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INFLATION AND INFLATION UNCERTAINTY IN THE EURO AREA

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Abstract

This paper estimates a time-varying AR-GARCH model of inflation producing measures of inflation uncertainty for the euro area, and investigates the linkages between them in a VAR framework, also allowing for the possible impact of the policy regime change associated with the start of EMU in 1999. The main findings are as follows. Steady-state inflation and inflation uncertainty have declined steadily since the inception of EMU, whilst short-run uncertainty has increased, mainly owing to exogenous shocks. A sequential dummy procedure provides further evidence of a structural break coinciding with the introduction of the euro and resulting in lower long-run uncertainty. It also appears that the direction of causality has been reversed, and that in the euro period the Friedman-Ball link is empirically supported, implying that the ECB can achieve lower inflation uncertainty by lowering the inflation rate.

JEL Classification: E31, E52, C22

Keywords: Inflation, Inflation Uncertainty, Time-Varying Parameters, GARCH Models, ECB, EMU

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1. Introduction

Inflation uncertainty, its linkages with actual inflation and its potential impact on real economic activity have been extensively analysed in the literature. Friedman (1977) was the first to suggest that higher average inflation could result in higher inflation uncertainty. This idea was developed by Ball (1992) in the context of a model in which higher inflation leads to increasing uncertainty over the monetary policy stance. The possibility of a negative effect of inflation on its uncertainty was then considered by Pourgerami and Maskus (1987), who pointed out that in an environment of accelerating inflation agents may invest more resources in inflation forecasting, thus reducing uncertainty (see also Ungar and Zilberfarb, 1993). Causality in the opposite direction, namely from inflation uncertainty to inflation, is instead a property of models based on the Barro–Gordon setup, such as the one due to Cukierman and Meltzer (1986).

Concerning the relationship between inflation uncertainty and real economic activity, some authors suggest that the former reduces the rate of investment by hindering long-term contracts (see, e.g., Fischer and Modigliani, 1978), or by increasing the option value of delaying an irreversible investment (see, e.g., Pindyck, 1991). Others argue that, to the extent that it is associated with increased relative price variation, it reduces the allocational efficiency of the price system (see Friedman, 1977). In contrast, Dotsey and Sarte (2000) show that inflation variability may increase investment through its impact on precautionary savings. Finally, Cecchetti (1993) suggests that a general equilibrium, representative agent model is not likely to yield a convincingly unambiguous result on the impact of uncertainty on real economic activity.

On the empirical side, a number of studies have investigated the relationship between inflation and inflation uncertainty, typically adopting an econometric framework of the GARCH type (see Engle, 1982), and providing mixed evidence (see Davis and Kanago, 2000 for a survey, and Baillie *et al*., 1996, Brunner and Hess 1993, Kontonikas 2004, Grier and Perry, 2000 for some specific contributions). Other authors take instead a VAR approach to analyse US data. In particular, Benati and Surico (2008) estimate structural VARs with time-varying parameters and stochastic volatility and report a decline in inflation predictability, showing that this can be caused by tough anti-inflation policies in the context of a sticky price model. Cogley *et al.* (2009) take a similar approach but focus instead on the inflation gap, and present evidence that this started decreasing in the early 1980s, when the inflation rate was reduced, and kept falling when Greenspan took over the chairmanship of the Fed. This result is then rationalised in terms of a new Keynesian DSGE model, within which a more stable long-run inflation target is the main factor leading to lower inflation volatility and persistence.

Empirical studies on the linkages between inflation uncertainty and real economic activity also report conflicting results both in terms of the sign (see, e.g., Holland, 1993) and of the magnitude and timing of the effects (see, e.g., Davis and Kanago, 1996, Cunningham, Tang, and Vilasuso, 1997, and Grier and Perry, 2000). Elder (2004) finds that in the US inflation uncertainty has significantly reduced real economic activity. This holds for the period prior to 1979, after 1982, and over the full post-1966 period and is robust to various specifications. This result is obtained by combining a VAR specification with a multivariate GARCH model.

The present study aims to contribute to this area of the literature by estimating a time-varying AR-GARCH model of inflation producing measures of inflation uncertainty for the euro area, and investigating the linkages between them in a VAR framework, also allowing for the possible impact of the policy regime change associated with the start of EMU in 1999. Among recent studies, O'Really and Whelan (2005) focus on inflation persistence and find relatively little instability in the parameters of the euro-area inflation process. Full-sample estimates of the persistence parameter are generally close to 1, and the hypothesis that this parameter has been stable over time cannot be rejected. Angeloni *et al.* (2006), using micro data on consumer prices and sectoral inflation rates from six euro-area countries spanning several years before and after the introduction of the euro, find no evidence of a shift around 1999. Finally, Altissimo *et al.* (2006) note that, for aggregate data, the degree of inflation persistence in the euro area appears to be very high for sample periods spanning multiple decades but falls dramatically once time variation in the mean level of inflation is allowed for; furthermore, the timing of the breaks generally coincides with observed shifts in the monetary policy regime.

The choice of focusing in the present study on the Harmonised Index of Consumer Prices (HICP) for the euro area as a whole reflects the ECB's mandate to achieve aggregate price stability in the Eurozone, the absence of instruments to fine-tune monetary policy to cyclical fluctuations in individual EMU countries, and the recognition that inflation differentials play a secondary role when calibrating the safety margin for admissible inflation in the euro area (ECB, 2003). The availability of reliable and easy-to-update measures of inflation uncertainty is particularly relevant for monetary policy purposes (Goodhart, 1999, and Greenspan, 2003). As Soderstrom (2002) notes, when there is uncertainty about the persistence of inflation, it is optimal for the central bank to respond more aggressively to shocks than when the parameters are known with certainty, in order to avoid undesirable outcomes in the future. According to Shuetrim and Thomson (2003), for certain shocks, taking into account parameter uncertainty can imply that a more, rather than less, activist use of the policy instrument is appropriate. While this finding is specific to their model specification, parameter estimates and shocks analysed, it contrasts with the widely held belief that the general implication of parameter uncertainty is a more conservative policy. Finally, Coenen (2007) argues that a cautious monetary policy-maker is well-advised to design and implement interest-rate policies under the assumption that inflation persistence is high when there is considerable uncertainty about its degree. Such policies are characterised by a relatively aggressive response to inflation developments and exhibit a substantial degree of inertia.

Surprisingly, studies on the relationship between inflation and inflation uncertainty for the euro area as a whole are distinctly rare. Most of them are either based on survey data (see, e.g., Giordani and Söderlind, 2003, and Arnold and Lemmen, 2008) or adopt a country-by-country approach (see Fountas *et al*. 2004, and Apergis, 2004). Our econometric framework is similar to that originally proposed by Evans (1991) and recently used by Berument *et al.* (2005) and Caporale and Kontonikas (2009). The latter authors analyse inflation uncertainty and its relation with inflation at country level in twelve EMU member states, and report that there is considerable heterogeneity across countries, even after the introduction of the euro, and also that the Friedman-Ball link appears to have weakened or even broken down in a number of cases in individual member states after the creation of EMU.

Our analysis extends theirs in several ways. First, as already mentioned, we focus on the euro area as a whole, consistently with the ECB's mandate. Second, we model inflation as a function of the unemployment rate rather than past inflation only, thus establishing a link between the present study and the literature on the Phillips curve in Europe. Third, we investigate empirically the relationship between inflation uncertainty and inflation in a bivariate VAR framework instead of a single-equation model.

More specifically, we take a two-step approach as in Grier and Perry (1998) and Caporale and Kontonikas (2009), namely we first estimate AR-GARCH models with time-varying parameters to generate a measure of inflation uncertainty, and then estimate the relationship between inflation and inflation uncertainty in a bivariate VAR context. The model we use in the first step regresses current inflation on lagged inflation and lagged unemployment. The advantages in terms of forecast accuracy deriving from including the unemployment rate as a measure of real economic activity in a model for inflation are discussed by Stock and Watson (1999) and Amisano and Giacomini (2007).

The paper is structured as follows. Section 2 outlines the methodology. Section 3 describes the data and the empirical results. Section 4 offers some concluding remarks, highlighting in particular the policy implications of our findings.

2. Econometric Framework

The GARCH models typically used in the literature have the drawback that they do not take into account the fact that short-run and long-run inflation uncertainty might be very different and affect inflation expectations in different ways.

As emphasised by Evans (1991), agents' temporal decisions are more likely to be affected by the conditional variance of short-run movements in inflation, whilst intertemporal decisions might be based mainly on changes in the conditional variance of long-term inflation. Moreover, one should distinguish between "structural uncertainty" (associated with the randomness in the time-varying parameters of the inflation process, and representing the propagation mechanism), which might originate, for instance, from unanticipated monetary policy changes, and "impulse uncertainty" (associated with the shocks hitting the conditional variance, which are propagated through the parameters of the inflation process).~

The econometric framework suggested by Evans (1991), and also used by Berument *et al.* (2005) and Caporale and Kontonikas (2009), has the advantage of yielding estimates of the various types of uncertainty discussed above, and is adopted here as well. More specifically, inflation is specified as a *k-th* order autoregressive process, AR(k), and is also a function of unemployment, the parameters being time-varying and the residuals following a GARCH(1,1) process. The model is the following:

$$\pi_{t+1} = \mathbf{X_t}\boldsymbol{\beta}_{\mathbf{t+1}} + e_{t+1} \qquad e_{t+1} \sim \mathrm{N}(0, \mathrm{h}_\mathrm{t})$$

$$\text{and} \quad X_t = [1 \quad \pi_t, \quad ..., \quad \pi_{t-k}, u_t \quad ..., u_{t-m}] \quad (1)$$

$$h_t = h + ae_{t-1}^2 + \lambda h_{t-1} \tag{2}$$

$$\boldsymbol{\beta}_{\mathbf{t+1}} = \boldsymbol{\beta}_{\mathbf{t}} + \mathbf{V_{t+1}} \qquad \text{where} \quad V_{t+1} \sim \mathrm{N}(0, \mathrm{Q}) \tag{3}$$

where π_{t+1} denotes the rate of inflation between t and t+1; $\mathbf{X_t}$ is a vector of explanatory variables known at time t containing current and lagged values of inflation and the unemployment rate u_t; e_{t+1} describes the shocks to the inflation process that cannot be forecast with information known at time t, and is assumed to be normally distributed with a time-varying conditional variance h_t. The conditional variance is specified as a GARCH(p,q) process, that is, as a linear function of past squared forecast errors, e^2_{t-i}, and past variances, h_{t-j}. Further, $\beta_{t+1} = [\beta^{\pi}_{0,t+1}, \beta^{\pi}_{1,t+1}, ..., \beta^{\pi}_{k,t+1}, \beta^{u}_{1,t+1}, ..., \beta^{u}_{m,t+1}]$ denotes the time-varying parameter vector, and $\mathbf{V_{t+1}}$ is a vector of shocks to $\boldsymbol{\beta}_{\mathbf{t+1}}$, assumed to be normally distributed with a homoscedastic covariance matrix $\mathbf{Q}$. The updating equations for the Kalman filter are:

$$\pi_{t+1} = \mathbf{X_t} E_t \boldsymbol{\beta}_{\mathbf{t+1}} + \varepsilon_{t+1} \tag{4}$$

$$H_t = \mathbf{X_t}\boldsymbol{\Omega}_{\mathbf{t+1|t}}\mathbf{X_t}' + h_t \tag{5}$$

$$E_{t+1}\boldsymbol{\beta}_{\mathbf{t+2}} = E_t\boldsymbol{\beta}_{\mathbf{t+1}} + [\boldsymbol{\Omega}_{\mathbf{t+1|t}}\mathbf{X_t}' H_{t-1}]\varepsilon_{t+1} \tag{6}$$

$$\boldsymbol{\Omega}_{\mathbf{t+2|t+1}} = [\mathbf{I} - \boldsymbol{\Omega}_{\mathbf{t+1|t}}\mathbf{X_t}' H_{t-1}\mathbf{X_t}]\boldsymbol{\Omega}_{\mathbf{t+1|t}} + \mathbf{Q} \tag{7}$$

where $\boldsymbol{\Omega}_{\mathbf{t+1|t}}$ is the conditional covariance matrix of $\boldsymbol{\beta}_{\mathbf{t+1}}$ given the information set at time t, representing uncertainty about the structure of the inflation process.

As Equation (5) indicates, the conditional variance of inflation (short-run uncertainty), H_t, can be decomposed into: (i) the uncertainty due to randomness in the inflation shocks e_{t+1}, measured by their conditional volatility h_t (impulse uncertainty); (ii) the uncertainty due to unanticipated changes in the structure of inflation $\mathbf{V_{t+1}}$, measured by the conditional variance of $\mathbf{X_t}\boldsymbol{\beta}_{\mathbf{t+1}}$, which is $\mathbf{X_t}\boldsymbol{\Omega}_{\mathbf{t+1|t}}\mathbf{X_t}' = S_t$ (structural uncertainty). The standard GARCH model can be obtained as a special case of this model if there is no uncertainty about $\boldsymbol{\beta}_{\mathbf{t+1}}$, so that $\boldsymbol{\Omega}_{\mathbf{t+1|t}} = \mathbf{0}$.

In this case, the conditional variance of inflation depends solely on impulse uncertainty. Equations (6) and (7) capture the updating of the conditional distribution of $\boldsymbol{\beta}_{\mathbf{t+1}}$ over time in response to new information about realised inflation. As indicated by Equation (6), inflation innovations, defined as ε_{t+1} in Equation (4), are used to update the estimates of $\boldsymbol{\beta}_{\mathbf{t+1}}$. These estimates are then used to forecast future inflation. If there are no inflation shocks and parameter shocks, so that $\pi_{t+1} = \pi_t = ... = \pi_{t-k}$ for all t, we can calculate the steady-state rate of inflation, π^*_{t+1}, as:

$$\pi^*_{t+1} = \beta^*_{0,t+1}\left[1 - \left(\sum_{i=1}^{k} \beta^{\pi}_{i,t+1}\right)\right]^{-1} \tag{8}$$

where

$$\beta^{*}_{0,t+1} = \left\lfloor \beta^{\pi}_{0,t+1} - \left(\sum_{j=1}^{m} \beta^{u}_{j,t+1}\right) u^{*} \right\rfloor \tag{9}$$

and u* is the steady-state unemployment rate, computed as the sample mean.

The conditional variance of steady-state inflation is then given by:

$$\sigma^2_t(\pi^{*}_{t+1}) = \nabla E_t \boldsymbol{\beta}_{\mathbf{t+1}} \boldsymbol{\Omega}_{\mathbf{t+1|t}} \nabla E_t \boldsymbol{\beta}^{'}_{\mathbf{t+1}} \tag{10}$$

where $$\nabla E_t \beta^{'}_{t+1} = \begin{bmatrix} \left[1-\left(\sum_{i=1}^{k}\beta^{\pi}_{i,t+1}\right)\right]^{-1} \\ E_t\beta^{*}_{0,t+1}\left[1-\left(\sum_{i=1}^{k}\beta^{\pi}_{i,t+1}\right)\right]^{-2} \\ \dots \\ E_t\beta^{*}_{0,t+1}\left[1-\left(\sum_{i=1}^{k}\beta^{\pi}_{i,t+1}\right)\right]^{-2} \end{bmatrix} \tag{11}$$

is a (k + m + 1 x 1) vector.

Having obtained uncertainty measures for each country, it is possible to analyse the link between inflation uncertainty and the level of inflation. In particular, we estimate a bivariate VAR model of inflation and steady-state inflation uncertainty, since the ECB focuses on long-run price stability. The model also includes a dummy variable to allow for possible structural breaks in the underlying relationship reflecting the introduction of the euro.

Specifically, the estimated model is the following:

$$\begin{pmatrix} \pi_t \\ unc_t \end{pmatrix} = A(L)\begin{pmatrix} \pi_{t-1} \\ unc_{t-1} \end{pmatrix} + BD_t + \begin{pmatrix} \varepsilon^{\pi}_t \\ \varepsilon^{unc}_t \end{pmatrix} \tag{11}$$

where unc_{t+1} represents steady-state uncertainty (i.e. $\sigma^2_t(\pi^{*}_{t+1})$), D_t is an intercept shift dummy variable, A(L) a matrix polynomial and B a 2×1 matrix. In the model specified above, the break date is imposed exogenously to coincide with the introduction of the euro in December 1998.

Subsequently, we carry out Granger-causality tests and also apply a sequential dummy approach to detect possible breaks endogenously. The motivation for the latter type of analysis comes from the literature arguing that rational agents are likely to react to the announcement of a regime switch before its implementation, and therefore breaks in the relationship of interest could have occurred before January 1999 (see Wilfling, 2004, and Wilfling and Maennig, 2001) .

3. Empirical Analysis

3.1 Data Description

Our preferred measure of inflation is the monthly rate of change of the seasonally unadjusted Harmonised Index of Consumer Prices (HICP) for the Eurozone as a whole.[1] The sample period is 1980:m1 – 2009:m2. Chart 1 below plots this series from 1990 onwards together with the corresponding year-on-year growth rate. Visual inspection suggests declining inflation and relatively subdued variability in the run-up to EMU and stable inflation afterwards with increased variability, particularly so towards the end of the sample.

Chart 1. EMU Inflation

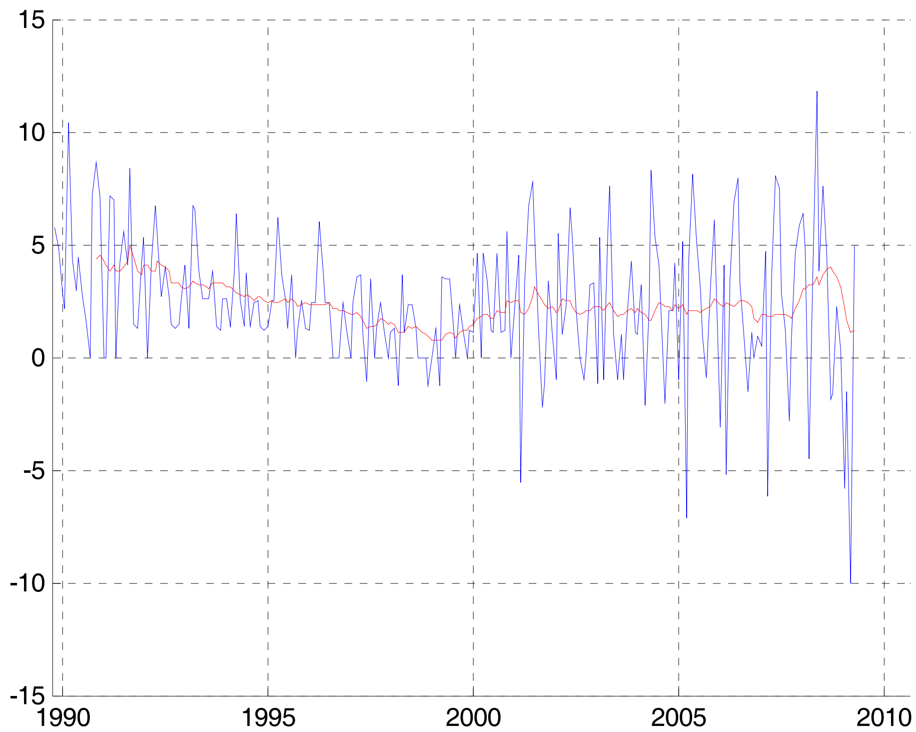


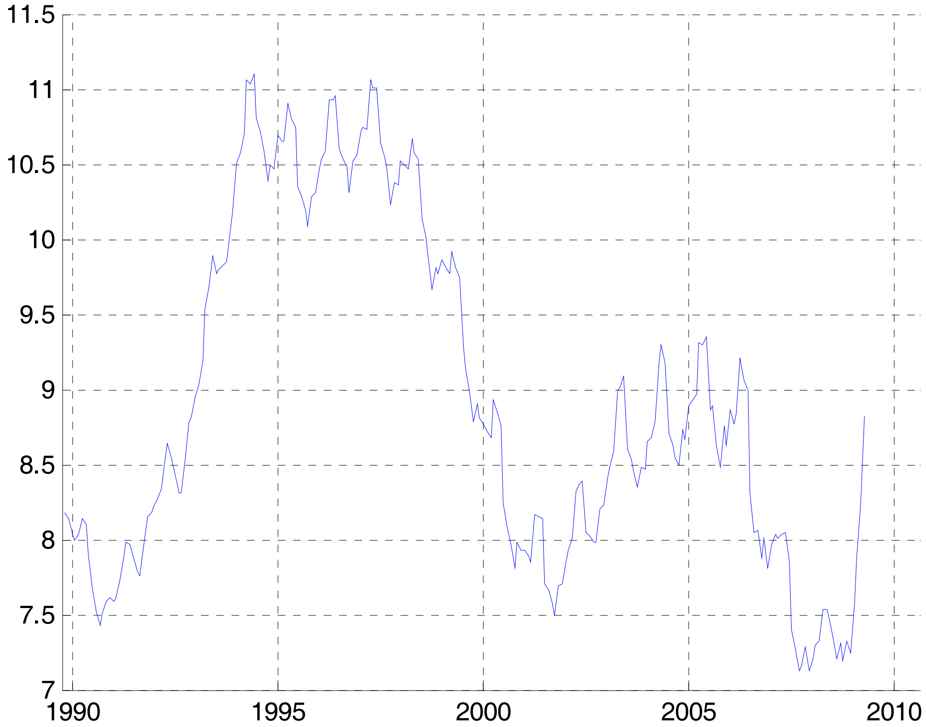
Chart 2 shows the seasonally unadjusted monthly unemployment rate. The series exhibits a cyclical pattern and is clearly linked to output developments over the same period. Finally Chart 3 contains the scatter plot of inflation against the unemployment rate. Consistently with the expectations-augmented Phillips curve framework (Phelps, 1967, and Friedman, 1968), it suggests that there is an inverse relationship between inflation and the unemployment rate and that a downward shift might have occurred at the start of EMU. Further, the curve appears to be flatter in the euro years. [2]

[1] "Euro area (changing composition) - HICP - Overall index, Monthly Index, Eurostat, neither seasonally nor working day adjusted". Eurostat code: VAL.IDX05.EMU.00. This series is available from January 1990, and has been extended backwards using a weighted average of the growth rates of the corresponding national series for France, Germany, Italy and Spain.

[2] "Euro area (changing composition) - Standardised unemployment rate, Total (all ages), Total (male & female), Eurostat, neither seasonally or working day adjusted, percentage of civilian workforce". The series is available from January 1993, with Eurostat code:

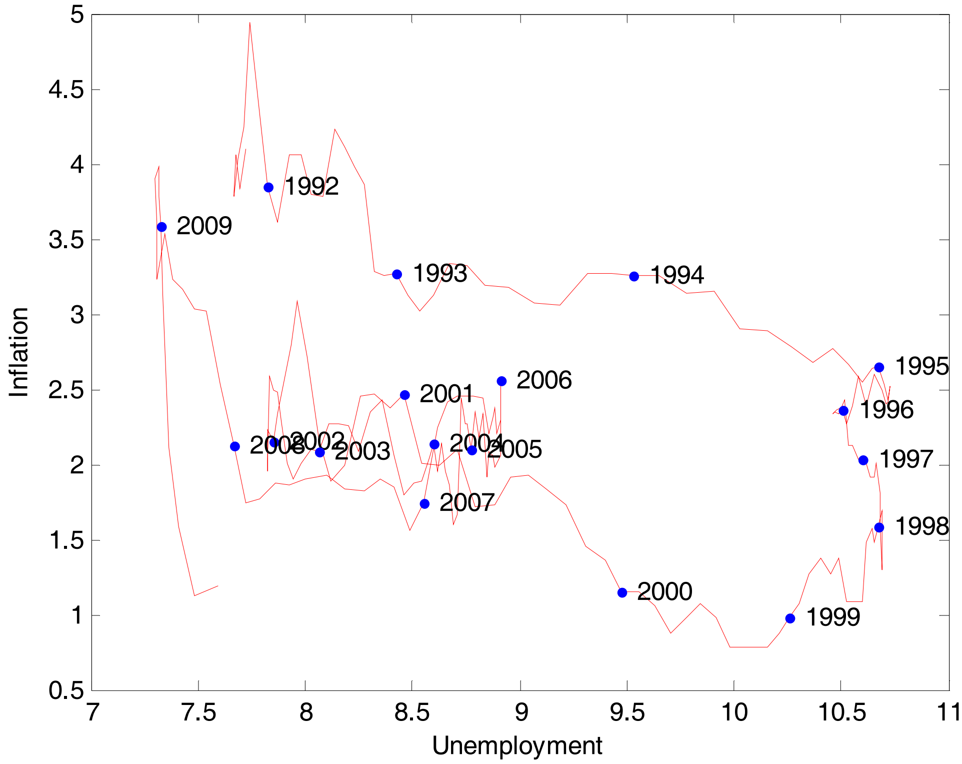
This might reflect declining inflation expectations in response to the adoption of a single monetary policy pursuing long-term price stability, as specified in the Maastricht Treaty. As the Bundesbank (2009) points out, all surveys for medium- to long-term inflation expectations in the euro area have indicated a rate just under 2% since the euro was launched in January 1999. This suggests that the Eurosystem was successful in creating and maintaining credibility, convincing individuals and enterprises that resolute action in terms of interest rates would be taken in the presence of serious inflationary risks.

Chart 2. EMU Unemployment rate



VAL.EZ.CURRENT.UNEM.TOT.M_PC_UNEM_TOT.NSA. It is linked backwards until February 1982 with the series "EU12 including West Germany - Standardised unemployment rate, Total (all ages), Total (male & female), Eurostat, neither seasonally nor working day adjusted, percentage of civilian workforce", Eurostat code: VAL.UNER.T.A0_99.ORIGINAL.EC_85.M. The two remaining years are completed by using the yearly growth rate of the series "Euro area - UNEMPLOYMENT, STANDARDISED (EU/ILO DEFINITION)(MU16) SA", from the BIS database: Macro-economic series (Blocks A-K, Q-W).

Chart 3. Inflation-unemployment cross-plot



3.2 Trend Inflation, Steady-State Inflation and Inflation Persistence

Our benchmark univariate model of the inflation process regresses current inflation on a constant term and three lags of both inflation and the unemployment rate. A Bayesian approach is taken for the estimation. We start from an appropriate but weakly informative prior, and use the 1980s as a pre-sample to update it. We report the results from 1990m1 onwards, therefore including the period immediately before the Maastricht Treaty, the convergence period to meet the requirements of the Treaty, and the first decade of EMU. After the estimation, and in order to check whether there is time variation in the relevant parameters of the models, we perform ADF tests on trend inflation, inflation persistence and the sum of the coefficients of unemployment.

Table 1. Tests for time variation

Coefficients of AR component			
	Constant	Sum(π_{t-1}, π_{t-2}, π_{t-3})	Sum(u_{t-1}, u_{t-2}, u_{t-3})
ADF test (p-value)	-2.072 (0.256)	-1.558 (0.503)	-2.528 (0.110)
Coefficients of GARCH component			
	h	a	λ
Parameter values	0.0005	0.0796	0.9161
T-stats based on covariance matrix	1.36	3.42*	37.98**
Squared standardized residuals serial correlation			
LB(1)	LB(2)	LB(3)	LB(4)
2.59	5.69	6.95	7.06

In all cases we conclude that there is time variation, either in the form of non-stationarity or in the form of structural breaks, which justifies the use of time-varying parameters in the AR equation. Similarly, we test for the significance of the ARCH and GARCH parameters of the model, finding that they are both significant. We report in Table 1 the Ljung-Box test statistic for serial correlation of the residuals at lags 1 to 4, showing that there is no residual autocorrelation at 5% level.

Charts 4-9 are based on the estimation results. Chart 4 shows trend inflation, namely the estimated time-varying constant from Equation (4), which declines steadily from the beginning of the 1990s up to the start of EMU, stabilising afterwards. The estimates of the individual coefficients on lagged unemployment (not reported) are not particularly informative, but it is clear that their sum, shown in Chart 5, is negative throughout the sample. This is consistent with our earlier interpretation of the data based on the Phillips curve model. It suggests again that this curve has become flatter since the introduction of the euro, possibly reflecting a more firm anchoring of inflation expectations.

Chart 4. EMU Trend inflation (month on month)

Chart 6 plots inflation persistence, defined as the sum of its autoregressive coefficients. This exhibits a positive trend becoming negative at the beginning of the third Phase of EMU, which may reflect lower and better-anchored inflation expectations in the post-1999 period, consistently with the ECB's mandate of achieving long-term price stability and with theory. For instance, Erceg and Levin (2003) show that inflation displays very little persistence if the long-run inflation target is constant. Similarly, Orphanides and Williams (2005) demonstrate that if long-run inflation expectations are well anchored, then inflation will be less persistent than if the public is uncertain about the long-run inflation target.

Chart 5. Total impact of unemployment on inflation

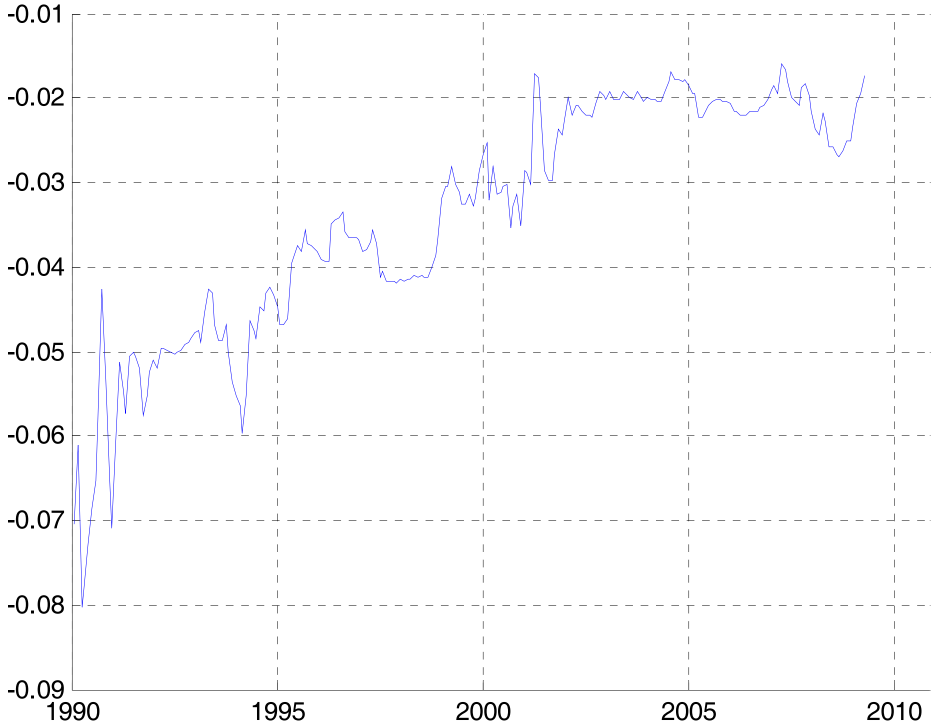


Chart 6. Inflation persistence

Chart 7 plots steady-state inflation as defined by Equation (8). It shows a gradual but marked decline for most of the sample period. From 1999, whilst the downward trend gradually flattens out, volatility appears to increase, especially in the early stages of EMU and in the most recent period.

Chart 7. Steady state inflation

3.3 Short-Run and Steady-State Inflation Uncertainty

Chart 8 shows short-run inflation uncertainty and its structural component estimated using Equation (5). A sharp decline in the early part of the sample is followed by an upward shift and higher volatility around a positive trend afterwards, the biggest spike coinciding with the latest financial turmoil. Decomposing short-run uncertainty into impulse and structural uncertainty reveals that the latter dominates, i.e. exogenous shocks are the main driving force of swings in short-run uncertainty, whilst structural uncertainty (which can be interpreted as reflecting unanticipated changes in monetary policy) is relatively stable throughout the sample period and does not play a significant role.

The focus of the ECB is, however, on medium- to long-term price stability, and therefore a more relevant measure to consider is steady-state inflation uncertainty (see Chart 9). This appears to follow a downward trend with the exception of the early years of the euro, when agents were still learning about the new monetary environment. The steady decline since then can be seen as an indication of the ECB's success in fulfilling its mandate as specified in the Maastricht Treaty.

Chart 8. Short-run and structural inflation uncertainty

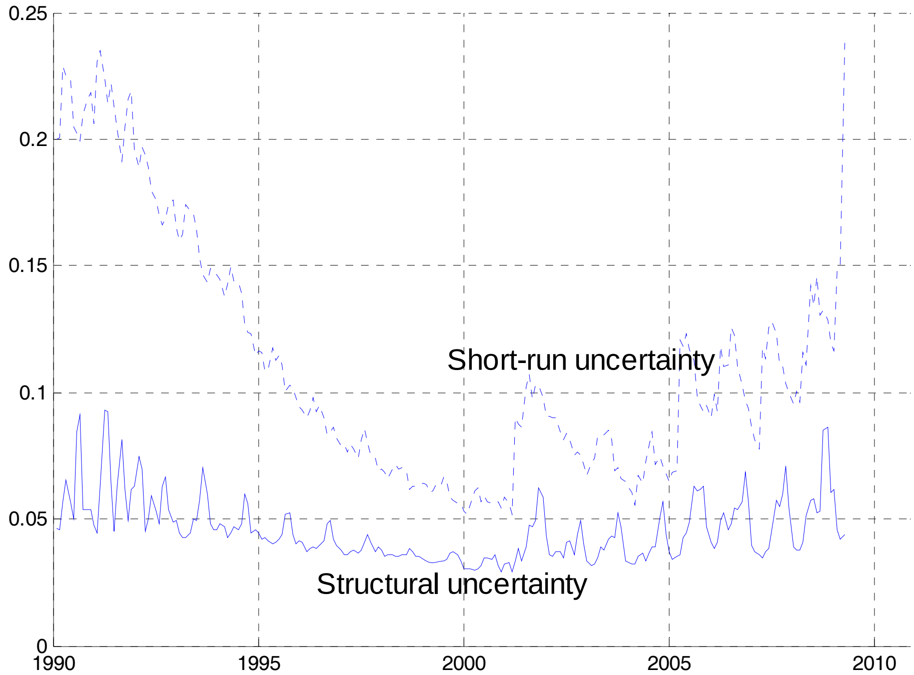


Chart 9. Steady-state inflation uncertainty

4. The Relationship between Inflation and Steady-State Inflation Uncertainty in a Bivariate VAR Framework

Next, we estimate a bivariate VAR for inflation and steady-state inflation uncertainty to test for causality and for the presence of structural breaks related to EMU. As a first step, the order of integration of the variables needs to be established. Standard ADF tests reject the null hypothesis of a unit root in levels in both cases. We then carry out the unit root tests proposed by Saikkonen and Lütkepohl (2002) and Lanne *et al.* (2002). These are based on first estimating the deterministic term by a generalized least squares (GLS) procedure under the unit root null hypothesis and subtracting it from the original series. Then an ADF-type test is performed on the adjusted series which also includes terms to correct for estimation errors in the parameters of the deterministic part. As in the case of the ADF statistic, the asymptotic null distribution is non-standard. Critical values are tabulated in Lanne *et al.* (2002). The test results when allowing for exogenous breaks indicate that steady-state uncertainty might not be stationary, but those including endogenously determined breaks again support stationarity of both series (see Table 2).

Table 2. Unit root tests on inflation and steady-state inflation uncertainty

		INFLATION		SS_UNC	
Test	Lag	Test statistic	Det. Comp.	Test statistic	Det. Comp.
ADF	3	-6.80 **	C, T, SD	-4.45 **	C, T, SD
	REC	-3.50 * (5)	C, T, SD	-4.45 ** (3)	C, T, SD
EXO_SB	3	-5.14**	C, T, SD 1998 M12	-2.74	C, T, SD 1998 M12
	REC	-3.03 * (5)	C, T, SD 1998 M12	-1.58 (5)	C, T, SD 1998 M12
END_SB	3	-8.14 **	C, T, SD 2008 M5	-3.58 **	C, T, SD 2001 M3
	REC	-1.52 (11)	C, T, SD 2008 M7	-3.58 ** (3)	C, T, SD 2001 M3

Table 3. Johansen Trace Test for: [infl, infl_ss_var]

Test 1: included lags (levels): 3, trend and intercept, seasonal dummies included					
Coint_rank	Test stat.	p_val	90%	95%	99%
r0 = 0	91.11	0 .0000	23.32	25.73	30.67
r0 = 1	19.71	0.0020	10.68	12.45	16.22
Test 2: included lags (levels): 1, trend and intercept, seasonal dummies included					
r0 = 0	137.26	0.0000	23.32	25.73	30.67
r1 = 1	21.37	0.0009	10.68	12.45	16.22
Test 3: included lags (levels): 8, trend and intercept, seasonal dummies included					
r0 = 0	67.20	0.0000	23.32	25.73	30.67
r1 = 1	16.20	0.0101	10.68	12.45	16.22

Johansen's cointegration tests confirm that both inflation and inflation uncertainty can be treated as stationary in levels (see Table 3). Therefore the remainder of the analysis is carried out under this assumption, and a VAR in levels was estimated.

Standard selection criteria suggested choosing lag length 14 for the bivariate VAR in levels. The deterministic component was specified to include a constant and a trend, as well as seasonal dummies and a shift dummy in 1998m12 to capture the introduction of the euro. Standard diagnostic tests indicate that the model is statistically adequate. Granger-causality tests imply uni-directional causality running from steady-state inflation uncertainty to inflation at the 5% confidence level, consistently with the Cukierman-Meltzer model, and bi-directional causality at the 10% level **(see Table 4)**.

Table 4. Test for Granger-causality, dummy 1998m12

H0: "infl_ss_var" does not Granger-cause "infl" – whole sample	
Test statistic l = 1.8757	pval-F(l; 14, 342) = 0.0279
H0: "infl" does not Granger-cause "infl_ss_var" – whole sample	
Test statistic l = 1.6509	pval-F(l; 14, 342) = 0.0644

The model was also estimated including a sequential intercept shift dummy in order to test endogenously for possible structural breaks. Chart 10 shows the sequential t-value of the corresponding coefficient in the equation for steady-state inflation uncertainty. As can be seen, a sizeable downward shift is apparent at the end of 2000. This is consistent with the earlier finding of declining and more stable steady-state inflation uncertainty in the euro years.

Chart 10. Sequential t-value of the dummy coefficient

The results of full-sample Granger-causality tests using the specification including a shift dummy for the break detected in 2000m1 are consistent with the earlier ones. However, when carried out separately over the two sub-samples 1990m1-1998m6 and 2001m3-200m2, such tests provide evidence of a reversal in causality, implying that the Cukierman-Meltzer link has been replaced by the Friedman-Ball one in the second subsample (see Table 5), and that the ECB can effectively reduce inflation uncertainty by bringing down the inflation rate.

Table 5. Test for Granger-causality, p – values

Ho	Whole sample	1980:m1-1998:m6	2001:m3-2009:m2
Inflation does not Granger-cause inflation variability	0.03	0.06	0.09
Inflation variability does not Granger-cause inflation	0.41	0.17	0.17

5. Conclusions

This paper estimates a time-varying AR-GARCH model of inflation for the euro area, and investigates its linkages with the resulting measures of inflation uncertainty in a bivariate VAR framework, also modelling the possible structural break resulting from the creation of EMU at the beginning of 1999. Obtaining accurate measures of inflation uncertainty is crucial for monetary authorities, since higher uncertainty requires more active policies, as pointed out by Soderstrom (2002). Our main findings are as follows. Steady-state inflation and inflation uncertainty have both declined steadily since the inception of EMU, whilst short-run uncertainty has increased, but mainly owing to exogenous shocks. A sequential dummy procedure provides further evidence of a break coinciding with the introduction of the euro and leading to lower long-run uncertainty.

Interestingly, our analysis suggests that a tough anti-inflation stance successfully reduces long-run uncertainty in the case of the Eurozone, whilst the opposite holds for the US, where tighter policies adopted by the Fed seem to lead to lower predictability (see Benati and Surico, 2008 and Cogley *et al.*, 2009). This might reflect differences in the mandate of the two central banks. Whilst the ECB is responsible for price stability only, the Fed pursues both price stability and full employment, two goals which at times can be in conflict. Moreover, whilst the ECB adopts a clear quantitative definition of price stability, no such definition is available in the case of the Fed.

It also appears that the direction of causality has been reversed, and that in the euro period the Friedman-Ball link is empirically supported, implying that the ECB can achieve lower inflation uncertainty by lowering the inflation rate. This is consistent with Fountas *et al.* (2004) and Conrad and Karanasos (2005) and with the idea that, given the ECB's mandate and its record so far, any long-lasting deviations from price stability in the Eurozone would surprise market participants and lead to higher inflation uncertainty.

Overall, these results give support to the view expressed by the President of the ECB, Jean-Claude Trichet, a few years ago, when he argued that the single monetary policy and its clear focus on long-run price stability had helped anchor medium- to long-run inflation expectations in the Euro area, thus reducing inflation uncertainty (Trichet 2004). Short-run inflation uncertainty might have risen since 1999, but this does not appear to reflect higher structural uncertainty associated with unanticipated monetary policy changes, but possibly remaining differences across member states, fiscal policy, and financial and commodity market price shocks.

This analysis can be extended in several directions. First, the regression model for the inflation process can be expanded to include other possible determinants such as industrial production, money growth, the yield curve, the exchange rate and credit. Second, whether the observed patterns can indeed be attributed to the role played by the ECB can be tested by comparing our findings with those obtained for a control group of countries, such as other European countries outside the Eurozone, the US, and Canada. Third, estimation of a time-varying VAR-GARCH model could be carried out instead of treating unemployment and other possible variables as exogenous. Fourth, the forecasting properties of the model could be investigated.

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