lagged response of inflation to monetary policy moves. Findings are relatively consistent across different measures of monetary stance, indicating that Greek inflation is rather inertial. However, we detect no major changes in the lead-lag relationships across sub-periods.

6.2 Policy Shocks

An alternative measure of inflation persistence is the number of lags it takes for inflation to have its maximum response to a monetary policy shock.[19] We estimate an unconstrained vector autoregressive model for the Greek economy with three endogenous variables: Δp_t, the quarterly change of CPI, y^{dev}, the deviation of the log of the real GDP from its H-P trend and i_t, the short-term nominal interest rate. Statistics that test the lag length result in the following model specification[20]

$$\mathbf{Y}_t = \mathbf{A} + \sum_{j=1}^{k} \mathbf{B}_j \mathbf{Y}_{t-j} + \mathbf{U}_t \qquad t=1,\ldots, T, \qquad k=2 \tag{5}$$

where $\mathbf{Y}_t$ is a (m×1) vector of jointly determined endogenous variables, $\mathbf{B}_j$ is the (m×m) coefficient matrix and $\mathbf{U}_t$ is a (m×1) vector of serially uncorrelated random errors.[21] The VAR model is estimated using quarterly data for the two sub-periods. Over the entire sample period, there have been several major shifts in the country's monetary policy regime. By estimating the VAR model over the complete period, there is the danger that the parameter estimates will be inconsistent due to these policy shifts. Thus, we choose two specific sub-samples that correspond to two particular

[18] Correlations between the measures of systematic monetary policy and inflation are also computed using core CPI inflation and WPI inflation. The results confirm the evidence for CPI inflation.

[19] See, for example, Christiano, Eichenbaum and Evans (2001) and Rotemberg and Woodford (1997) for the US economy, and Smets and Wouters (2002) and Fagan, Henry and Mestre (2001) for the euro area.

[20] We used the Akaike Information Criterion and the Schwartz Criterion. Both information criteria indicate a two-lag VAR.

[21] The null hypothesis of a unit root in the variables of the model cannot be accepted at 5% or 10% statistical level of significance: Δp_t and y^{dev} are I(0), while the interest rate seems to be I(1). Thus, we apply the standard analysis and not cointegration analysis in estimating the VAR model. Besides, our main interest is to detect short-run interactions between the variables in the system and not to identify

regime periods. The 1980s were characterised by monetary expansion, a sliding drachma policy and high inflation rates. Beginning in the early-1990s, macroeconomic policy in Greece primarily focused on disinflation. As we have already mentioned, the Bank of Greece announced a specific exchange rate target in order to bring inflation down to single digit levels.

To measure inflation persistence, we need to quantify the number of time periods between the timing of the monetary policy shock and the maximum response of inflation. To do this, we need to estimate an impulse response function that traces the effect of a one-time shock to one of the innovations on current and future values of the endogenous variables. Monetary policy shocks are identified as innovations in the equation for the short term interest rate.[22] Based on conventional economic theory, the interest rate responds to the monetary policy shock and affects output and inflation.

Figure 7 shows the response of CPI inflation to a one standard deviation of shock to the innovation of the nominal interest rate by means of generalized impulses proposed by Pesaran and Shin (1998).[23] As can be seen, the behaviour of the inflation rate exhibits the pattern of the so-called ‘price puzzle’ (Sims 1992), which is a common finding for impulse responses of estimated unconstrained vector autoregressions. In other words, the sign of the impulse response of inflation to a positive monetary policy shock is positive (i.e inflation increases after a contractionary monetary shock) rather than negative, as expected from the theory.[24]

As is evident from Figure 7, for the first sub-period the response of the CPI inflation rate to a generalised one standard deviation shock to the innovation on the short-term lending rate is negative for the first two quarters. Then, it becomes positive (indicating the existence of the price puzzle) and it peaks after 14 quarters. For the

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long-run relationships. Moreover, another constraint for cointegration analysis arises from the small number of observations in each sub-sample.

[22] We consider the short-term bank lending rate to enterprises (quarterly averages).

[23] A generalised decomposition constructs an orthogonal set of innovations that does not depend on the ordering of the variables of the VAR.

[24] Brissimis, *et al.* (2001) also find a price puzzle for the Greek economy when they use a VAR model to capture the elements of the transmission mechanism of monetary policy. One interpretation of the existence of the price puzzle proposed in the literature is that central banks have access to superior information compared to the private sector and, thus, they have better forecasts of expected inflation. On the basis of these forecasts, they raise their interest rates although to a lesser extent than required to completely offset inflationary pressures. Stiglitz (1992) offers another explanation of the perverse response of inflation. In an imperfectly competitive environment firms have an incentive to raise prices so as to increase their cash flows before economic activity declines after a monetary tightening.

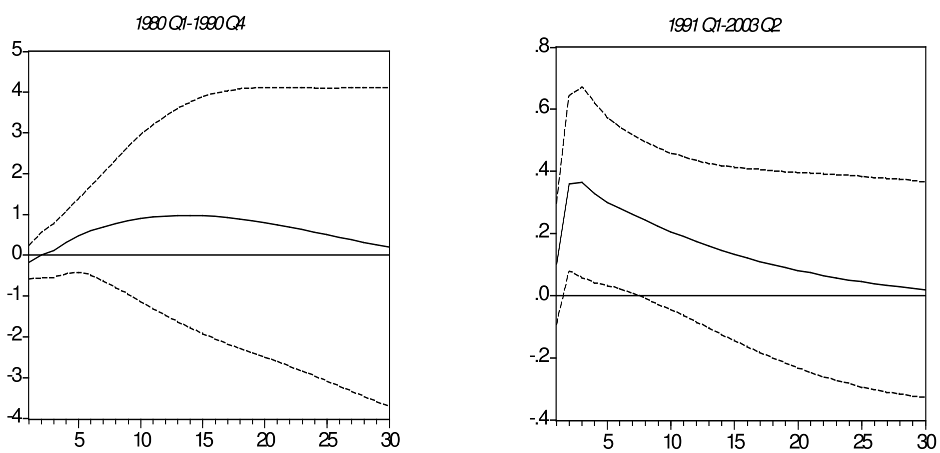
disinflation period after 1991, the inflation rate displays a similar perverse reaction to a contractionary monetary policy shock. However, it exhibits a quicker response (incorrectly positive; inflation increases immediately after the monetary shock and it peaks in the third quarter).

Similar results are found when we use the 12-month Treasury bill rate as a monetary policy instrument (Figure 8). For the high inflation period, the response has the correct sign (negative) and it leads 10 quarters. However, for the disinflation period, the inflation response exhibits a price puzzle.[25]

The picture is quite different when we consider the reserve money equation as a monetary policy reaction function. As can be seen in Figure 9, the sign of the impulse response in inflation to a monetary policy shock is negative, as expected, and indicates a 8-quarter lag between the shock and inflation peak effect post-1991. Pre-1991, it indicates a 3-quarter lag. Then, however, it becomes positive.

Figure 7

Generalised Impulse Response of CPI Inflation to One Standard Deviation Shock in the Equation for Short-term Lending Rate



[25] The inclusion of M3 money growth rate and a time trend in the set of the exogenous variables in our basic VAR model does not seem to help considerably with the price puzzle.

Figure 8

Generalised Impulse Response of CPI Inflation to One Standard Deviation Shock in the Equation for 12-month Treasury Bill Rate

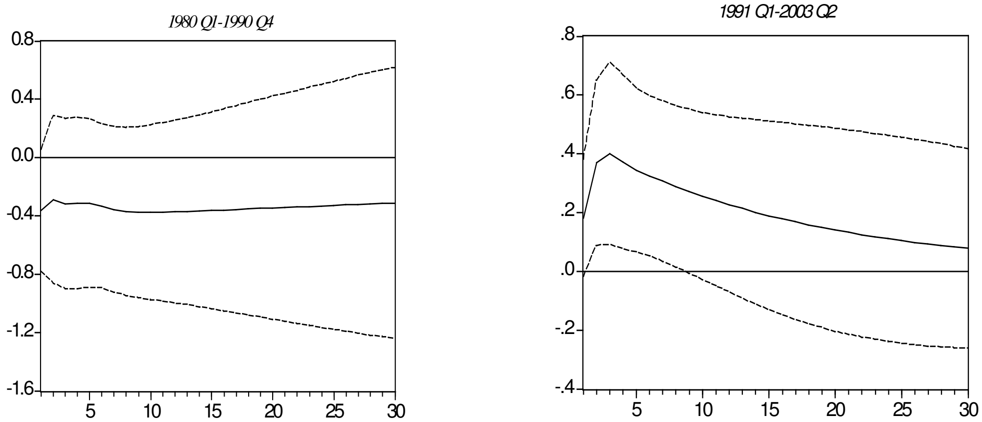
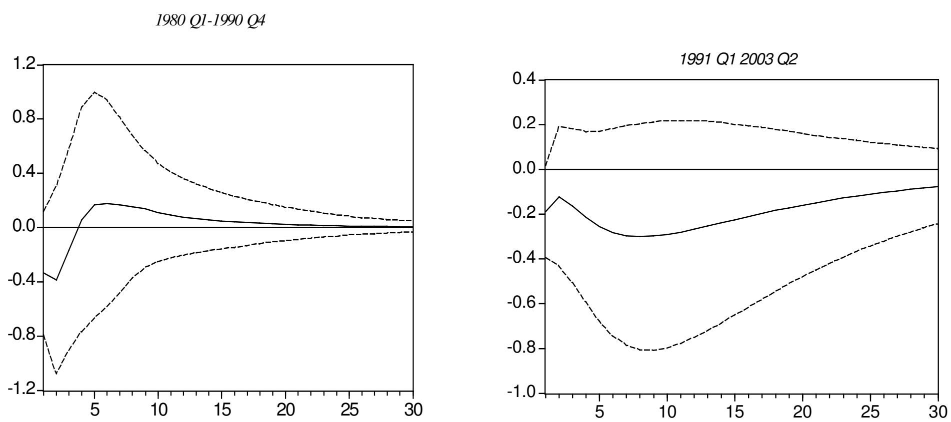


Figure 9

Generalised Impulse Response of CPI Inflation to One Standard Deviation Shock in the Equation for Reserve Money



7. Concluding Remarks

This paper has presented a model-free assessment of the time profile of inflation persistence for the post-1975 Greek economy. We have adopted three alternative definitions of inflation persistence proposed in the empirical literature (see Batini, 2002), i.e. (1) by looking at the autocorrelation properties of the inflation series, (2) by quantifying the lags of the inflation response to systematic monetary policy, and (3) by quantifying the lags of the inflation response to monetary policy shocks.

The first definition concerns the empirical analysis of inflation dynamics. It asks whether inflation inertia may be an intrinsic structural phenomenon of the economy or a policy regime-variant feature of the inflation time series. The second definition refers to the cost of disinflation. For monetary policymaking, knowledge of the exact extent of transmission lags from monetary policy to the inflation rate is crucial. Since a priority of monetary policy in Greece, as a member of the euro area, is to pursue the objective of price stability, the ability to quantify and model the lagged response of inflation to changes in monetary stance, helps policy makers to determine the cost of reducing inflation in terms of output loss. Finally, the third definition of inflation persistence accounts for the delays in the response of inflation persistence to monetary policy shocks.

However, some reservations arose. A model could not capture the extent of all kinds of inflation persistence. For example, a model might account for the sluggish response accompanied with systematic changes in monetary policy but it could fail to account for the delays in effect of the non-systematic component of monetary policy. Besides, policy shocks are only a minor determinant of inflation persistence compared to real shocks, such as shocks to technology. (For these reservations, see Batini, 2002 and Woodford, 1998).

Based on these alternative definitions of persistence, in this paper we have asked, first, whether the positive serial correlation of Greek CPI inflation has changed over time, second, whether there exists a sluggish response of inflation to systematic and non-systematic changes in monetary policy and, third, whether this observed delay has altered across changes in the monetary policy framework. To capture possible shifts in inflation persistence, we utilised different empirical methodologies, i.e. autoregressive univariate modelling, random coefficients modelling and vector autoregressive modelling for Greek inflation using macro data. We found that Greek inflation is rather persistent according to all definitions used and methodologies employed.

Empirical findings can be summarised as follows. First, classical estimations show that CPI inflation persistence was high during the inflationary period and the first six years of the disinflationary period (the persistence parameter took the value of 0.70 and approximately 0.78). Only after 1997, did it start to decline. Second, there appears to have been a sizeable shift in the mean of Greek inflation over the past twenty-five years, but only a small shift in inflation persistence. Third, rolling regressions were applied as an alternative way of analysing the influence of structural breaks in the coefficients of the AR process. We found that the sample mean of the inflation rate dropped dramatically in the 1990s, while the persistence parameter started to decline only by the end of the sample. The decline appeared earlier when the window size was smaller. Fourth, RC estimation (employing the year-on-year inflation rate rather than the quarter-on-quarter) produces slightly lower estimated values for the persistence parameter. Fifth, the estimated correlations between CPI inflation and policy stance measures as well as the results from a basic VAR model reveal that there was a delay in the inflation response to changes in monetary policy under the period of consideration. This delay seems to have changed little over time, as shown by the results obtained by dividing the sample to cover different sub-periods.

Our analysis has used macro data. However, an interesting topic for further research is to present evidence on the degree of inflation persistence looking at price data on a sectoral level rather on an aggregate level, and explain whether cross-section variance in inflation persistence and/or in the level of the inflation rate can be determined by the structural features of different sectors.

Appendix: Data Sources and Definitions

CPI, Consumer Price Index (1999=100), sample period 1975 Q1-2003 Q2, quarterly averages seasonally adjusted. Source: National Statistical Service of Greece.

Short-term bank lending rate to enterprises (quarterly averages), sample period 1975 Q1-2003 Q1, in percent. Source: Bank of Greece.

12-month Treasury Bill rate, sample period 1975 Q1-2003 Q2, (quarterly averages) in percent. Source: Bank of Greece.

Gross Domestic Product, market prices, at 1995 constant prices, in million euro (ESA 95), quarterly data, no seasonal adjustment, sample period 1975 Q1-2003 Q1. Source: National Statistical Service of Greece.

Gross Domestic Product, market prices, at current prices, in million euro (ESA 95), quarterly data, no seasonal adjustment, sample period 1975 Q1-2002 Q4. Source: National Statistical Service of Greece.

M3 money, M1 money, ECB definition, in million euro, no seasonal adjustment, end-of quarter data, sample period 1980 Q1-2003 Q1. Source: Bank of Greece.

Reserve money, sample period 1980 Q1-2003 Q2, in million drachmas, end-of quarter data, no seasonal adjustment. Source: Bank of Greece.

Nominal wages, minimum monthly salary of white colour workers in manufacturing, in thousands of drachmas, quarterly data, no seasonal adjustment, sample period 1975 Q1-2002 Q4. Source: National General Collective Agreements.

Annex Box
A Brief Chronology of the Major Monetary Policy Changes in Greece, 1975-2003

Year	*Monetary Policy Changes*	Comments
1975	Policy of rapid sliding depreciation of the drachma. Inflation rose following an accommodating monetary policy stance. Central bank financing of fiscal deficits.	Managed float. The drachma's uncoupling from the dollar and its float against a basket of currencies amounted to a policy of rapid sliding depreciation, in order to counterbalance the effects of wage increases on the country's international competitiveness.
1982	Abolition of the Currency Committee. Its responsibilities were transferred to the Bank of Greece.	The Committee, comprised of five ministers and the Governor of the Bank of Greece, had decided on monetary and credit policies and targets since 1946.
1983	Drachma's devaluation.	Accommodative macroeconomic policies. A crisis in the balance of payments, stagflationary developments.
1985	Drachma's devaluation in the context of a two-year stabilisation programme.	Switch to an anti-inflationary policy. The programme led to a narrowing of the current account deficit and to a fall in inflation.
1987	Financial liberalization and deregulation.	Deregulation measures improved the functioning of financial markets. The process was gradual and was completed in 1994.
1988	Abandonment of the stabilisation programme.	Relaxation of wage controls, new inflation crisis, fiscal deficits.
1991-94	Attention focused on disinflation. Efforts to meet the Maastricht criteria.	A strict incomes policy, tightened monetary and fiscal policy stance. Inflation fell to a more moderate level, sluggish economic growth.
1995	Hard drachma policy. The attainment of price stability became the primary objective of the monetary policy. To this end, the Bank of Greece relied on two intermediate targets; an exchange rate target and a target range for M3 growth.	The Bank of Greece adopted, for the first time, a nominal exchange rate anchor as the main intermediate target for monetary policy. By limiting the year-on-year depreciation of the drachma against the ECU to a rate that did not fully accommodate the inflation differential between Greece and the EU average, the exchange rate anchor was used as an important mechanism for disinflation.
1997	Central Bank Independence.	The Bank of Greece was given by law control over exchange rate policy and was

		provided with a mandate to attain monetary and price stability.
1998	ERM entry policy.	On 16 March the drachma joined the European Exchange Rate Mechanism (ERM) at a central rate that implied 12.3% devaluation against the ECU. Budgetary measures, a strict incomes policy and the introduction of structural reforms, mainly in the financial sector supported the new exchange rate.
1999	Entry into ERM II	On 1 January the drachma began to participate to ERM II (at a central rate of 353.109 per ECU with a standard fluctuation band of ±15%).
2001	Greece entered the euro area.	On 17 January 2000 the drachma's central rate against the euro was revalued by 3.5% to 340.750 drachmas per euro. On 19 June the ECOFIN Council admitted Greece into the euro area, effective 1 January 2001.

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