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by Guglielmo Maria Caporale,
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by Guglielmo Maria Caporale [2], Luca Onorante [3]
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In 2010 all ECB publications feature a motif taken from the €500 banknote.

NOTE: This Working Paper should not be reported as representing the views of the European Central Bank (ECB). The views expressed are those of the authors and do not necessarily reflect those of the ECB.

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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 their linkages 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, consistently with the idea 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

Non-technical summary

The launch of EMU in 1999 changed significantly the European macroeconomic environment and the way agents formulate their expectations. Inflation and inflation expectations are among the variables which most likely were affected by this change, for three main reasons. First, the Maastricht Treaty established a new institution, the ECB, with the task of ensuring price stability for the euro area as a whole. Second, the ECB adopted a new monetary strategy, i.e. the two-pillar framework, and a set of new instruments. Third, new monetary policy transmission channels emerged while others disappeared (e.g. idiosyncratic exchange rate adjustment).

Obtaining accurate measures of inflation uncertainty is crucial for monetary authorities, since higher uncertainty requires more active policies, as pointed out by Soderstrom (2002). This paper focuses on the relationship between inflation and inflation uncertainty in the euro area. It contributes to the literature by estimating a time-varying AR-GARCH model of inflation for the euro area as a whole with the aim of producing measures of inflation uncertainty and investigating the linkages between these two variables in a VAR framework, also allowing for the possible impact of the policy regime change associated with the start of EMU.

In contrast to most existing studies, our analysis distinguishes between short-run and steady-state uncertainty. The former includes both structural uncertainty (associated with the randomness in the time-varying parameters of the inflation process which might result, for instance, from economic agents adapting to a new economic environment) and impulse uncertainty, which reflects exogenous shocks to the system. Both types of short-run uncertainty affect forecast errors. Steady-state uncertainty instead reflects uncertainty when variables are at their steady state levels and there are no shocks to the system.

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 stabilised both in its impulse and parameter uncertainty component. A sequential dummy procedure provides further evidence of a break coinciding with the introduction of the euro and leading to lower long-run uncertainty.

Our analysis suggests that a tough anti-inflation stance successfully reduces long-run uncertainty in the case of the euro area, and 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.

1. Introduction

The launch of EMU in 1999 changed significantly the European macroeconomic environment and the way agents formulate their expectations. The introduction of the euro, the transfer of monetary policy from domestic authorities to the European Central Bank (henceforth ECB) and the almost complete elimination of residual barriers to financial and economic integration were all part of this process.

Inflation and inflation expectations are among the variables which most likely were affected by this change, for three main reasons. First, the Maastricht Treaty established a new institution, the ECB, with the task of ensuring price stability for the euro area as a whole. Second, the ECB adopted a new monetary strategy, i.e. the two-pillar framework, and a set of new instruments. Third, new monetary policy transmission channels emerged while others disappeared (e.g. idiosyncratic exchange rate adjustment).

This paper focuses on the relationship between inflation and inflation uncertainty in the euro area. Surprisingly, studies of this type 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). The present study contributes to this area of the literature by estimating a time-varying AR-GARCH model of inflation for the euro area as a whole with the aim of producing measures of inflation uncertainty and investigating the linkages between these two variables in a VAR framework, also allowing for the possible impact of the policy regime change associated with the start of EMU.

In contrast to most existing studies, our analysis distinguishes between short-run and steady-state uncertainty. The former includes both structural uncertainty (associated with the randomness in the time-varying parameters of the inflation process which might result, for instance, from economic agents adapting to a new economic environment) and impulse uncertainty, which reflects exogenous shocks to the system. Both types of short-run uncertainty affect forecast errors. Steady-state uncertainty instead reflects uncertainty when variables are at their steady state levels and there are no shocks to the system.

As emphasised by Evans (1991), uncertainty is not the same as variability. If agents have very little information about inflation, they may view the future as highly uncertain even though the econometrician observes little *ex post* volatility. Conversely, there may be very little uncertainty associated with a large change in actual inflation because agents have a good deal of information in advance.

The need to distinguish between different types of uncertainty is motivated by economic theory suggesting that economic agents take two types of decisions: intra-temporal and intertemporal ones. The former are more likely to be affected by the conditional variance of short-run movements in inflation, as shown by Lucas (1973) with respect to production and by Deaton (1977) for consumption. The latter instead might be influenced by long-run inflation uncertainty through its impact on interest rates, risk premia and debt maturity as suggested by Klein (1975) amongst others.

The econometric framework we use to identify the different types of uncertainty is similar to that originally proposed by Evans (1991) and recently used by Berument *et al.* (2005) and Caporale and Kontonikas (2009), which consists of two steps. First, an AR-GARCH model with time-varying parameters is estimated to generate different measures of inflation uncertainty based on lagged values of inflation only. The use of the Kalman filter at this stage of the analysis is consistent with agents forming expectations using available information efficiently. In the second step, taking these expectations and the related measures of uncertainty as given, the relationship between inflation and inflation uncertainty is tested in a univariate framework.

The main advantage of this procedure is that in the first step we put ourselves in the position of economic agents formulating expectations within the model. In the second step we adopt the point of view of an external observer using the most appropriate framework to analyse empirically the interaction between the variables of interest. A potential shortcoming is that one of the two variables used in step two is a generated regressor and therefore a stochastic variable. This may reduce the efficiency of the estimation.

We extend this framework along two directions. First, we consider other variables, beside lagged inflation values, which are likely to affect inflation expectations. The set of new variables includes: the unemployment rate, a measure of the output gap, the rate of change of M3, nominal short- and long-term interest rates, the interest rate differential, and the nominal effective exchange rate. Second, we analyse the relationship between actual inflation and inflation uncertainty in a multivariate context, better suited to deal with the possibility of reverse causality.

The choice of focusing on euro area inflation measured as the month-on-month rate of change of the Harmonised Index of Consumer Prices (HICP) reflects the ECB mandate to achieve aggregate price stability in the euro area and the absence of instruments to fine-tune monetary policy to cyclical fluctuations in individual EMU countries.

The paper is structured as follows. Section 2 contains a brief literature review. Section 3 outlines the methodology. Section 4 describes the data and the empirical results. Section 5 offers some concluding remarks, highlighting in particular the policy implications of our findings.

2. Literature review

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 allocative 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.

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 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 in contrast 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.

3. Econometric Framework

Usually inflation uncertainty is estimated by means of GARCH models which 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. The econometric framework we employ has the advantage of yielding estimates of the various types of uncertainty discussed above. More specifically, inflation is specified as a *k-th* order autoregressive process, AR(k), and is also a function of other relevant economic variables , the parameters being time-varying and the residuals following a GARCH(1,1) process. The model is the following:

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

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

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

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 including a constant term, inflation, the unemployment rate, the rate of change of M3 in nominal terms, the rate of change of the nominal effective exchange rate, the differential between long- (10 years) and short-term (3 months) interest rates;[1] 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 (1, 1) process, that is, as a linear function of past squared forecast errors, e^2_{t-i}, and past variances, h_{t-j}. Further, $\boldsymbol{\beta}_{\mathbf{t+1}}$ 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 $\mathbf{Q}$.

At each point in time, economic agents form their inflation expectations and update their estimation of the parameters β on the basis of the available information. The optimal predictor in this context is the Kalman filter, including equations (1) to (3) and the following updating equations:

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

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

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

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

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

[1] 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). We extend the model to include also other potential macroeconomic determinants of 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 $\mathbf{X_t \beta_{t+1}}$, which is $\mathbf{X_t \Omega_{t+1|t} X_t}' = S_t$ (structural uncertainty). The standard GARCH model can be obtained as a special case of this model if there is no uncertainty about $\mathbf{\beta_{t+1}}$, so that $\mathbf{\Omega_{t+1|t}} = \mathbf{0}$. 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 $\mathbf{\beta_{t+1}}$ 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 $\mathbf{\beta_{t+1}}$. These estimates are then used to forecast future inflation. If there are no inflation 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

$$\beta^*_{0,t+1} = \left[\beta_{0,t+1} + \left(\sum_{j=1}^{m} \beta^{\upsilon}_{j,t+1}\right)\upsilon^* + \left(\sum_{j=1}^{m} \beta^{\mu}_{j,t+1}\right)\mu^* + \left(\sum_{j=1}^{m} \beta^{\xi}_{j,t+1}\right)\xi^* + \left(\sum_{j=1}^{m} \beta^{\delta}_{j,t+1}\right)\delta^*\right] \tag{9}$$

and υ^* is the steady-state unemployment rate, μ^* the steady-state rate of change of M3 in nominal terms, ξ^* the rate of change of the nominal effective exchange rate, and δ^* the differential between long- and short-term interest rates. In all cases steady-state values are computed as sample means. The conditional variance of steady-state inflation is then given by:

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

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

is a $(k + m + 1 \times 1)$ vector. Having obtained the three uncertainty measures, i.e. impulse uncertainty, structural uncertainty and steady-state uncertainty, we analyse their links with inflation by estimating 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 as defined in (10), 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).

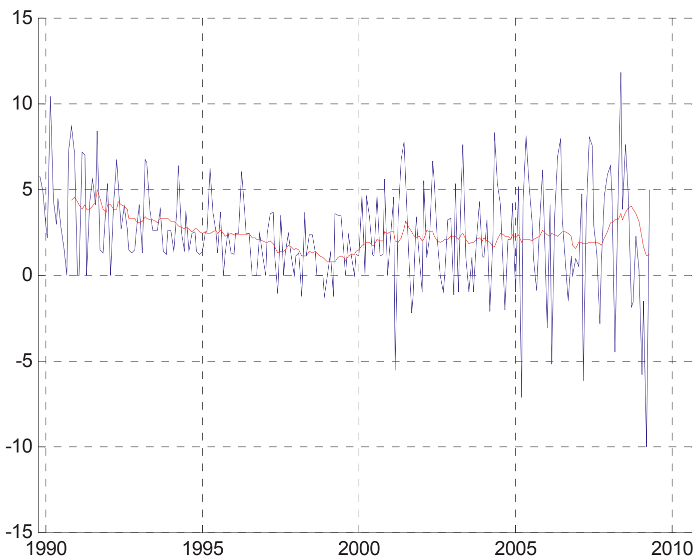
4. Empirical Analysis

4.1 Data Description

The vector of explanatory variables includes: inflation, the unemployment rate, the rate of change of M3 in nominal terms, the rate of change of the nominal effective exchange rate, the differential between long- (10 years) and short-term (3 months) interest rates. The sample period is 1980:m1 – 2009:m2.

Our preferred measure of inflation is the monthly rate of change of the seasonally unadjusted Harmonised Index of Consumer Prices (HICP) for the euro area.[2] Chart 1 below plots this series from 1990 onwards, scaled up by a factor 12, 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.

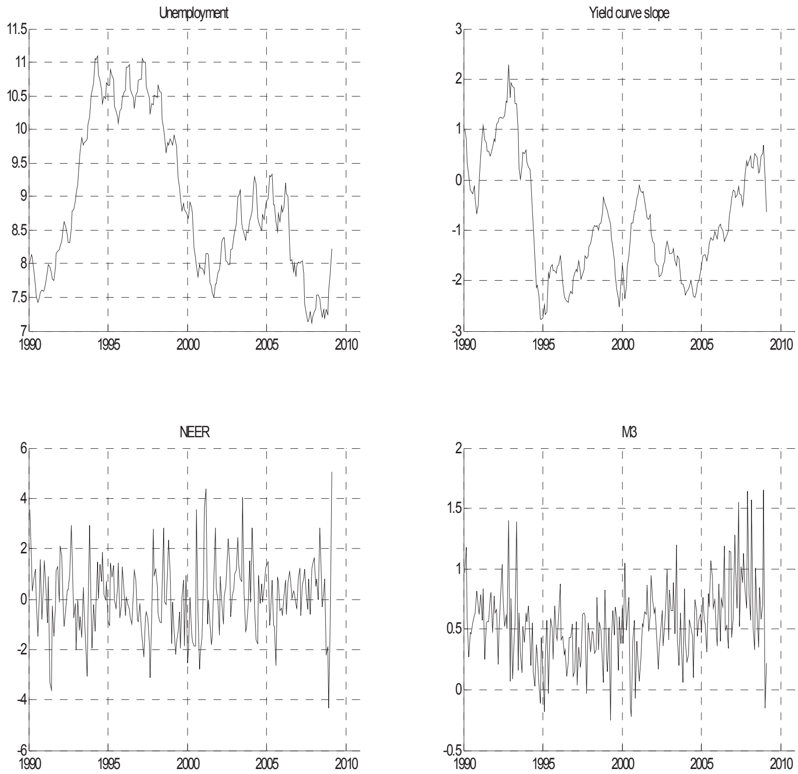
Chart 1. Euro area Inflation



[2] The data appendix provides details of the time series used for the analysis. Alternative series for inflation have been tested, leading to comparable results.

Chart 2 shows the four macroeconomic variables which we include in the model in addition to lagged inflation. The upper left panel shows the euro area seasonally unadjusted monthly unemployment rate. This series exhibits a cyclical pattern.

Chart 2. Unemployment, yield curve slope, NEER change, M3 change,



The upper right panel shows the yield curve, namely the difference between the short- and the long-term interest rate. This series has negative and fluctuating values over most of the sample period (implying a positively sloped yield curve). Slope inversion is observed only at times of heightened inflationary pressures. The lower left panel shows the logarithm of the nominal effective exchange rate in first differences. This series exhibits a jagged pattern and is consistent with the euro appreciation after 2001. Finally, the lower right panel shows the logarithm of M3 in first differences. This series has a very uneven pattern which is consistent with M3 growing at an increasing pace in the second half of the sample period.

4.2 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 inflation and the other regressors as indicated above. 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, 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 the estimated coefficients. Twelve out of sixteen coefficients (at least one for each variable) exhibit 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 (these tests are available on request). Similarly, we test for the significance of the ARCH and GARCH parameters of the model, finding that they are both significant (Table 1, upper panel). In Table 1 we also report the Ljung-Box test statistic for serial correlation of the residuals at lags 1 to 4, showing that there is no residual autocorrelation at the 5% level.

Table 1. Tests for time variation

Coefficients of GARCH component			
	h	a	Λ
Par. Val.	0.0008	0.0793	0.9060
T-stats on cov. matrix	0.7663	2.3525	19.0580
Squared standardized residuals serial correlation			
LB(1)	LB(2)	LB(3)	LB(4)
2.00	2.88	4.06	7.15

Charts 3-8 are based on the estimation results. Chart 3 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. Chart 4 shows the sum of the estimated coefficients for each of the four regressors other than inflation together with standard confidence bands. The unemployment coefficients (upper left panel) are consistent with a Phillips curve interpretation and with the well-documented flattening of the curve itself in the most recent years.[7] The cumulated yield curve coefficients (upper right panel) have the expected sign. In particular, in the first half of the sample, with the exception of the 1992 EMS crisis, the estimated coefficients are consistent with monetary policy being used in many countries to reduce inflation to values compatible with the Maastricht inflation criterion. Neither the coefficients on the nominal exchange rate (lower left panel), nor those on the rate of change of M3 (lower right panel) are statistically significant.[8]

[7] For evidence in the US and in the euro area see, respectively, Atkeson and Ohanian (2001) and Fischer, Lenza, Pill and Reichlin (2009).

[8] However, in the case of the exchange rate, shrinking confidence bands with respect to the pre-EMU period are consistent with a progressive stabilisation of market expectations.

Chart 3. EMU Trend inflation (month on month)

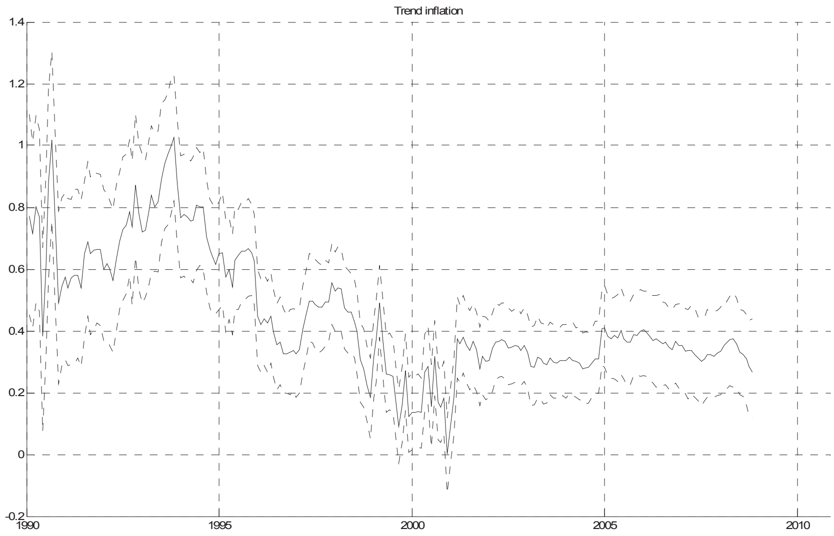


Chart 4. Time-varying beta coefficients

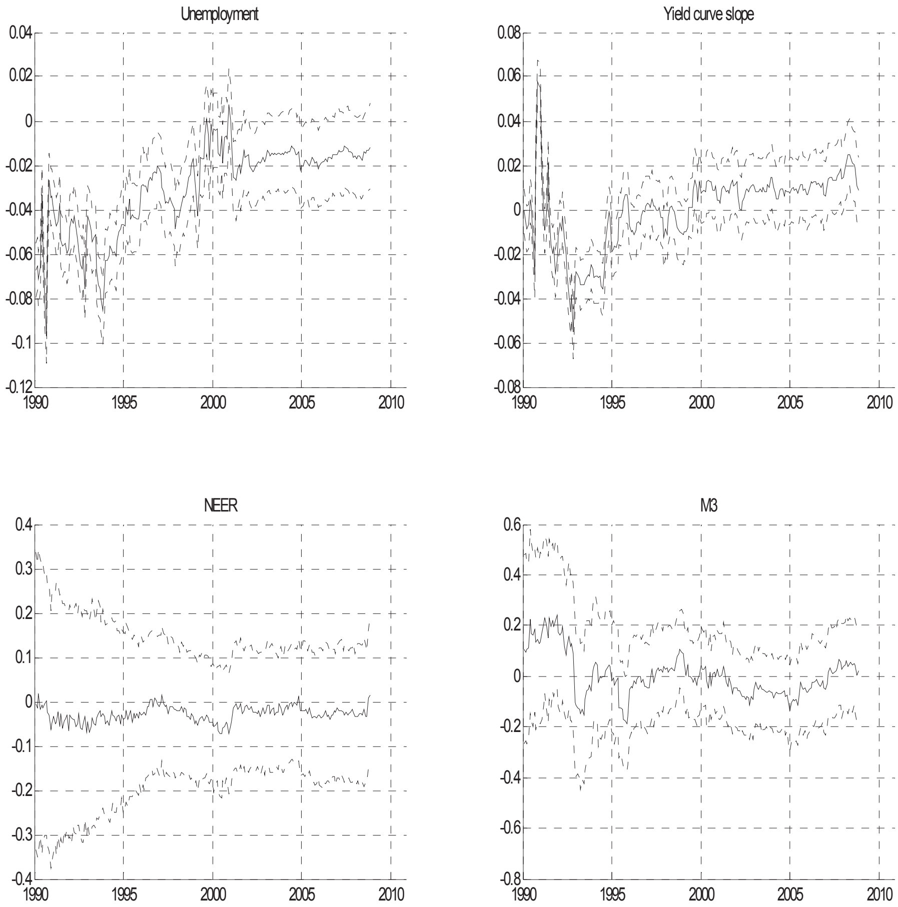
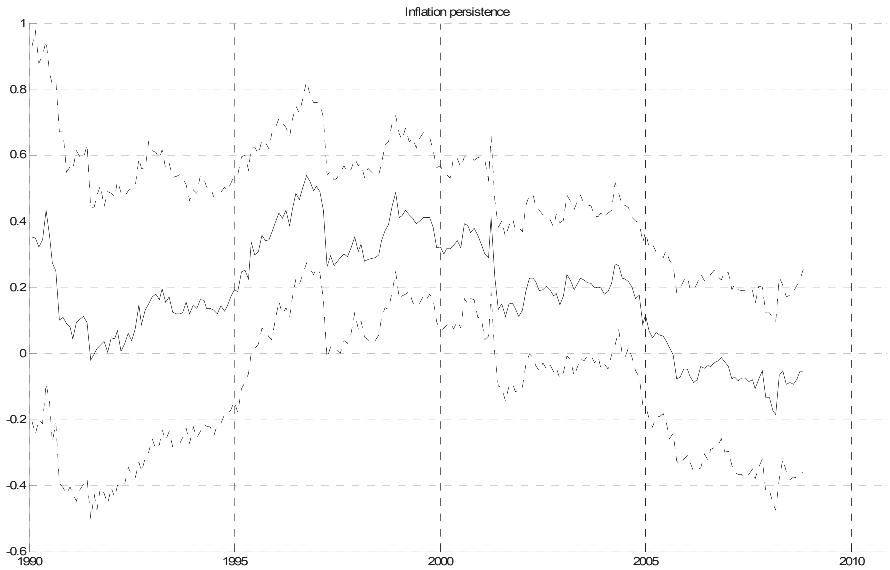


Chart 5 plots inflation persistence, defined as the sum of its autoregressive coefficients. This exhibits a downward sloping trend at the beginning of the third Phase of EMU, which may reflect lower and better-anchored inflation expectations in the post-1999 period, consistently with the ECB mandate 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) argue 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. Inflation persistence



Among recent studies, O'Reilly 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; furthermore, the timing of the breaks generally coincides with observed shifts in the monetary policy regime.

Chart 6 plots steady-state inflation as defined by Equation (8). The chart shows a marked decline in the pre-EMU period followed by stabilisation afterwards.

Chart 6. Steady state inflation

4.3 Short-Run and Steady-State Inflation Uncertainty

Chart 7 shows short-run inflation uncertainty together with its structural component (parameter uncertainty, Equation 5). The sharp decline in the early part of the sample is mainly due to diminishing impulse uncertainty but also to the parameters becoming more stable (which can be interpreted as the effect of monetary policy becoming more predictable). After the start of EMU both components of short-run uncertainty remain low and stable.

Chart 7. Short-run and structural inflation uncertainty

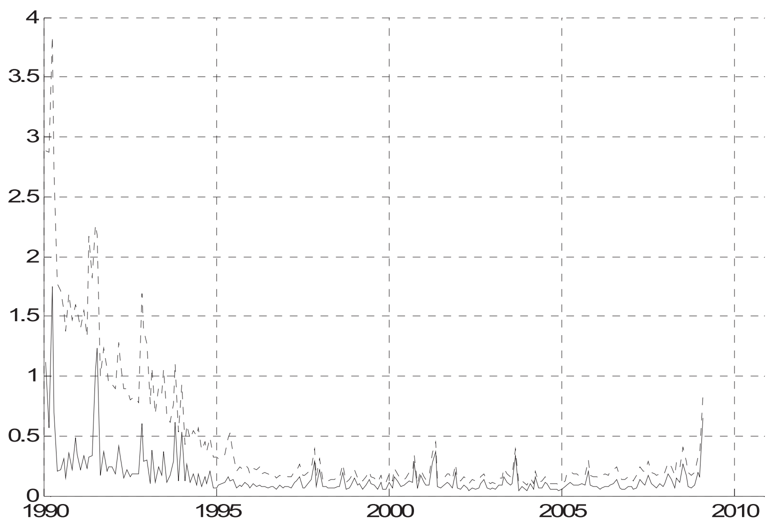


Chart 8. Steady-state inflation uncertainty

Finally, we conducted a robustness check using seasonally adjusted HICP. This series is also 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. Even though regressing seasonally adjusted on seasonally unadjusted series is not appropriate in econometric terms, we do it here for the purpose of checking whether the results are robust to whether or not a seasonal filter is used for the HICP series. As expected, we find that the contemporaneous presence of seasonal adjusted and seasonally unadjusted series biases slightly downwards the estimated inflation persistence, but this effect is compensated by the constant and leads to an almost identical estimate of steady-state inflation. The results in terms of uncertainty are also very similar, with small differences in levels but not in shape. These robustness results are available upon request.

5. 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). The test results support stationarity of both series, are robust to the choice of lag length and indicate that a structural break in structural inflation uncertainty might have occurred at the start of EMU (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	13	-2.77 *	C	-1.94 *	0
EXO_SB	7	-4.86 **	C, 1998 M12	-2.84 *	C, **1998 M12**
END_SB	9	-4.21 **	C, 2008 M4	-2.97 *	C, **2001 M3***

Table 3. Cointegration tests: [infl, infl_ss_var]

Test 1: included lags (levels) 14, intercept – Johansen trace test					
Coint_rank	Test stat.	p_val	90%	95%	99%
r0 = 0	15.68	0.0453	13.42	15.41	19.62
Test 2: included lags (levels) 14, intercept – S&L cointegration trace test					
r0 = 0	17.33	0.0017	8.18	9.84	13.48

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 is estimated.
Standard selection criteria suggest choosing lag length 13 for the bivariate VAR in levels. The deterministic component is specified to include a constant. A shift dummy in 1998m12, included to capture the introduction of the euro, is found to be statistically significant with a sign consistent with diminishing uncertainty. Standard diagnostic tests indicate that the model is statistically adequate.

Granger-causality tests imply uni-directional causality running from inflation to steady-state uncertainty at the 5% confidence level, consistently with the Friedman-Ball hypothesis **(see Table 4)**.

Table 4. Test for Granger-causality

H0: "infl_ss_var" does not Granger-cause "infl" – whole sample	
Test statistic = 1.4666	pval-F(l; 13, 368) = 0.1274
H0: "infl" does not Granger-cause "infl_ss_var" – whole sample	
Test statistic = 1.9398	pval-F(l; 13, 368) = 0.0249

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, the dummy is statistically significant (below -2) between 1997 and 2001.

Chart 10. Sequential t-value of the dummy coefficient

Re-estimating the model over the two sub-samples 1990m1-1998m6 and 2001m3-2009m2 and testing for Granger-causality, we find no evidence of significant causality in the first part of the sub-sample and strong causality from inflation to inflation uncertainty in the second sub-sample (see Table 5). While Granger-causality is only a pre-condition for true causality, our results suggest that the ECB can effectively reduce inflation uncertainty in the euro area by preserving low and stable inflation.

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

	1990m1-1998m6	2001m3-2009m2
Inflation uncertainty does not Granger-cause inflation	0.1946	0.3014
Inflation does not Granger-cause inflation uncertainty	0.3144	0.0014

6. 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 steady-state 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 stabilised both in its impulse and parameter uncertainty component. 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 euro area, 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).

It also appears that in the euro period the Friedman-Ball link is empirically supported, suggesting that the ECB might be able to achieve lower inflation uncertainty by maintaining low and stable inflation. This is consistent with Fountas *et al.* (2004) and Conrad and Karanasos (2005) and with the idea that, given the ECB mandate and its record so far, any long-lasting deviation from price stability in the euro area 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).

This analysis can be extended in several directions. First, 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 euro area, the US, and Canada. Second, estimation of a time-varying VAR-GARCH model could be carried out instead of treating unemployment and other possible variables as exogenous. Third, the forecasting properties of the model could be investigated. Fourth, survey data could also be used to compute inflation uncertainty and these results compared to those obtained using our framework.

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Data appendix

The following series are used in the baseline estimation:

- Euro area (changing composition) - HICP - Overall index, Monthly Index, Eurostat, neither seasonally nor working day adjusted”. 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.
- Euro area (changing composition) - Standardised unemployment rate, Total (all ages), Total (male & female), Eurostat, neither seasonally or working day adjusted, percentage of civilian workforce”. It is linked backwards 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”.
- Euro area (changing composition) - Money Market - Euribor 3-month - Historical close - Euro, provided by Reuters. Euro area (changing composition) - Benchmark bond - Euro area 10-year Government Benchmark bond yield - Yield - Euro
- Nominal effective exch. rate, 12 group of currencies (AU, CA, DK, HK, JP, NO, SG, KR, SE, CH, GB, US), changing composition of the euro area against Euro
- Euro area (changing composition), Index of Notional Stocks, MFIs, central government and post office giro institutions reporting sector - Monetary aggregate M3, All currencies combined - Euro area (changing composition) counterpart, Non-MFIs excluding central government sector, Annual growth rate, data Working day and seasonally adjusted,