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## The Euro and inflation uncertainty in the European Monetary Union

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This paper adopts a time-varying GARCH framework to estimate short-run and steady-state inflation uncertainty in 12 EMU countries, and then investigates their relationship with inflation. The effects of the Euro introduction in 1999 are examined by utilising a dummy variable. Tests for endogenously determined breaks are also employed. We find a considerable degree of heterogeneity across EMU countries in terms of average inflation, its degree of persistence, and both types of uncertainty, whilst the trend component of inflation is generally decreasing. Various breaks in the relationship between inflation and inflation uncertainty are found, frequently well before the Euro introduction.

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### 1. Introduction

The introduction of the Euro and of a common monetary policy in 1999 undoubtedly represented a major policy regime shift for the member countries of the European Monetary Union (EMU). This could have affected both inflation expectations and inflation uncertainty, as, at least initially, agents might not have been certain of the objective function and the policy preferences of the European Central Bank (ECB), and of how they might compare to those of the national central banks previously in charge of monetary policy (for instance, the ECB might have been perceived as less credible than the Bundesbank, which had an established anti-inflation reputation). Uncertainty about the policy

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preferences of the new monetary authorities might also result in higher inflation forecast errors. According to the Maastricht Treaty, although the primary objective of the ECB is price stability (which the ECB has interpreted as an annual Euro area inflation rate below, but close to, 2% in the medium run), it should also be concerned with output and employment (albeit without prejudicing its main objective). The monetary policy framework adopted by the ECB to fulfil these tasks is based on two analytical perspectives or two “pillars”, namely economic analysis and monetary analysis,<sup>1</sup> and the ECB has repeatedly stated that achieving price stability is the most effective way to contribute to output and employment growth, but nevertheless higher uncertainty might have characterised the new economic environment.

Analysing survey data, [Heinemann and Ullrich \(2006\)](#) do not find significant differences in the inflationary credibility of the ECB compared to the Bundesbank, and hence no permanent change in inflation expectations. However, their analysis suggests that the higher uncertainty characterising the period leading up to EMU led to a temporary change in expectation formation, with agents relying more heavily on backward-looking expectations, before reverting to the normal mechanisms once the ECB had established its inflation credibility.

As for inflation uncertainty, in a recent review of the performance of the ECB in the first few years of the new regime, its President, Jean-Claude Trichet, has expressed the view that “... the ECB has, despite substantial adverse price shocks, successfully kept inflation and inflation expectations at low levels by historical standards. The single monetary policy and its clear focus on the maintenance of price stability have helped to anchor inflation expectations in the Euro area over the medium and the long-term. This has facilitated a reduction of inflation uncertainty and the associated risk premia” (see [Trichet, 2004](#), p. 2).

In this paper, we adopt an appropriate econometric framework to analyse empirically whether the new policy regime with the ECB setting a common interest rate for the EMU countries has in fact different features, in particular whether the link between inflation and inflation uncertainty has changed. Specifically, we use a time-varying model with a GARCH specification for the conditional volatility of inflation, as in [Evans \(1991\)](#), and obtain estimates for 12 EMU countries, over the period 1973–2004, using monthly data. The adopted framework enables us to distinguish between different types of inflation uncertainty which can affect the inflation process. Next, we estimate the relationship between inflation and inflation uncertainty taking into account the possibility of breaks. Dummy variables corresponding to the introduction of the Euro are initially incorporated into the model. However, as the mere announcement of a regime switching from floating to fixed rates at a given future date can determine changes in the behaviour of rational agents prior to the fixing, we also determine endogenously the break dates using a procedure developed by [Bai and Perron \(1998, 2003\)](#). This allows us to investigate whether adjustment took place much before the introduction of the Euro. This type of analysis is motivated by some theoretical literature demonstrating that rational agents will react to the announcement of a regime switch from floating to fixed rates well before the change occurs (see [Wilfling, 2004](#); [Wilfling and Maennig, 2001](#)).

Our empirical findings enable us to shed light on the difficulties encountered by the ECB in fulfilling its mission in the new environment resulting from the introduction of the Euro. In particular, we find that in the post-Euro period there are still significant inflation differentials across countries, implying different real interest rates for the various EMU members given a single nominal interest rate. Further, heterogeneity in both short-run and steady-state uncertainty occurs even in the presence of a common currency: although the former is only to be expected given the well-known lags of monetary policy, the latter clearly makes the ECB's policy objective of long-run price stability hard to achieve. Moreover, the relationship between inflation and inflation uncertainty

<sup>1</sup> Economic analysis aims at assessing the short- to medium-term determinants of price developments focusing on real activity and financial conditions in the economy. Monetary analysis focuses on a longer-term horizon taking into account the long-run relationship between money and prices. A reference value of 4.5% for the growth rate of broad money (M3) that is compatible with price stability has been calculated using the quantity theory equation. The ECB has stated, though that “monetary policy does not react mechanically to deviations of M3 growth from the reference value” (see [The Monetary Policy of the ECB, 2004](#)). As [Rudebusch and Svensson \(1999, p. 1\)](#) point out, the ECB strategy “appears to be a combination of a weak type of monetary targeting and an implicit form of inflation targeting”.

appears to have broken down in a number of cases, implying that the ECB faces an even harder task in keeping inflation under control. Under the circumstances, the ECB can be deemed to have performed at least satisfactorily, but it is possible that some suggested changes to its analytical framework might make its policies even more effective.

The layout of the paper is the following. Section 2 reviews the relevant literature. Section 3 outlines the empirical framework. Section 4 presents the empirical results. Section 5 summarises the main findings and discusses their policy implications.

## 2. A brief literature review

The relationship between inflation and inflation uncertainty has received increased attention in recent years. Friedman (1977) first argued that higher average inflation would result in more inflation uncertainty. This idea was developed by Ball (1992). In his model, in the presence of two types of policymakers with different preferences, who stochastically alternate in power, higher inflation generates higher inflation uncertainty, as agents do not know when monetary authorities with a tougher stance on inflation will replace the current ones. In contrast to the Friedman–Ball hypothesis, the effect of inflation of its uncertainty can also be negative. As Pourgerami and Maskus (1987) point out, in an environment of accelerating inflation agents may invest more resources in forecasting inflation, thus reducing uncertainty about inflation (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 set up, such as the one due to Cukierman and Meltzer (1986), in which there is an incentive for policymakers to create inflation surprises to raise output growth.

A number of empirical studies have investigated the relationship between inflation and inflation uncertainty, typically adopting an econometric framework of the GARCH type to measure uncertainty (see Engle, 1982), and providing mixed evidence (see Davis and Kanago, 2000 for a survey). Most studies, such as Grier and Perry (1998), adopt a two-step procedure: they estimate GARCH models to generate a measure of inflation uncertainty, and then carry out Granger causality tests (see also Fountas et al., 2004).<sup>2</sup> Using data for the G7 countries, they find strong evidence of causality running from inflation uncertainty to inflation, but less empirical support for causality in the opposite direction (see also Baillie et al., 1996). Various studies focus on the US, again with mixed results. Brunner and Hess (1993) and Grier and Perry (1998, 2000), *inter alia*, find evidence of a Friedman effect, with Baillie et al. (1996) reporting the opposite. More recently, the impact of inflation targeting on this relationship has been analysed. Kontonikas (2004) reports that the adoption of an explicit target in the UK has resulted in lower inflation persistence and long-run uncertainty.

Fountas et al. (2004) argue that in the context of EMU the linkages between inflation, inflation uncertainty and output growth have even more important implications for monetary policy, since price stability becomes an even more crucial policy objective for the ECB if inflation is found to affect inflation uncertainty. Further, asymmetries in the effects of inflation uncertainty on output across member countries could make a common monetary policy a less effective stabilisation tool. In fact their empirical analysis, based on EGARCH models, provides evidence supporting the Friedman hypothesis and the presence of asymmetric real effects. However, their sample period is 1960–1999, and hence does not include the new monetary policy setting resulting from the introduction of the Euro, whose effects on inflation we wish to examine. Further, their analysis does not distinguish between different types of inflation uncertainty, whilst the approach taken in the present study, as explained below,

<sup>2</sup> Note that Pagan (1984) criticizes this two-step procedure for its misspecifications due to the use of generated variables from the first stage as regressors in the second stage. According to this line of reasoning, the simultaneous conditional mean and variance estimation as in a GARCH-in-mean (GARCH-M) model is more efficient than a two-step approach. However, as pointed out by Grier and Perry (1998) and Fountas et al. (2004), Pagan's approach has the drawback that it does not allow testing possible lagged effects of inflation uncertainty on inflation, which might exist at the monthly or quarterly frequency. Fountas et al. (2004) report the estimation results of an EGARCH-M model, which confirm that a simultaneous approach does not detect the causal effect of inflation uncertainty on inflation. For these reasons, following Grier and Perry (1998) we also adopt the two-step procedure.

enables us to measure separately the impact of short-run (structural and impulse) and long-run uncertainty.<sup>3</sup>

### 3. Econometric framework

All the methods discussed above have the drawback that they do not take into account the fact that uncertainty about the long- and short-term prospects for inflation might differ significantly 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, 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), reflecting, for example, changes in the variance of structural disturbances such as price shocks (see Berument et al., 2005). Apart from Evans (1991), other studies that attempt to decompose US inflation uncertainty into its long-run and short-run components include Ball and Cecchetti (1990), Evans and Wachtel (1993), and Guler and Ozlale (2005).<sup>4</sup>

The econometric framework suggested by Evans (1991), and also adopted by Berument et al. (2005) in their analysis of the linkages between UK inflation uncertainty and interest rates, has the advantage over alternative approaches of yielding estimates of the various types of uncertainty discussed above. Following Evans (1991), in the present study we utilise a GARCH model with time-varying parameters, which are estimated using Kalman filtering, for Euro area inflation rates. More specifically, inflation is specified as a  $k$ -th order autoregressive process, AR( $k$ ), with time-varying parameters, the residuals of this equation following a GARCH(1,1) process. The model is the following:

$$\pi_{t+1} = \mathbf{X}_t \boldsymbol{\beta}_{t+1} + e_{t+1} \text{ where } e_{t+1} \sim N(0, h_t) \text{ and } \mathbf{X}_t = [1, \pi_t, \dots, \pi_{t-k}] \quad (1)$$

$$h_t = h + \alpha e_{t-1}^2 + \lambda h_{t-1} \quad (2)$$

$$\boldsymbol{\beta}_{t+1} = \boldsymbol{\beta}_t + \mathbf{V}_{t+1} \text{ where } \mathbf{V}_{t+1} \sim N(\mathbf{0}, \mathbf{Q}) \quad (3)$$

where  $\pi_{t+1}$  denotes the rate of inflation between  $t$  and  $t+1$ ;  $\mathbf{X}_t$  is a vector of explanatory variables known at time  $t$ ;  $e_{t+1}$  describes the shocks to the inflation process that cannot be forecast with information known at time  $t$ ;  $e_{t+1}$  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,  $e_{t-i}^2$ , and past variances,  $h_{t-j}$ . Further,

<sup>3</sup> Other strands of the literature analyse the relationship between inflation and its uncertainty using long-memory models (see Conrad and Karanasos, 2006), and possible asymmetries (see Brunner and Hess, 1993): examples are the EGARCH model of Nelson (1991) used in Fountas et al. (2004); the Threshold GARCH (TGARCH) model of Zakoian (1994), and the component GARCH (CGARCH) model of Engle and Lee (1993) (both these models are estimated by Grier and Perry, 1998 and Kontonikas, 2004).

<sup>4</sup> Evans and Wachtel (1993) stress that the assumption of fixed parameters in the inflation process overestimates the degree to which agents can forecast inflation, and consequently underestimates inflation uncertainty. They decompose the sources of inflation uncertainty into two components: “regime uncertainty component” and “certainty equivalence component”. The second component ignores uncertainty about future inflation regimes and reflects only the variance of future shocks to the inflation process. The first component reflects the agents' uncertainty about the characteristics of the current policy regime or even future regimes, if there is a possibility that the regime will change. Thus, cross-countries differences in the conduct of monetary policy may account for the differences in the average levels of uncertainty. This decomposition allows inflation uncertainty to change over time as agents keep updating their information on the current regime and their expectations about the future regime. See also the comment by Brunner (1993).

$\beta_{t+1} = [\beta_{0,t+1}, \beta_{1,t+1}, \dots, \beta_{k,t+1}]'$  denotes the time-varying parameter vector,<sup>5</sup> and  $\mathbf{V}_{t+1}$  is a vector of shocks to  $\beta_{t+1}$ , assumed to be normally distributed with a homoscedastic covariance matrix  $\mathbf{Q}$ . The updating equations for the Kalman filter are

$$\pi_{t+1} = \mathbf{X}_t E_t \beta_{t+1} + \varepsilon_{t+1} \tag{4}$$

$$H_t = \mathbf{X}_t \Omega_{t+1|t} \mathbf{X}'_t + h_t \tag{5}$$

$$E_{t+1} \beta_{t+2} = E_t \beta_{t+1} + [\Omega_{t+1|t} \mathbf{X}'_t H_{t-1}] \varepsilon_{t+1} \tag{6}$$

$$\Omega_{t+2|t+1} = [\mathbf{I} - \Omega_{t+1|t} \mathbf{X}'_t H_{t-1} \mathbf{X}_t] \Omega_{t+1|t} + \mathbf{Q} \tag{7}$$

where  $\Omega_{t+1|t}$  is the conditional variance/ covariance matrix of  $\beta_{t+1}$  given the information set at time  $t$ , representing uncertainty about the structure of the inflation process.

As Eq. (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  $\mathbf{X}_t \beta_{t+1}$ , which is  $\mathbf{X}_t \Omega_{t+1|t} \mathbf{X}'_t = S_t$  (structural uncertainty). The standard GARCH model can be obtained as a special case of our model if there is no uncertainty about  $\beta_{t+1}$ , so that  $\Omega_{t+1|t} = \mathbf{0}$ . In this case, the conditional variance of inflation depends solely on impulse uncertainty.<sup>6</sup> Eqs. (6) and (7) capture the updating of the conditional distribution of  $\beta_{t+1}$  over time in response to new information about realised inflation. As indicated by Eq. (6), inflation innovations, defined as  $\varepsilon_{t+1}$  in Eq. (4), are used to update the estimates of  $\beta_{t+1}$ . These estimates are then used to forecast future inflation.

If there are no inflation shocks and parameter shocks, so that  $\pi_{t+1} = \pi_t = \dots = \pi_{t-k}$  for all  $t$ , we can calculate the steady-state rate of inflation,  $\pi_{t+1}^*$ , as:

$$\pi_{t+1}^* = \beta_{0,t+1} \left[ 1 - \sum_{i=1}^k \beta_{i,t+1} \right]^{-1} \tag{8}$$

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

$$\sigma_t^2(\pi_{t+1}^*) = \nabla E_t \beta_{t+1} \Omega_{t+1|t} \nabla E_t \beta'_{t+1} \tag{9}$$

$$\text{where } \nabla E_t \beta'_{t+1} = \begin{bmatrix} \left( 1 - \sum_{i=1}^k E_t \beta_{i,t+1} \right)^{-1} \\ E_t \beta_{0,t+1} \left( 1 - \sum_{i=1}^k E_t \beta_{i,t+1} \right)^{-2} \\ \dots \\ E_t \beta_{0,t+1} \left( 1 - \sum_{i=1}^k E_t \beta_{i,t+1} \right)^{-2} \end{bmatrix} \text{ is a } (k + 1 \times 1) \text{ vector.} \tag{10}$$

<sup>5</sup> Evans (1991) provides a theoretical justification for the random walk specification of the time-varying parameter vector. He argues (p. 176): "Suppose, for example, that all the structural variations in inflation reflect changes in monetary policy which in turn are due to changing views about the structure of the economy. Since it would be very hard to predict any future change in policy and hence movements in  $\beta_t$  under these conditions,  $E_t \beta_{t+1} = E_t \beta_t$  as implied by a random walk for  $\beta_t$ ."

<sup>6</sup> As Evans (1991) argues, if there is uncertainty about  $\beta_{t+1}$ ,  $h_t$  will tend to understate the true conditional variance since  $S_t > 0$ .

Having computed short-run and steady-state uncertainty measures for each country, we then proceed, in the second part of our empirical investigation, to analyse the links between the various types of inflation uncertainty and the level of inflation, and to examine the impact of the Euro. Specifically, we regress the two uncertainty measures against past inflation.<sup>7</sup> Moreover, we include a dummy variable to allow for possible intercept and slope changes in the underlying relationship between inflation uncertainty and past inflation reflecting the introduction of the Euro. The estimated model is the following:

$$\text{unc}_{t+1} = \gamma_0 + \gamma_1 D_{t+1} + (\gamma_2 + \gamma_3 D_{t+1}) \pi_t + \theta_{t+1} \quad (11)$$

where  $\text{unc}_{t+1}$  represents in turn steady-state uncertainty (i.e.  $\sigma_t^2(\pi_{t+1}^*)$ ) and short-run uncertainty (i.e.,  $H_t$ ), and  $D_{t+1}$  is a dummy variable equal to zero during the pre-Euro period and one during the Euro period.

In the model specified above, the possible structural break in the relationship between inflation and inflation uncertainty in the Euro area is exogenously fixed at January 1999. However, the mere announcement of a regime switching from floating to fixed rates could have induced changes in the behaviour of rational agents and thereby could have affected the inflation-uncertainty relationship prior to the fixing in 1999 (see Wilfling, 2004; Wilfling and Maennig, 2001). Hence, we also apply the procedure developed by Bai and Perron (1998, 2003) for multiple structural change models, which enables one to determine endogenously the number of breaks and the break dates. The procedure considers all possible models under the assumption of a given number of breaks and a given minimum distance between the break dates. The selected “optimal” model is then the one which minimises the sum of squared residuals and some information criteria. In our application we allow for up to three possible breaks, and use the Bayesian Information Criterion (BIC) to choose the best specification.<sup>8</sup>

## 4. Empirical analysis

### 4.1. Data

Inflation is measured as the first difference of the logarithm of the seasonally adjusted consumer price index (CPI),  $\pi_{t+1} = 100(\ln \text{CPI}_{t+1} - \ln \text{CPI}_t)$ , using monthly data for 12 EMU countries (Germany, France, Italy, Spain, Portugal, Greece, Ireland, Finland, Belgium, Netherlands, Luxembourg, Austria) over the period 1973–2004. Six years of the Euro period are included in our sample,<sup>9</sup> allowing us to study the effects of the 1999 policy regime shift on inflation uncertainty over a reasonably long horizon. The data are obtained from OECD’s *Main Economic Indicators: Historical Statistics*.

### 4.2. Unit root test results

Establishing the order of integration of inflation is important for our subsequent empirical analysis. In columns 2–5 of Table 1 we report the results from ADF (see Dickey and Fuller, 1979, 1981) and KPSS (see Kwiatkowski et al., 1992) unit root tests with an intercept and a deterministic linear trend. Overall, the results suggest that inflation in our sample countries has a unit root. This is consistent with the findings of Rapach and Weber (2004), who also conclude that inflation is non-stationary using a sample of OECD countries and the Ng–Perron unit root test (see Ng and Perron, 2001).

<sup>7</sup> In contrast to Evans (1991, p. 180) where “the regressions use the month-to-month changes in the variances and inflation because inflation has a unit root and all three variances are complicated functions of past inflation”, we utilise the *levels* of the series since inflation was found to be stationary when structural breaks were taken into account.

<sup>8</sup> An alternative, sequential procedure is also discussed by Bai and Perron (2003).

<sup>9</sup> As Greece adopted the Euro only in January 2001, the corresponding sub-sample is four years.

**Table 1**

Unit root tests, 1972–2004.

Country	ADF		KPSS		Lee and Strazicich						
	Constant	Constant and trend	Constant	Constant and trend	Model A – breaks in constant			Model C – breaks in constant and trend			
					$\hat{\tau}$ -stat	Break dates		$\hat{\tau}$ -stat	Break dates		
Germany	-2.319	-2.605	2.235***	0.244***	-4.06**	1990.10	1991.09		-8.52***	1991.09	2000.12
Italy	-1.144	-3.375*	1.999***	0.273***	-5.92***	1984.01	1984.06		-7.62***	1975.12	1984.03
France	-1.461	-2.896	1.646***	0.385***	-3.91**	1983.04	1988.08		-6.6***	1983.06	1999.12
Spain	-1.006	-2.751	3.222***	0.451***	-2.76	1975.05	1976.04		-10.96***	1976.04	1977.01
Portugal	-1.579	-3.568**	2.634***	0.277***	-5.66***	1976.07	1976.11		-11.84***	1976.03	1977.01
Greece	-2.006	-2.704	1.845***	0.34***	-6.15***	1978.09	1993.11		-10.3***	1975.08	1979.01
Ireland	-1.595	-2.539	2.232***	0.262***	-3.25	1975.04	1975.12		-11.3***	1980.03	1981.10
Finland	-1.181	-3.132	2.037***	0.222***	-4.67***	1976.11	1983.03		-8.44***	1975.12	1982.05
Belgium	-1.929	-2.765	2.121***	0.265***	-6.13***	1975.09	1976.09		-8.5***	1976.01	1985.05
Netherlands	-1.787	-1.965	2.578***	0.864***	-3.32	1977.04	1978.06		-9.19***	1976.03	1991.07
Luxembourg	-2.438	-2.523	3.195***	0.304***	-5.01***	1979.08	1984.10		-8.85***	1998.12	2001.01
Austria	-2.369	-2.684	3.282***	0.301***	-3.16	1984.01	1984.07		-8.27***	1976.07	1984.01

Note: The number of lagged difference terms in the regressions was chosen by the modified Akaike criterion in the ADF test. The Andrews bandwidth was used in the KPSS test. In the Lee–Strazicich test, the number of lagged difference terms was chosen by the 't-sig' approach suggested by Perron (1997). We set an upper bound of 12 for the lag length and test down until a significant (at the 10% level) lag is found. The reported ADF and Lee and Strazicich statistics test the null hypothesis that inflation contains a unit root. The reported KPSS statistics test the null hypothesis that inflation is stationary. The following asterisks \*\*\*, \*\*, \* indicate rejection of the null hypothesis at the 1, 5, 10% level of significance, respectively.

A potential shortcoming of the ADF unit root test is that a stationary variable that is subject to structural breaks may appear to be non-stationary. Since Perron (1989), it has been recognised that ignoring an existing structural break results in a greater tendency to under-reject the null of unit root when the stationary alternative is true.<sup>10</sup> Perron's (1989) initial approach was to allow for a single exogenously imposed structural break under both the null and alternative hypotheses. Subsequent studies have emphasised the need to determine the break endogenously from the data (see, e.g., Perron, 1997). Given the relatively long span of our inflation series, in order to account for the possibility of more than one structural break, we utilise the endogenous two-break unit root test of Lee and Strazicich (2003). This test counterbalances the potential loss of power of tests that overlook the possibility of more than one break. Unlike the Lumsdaine and Papell (1997) two-break unit root test,<sup>11</sup> the Lee and Strazicich (2003) test includes breaks under both the null and the alternative hypotheses, with rejections of the null unambiguously implying trend stationarity. Two models are considered, one that allows for two breaks in the level of inflation (Model A), and one that allows for two breaks in the intercept and the trend of inflation (Model C).

In columns 6–11 of Table 1 we report the results from the Lee and Strazicich (2003) two-break unit root test. Allowing for two shifts in the intercept and the slope of inflation provides greater support for inflation stationarity, as compared to the model that allows only for level breaks, since unit root rejection rates increase from 8/12 to 12/12 when we switch from Model A to Model C. Thus, inflation can be treated as a stationary variable in all sample countries when breaks in the intercept and the slope of the series are taken into account. This finding is very important since it allows us to proceed further and extract steady-state inflation and the corresponding uncertainty from the inflation series, something which could not be conceptualised if there was no mean reversion and the variance of inflation was explosive.<sup>12</sup> Regarding the dates of the breaks, Model C indicates that at least one of the

<sup>10</sup> Clark (2006) and Levin and Piger (2003) among others allow for the possibility of structural breaks when examining inflation persistence. Rapach and Wohar (2005) also test for breaks and find pervasive evidence of shifts in the level of inflation in a wide range of European countries.

<sup>11</sup> The null hypothesis in the endogenous two-break unit root test of Lumsdaine and Papell (1997) assumes no structural breaks, while the alternative does not necessarily imply broken trend stationarity. Thus, rejecting the null may be interpreted as rejection of a unit root with no structural break, and not necessarily as rejection of a unit root *per se*.

<sup>12</sup> We would like to thank a referee for raising this point.

breaks takes place around the mid and late-1970s, a period of high and volatile inflation. In general, breaks tend to occur prior to 1999 with only a few instances (Germany, France and Luxembourg) of a break around the introduction of the Euro, or during the post-Euro period.

#### 4.3. Time-varying GARCH model estimates

We have estimated a time-varying GARCH model for inflation with Kalman filtering, as described in Section 3. Table 2 reports the preferred specifications. The  $t$ -tests of significance applied to estimates of the main diagonal of  $\mathbf{Q}$  indicate that overall there is time variation in the parameters of the model. In each country, at least one of the parameters' variance estimates ( $\sigma_i$ ) is significantly different from zero. Statistically significant time variation appears to be present either in the intercept of the inflation model (e.g. Ireland), or in some of the lags (e.g. France), or in both intercept and lags (e.g. Germany). The ARCH ( $\alpha$ ) and GARCH ( $\lambda$ ) parameter estimates are positive and statistically significant in most cases. In a number of countries (Italy, Spain, Portugal, Greece, Ireland, and Finland) the sum of the ARCH and GARCH coefficients almost equals one, indicating that volatility shocks are very persistent. Ljung–Box test statistics of the squared standardised residuals at lags one to six<sup>13</sup> show that, with the exception of a few lags in a small minority of countries, there are no remaining GARCH effects at the 5% level of significance. Hence, overall, Table 2 indicates that there is time variation in the parameters of the inflation model and that the GARCH(1,1) conditional variance specification captures the dynamics of inflation volatility.

Figs. 1–3 are based on the estimation results. Fig. 1 plots actual inflation and steady-state inflation in the EMU countries over the period 1980.01–2004.11. In the early years of the new monetary regime the Euro area was affected by a variety of price shocks such as the tripling of oil prices between early 1999 and mid-2000, the depreciation of the common currency over this period, and finally, in 2001, significant increases in food prices, due to a series of livestock epidemics. This is evident across the EMU countries in the plots of actual inflation. Average monthly inflation rates vary considerably in the EMU area, ranging from 0.2% in Germany to 1% in Greece. Similarly to the former country, mean monthly inflation rates in the Benelux countries (Belgium, Netherlands, Luxembourg) and Austria were low: 0.26%, 0.21%, 0.26% and 0.23%, respectively. Steady-state inflation follows similar patterns. The Club-Med countries were the worst performers in terms of annualised steady-state inflation rate: Greece 12%, Portugal 9.8%, Spain 6.6%, Italy 6.4%, while the corresponding value for Germany, Austria and the Benelux countries was around 3%. Ireland, France, and Finland fall somewhere in the middle with annualised steady state inflation rates of 5.2%, 4.5%, and 4.2%, respectively.

Busetti et al. (2006) also present evidence of diverging behaviour in the inflation rate of the EMU countries since 1999. Such inflation differentials are often found even within monetary unions, where many economic differences may survive. The ECB itself admits that monetary policy can only affect price developments in the Euro area as a whole and cannot influence inflation differentials across regions (see *The Monetary Policy of the ECB, 2004*). Nevertheless, from the viewpoint of monetary policy effectiveness in stimulating economic growth, inflation rates in EMU countries should converge in order for changes in the Euro-wide nominal interest rate to be translated into similar real interest rate changes across member countries.

Fig. 2 plots short-run uncertainty and steady-state uncertainty.<sup>14</sup> The former appears to have decreased over time in Italy, Spain, Portugal, Greece, Ireland, Finland and Belgium. In Germany a large temporary increase in short-run uncertainty can be noticed around the time of the re-unification in the early 1990s. Short-run uncertainty in the Netherlands and Austria is relatively stable, apart from occasional temporary shocks. The same applies to Luxembourg, with the exception of a large

<sup>13</sup> The performance of the Ljung–Box test is affected by the number of lags ( $k$ ) that are utilised. Tsay (2002) suggests that the choice of  $k = \ln(\text{sample size})$  provides better power performance. Thereby, we set an upper limit for  $k$  equal to  $\ln(\text{sample size}) \approx 6$ .

<sup>14</sup> Tests for the equality of mean and variance between short-run and steady-state uncertainty indicate that there are statistically significant differences between the two series in our sample countries. Results are not included to save space, but are available from the authors upon request.



**Table 2**

Kalman filter estimates, 1972–2004.

Country	AR specification and coefficient variance estimates	GARCH estimates			Squared standardized residuals diagnostics					
		$h$	$\alpha$	$\lambda$	LB(1)	LB(2)	LB(3)	LB(4)	LB(5)	LB(6)
Germany	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_{t-1} + e_{t+1}$ $\sigma_0 = 0.025^{***}$ $\sigma_1 = 0.019^{**}$	0.017 <sup>***</sup>	0.244 <sup>**</sup>	0.464 <sup>***</sup>	0.117 [0.732]	0.595 [0.743]	3.041 [0.385]	3.041 [0.551]	3.863 [0.569]	4.23 [0.646]
Italy	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_t + \beta_{2,t+1}\pi_{t-2} + \beta_{2,t+1}\pi_{t-6} + \beta_{2,t+1}\pi_{t-11} + e_{t+1}$ $\sigma_0 = 0.03^{***}$ $\sigma_1 = 0.005$ $\sigma_2 = 0.016$ $\sigma_3 = 0.002$ $\sigma_4 = 0.024^{***}$	0.0001	0.105 <sup>***</sup>	0.8947 <sup>***</sup>	1.585 [0.208]	1.92 [0.338]	8.782 [0.032]	8.827 [0.065]	9.972 [0.076]	10.652 [0.099]
France	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_t + \beta_{2,t+1}\pi_{t-7} + e_{t+1}$ $\sigma_0 = 0.007$ $\sigma_1 = 0.049^*$ $\sigma_2 = 0.051$	0.247	0.356	0.299 <sup>***</sup>	0.038 [0.846]	2.817 [0.245]	3.047 [0.385]	3.055 [0.549]	3.131 [0.68]	5.978 [0.426]
Spain	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_t + \beta_{2,t+1}\pi_{t-2} + \beta_{2,t+1}\pi_{t-7} + \beta_{2,t+1}\pi_{t-11} + e_{t+1}$ $\sigma_0 = 0.023^{***}$ $\sigma_1 = 0.00$ $\sigma_2 = 0.021^{**}$ $\sigma_3 = 0.008$ $\sigma_4 = 0.018^{***}$	0.001	0.108 <sup>***</sup>	0.891 <sup>***</sup>	0.186 [0.666]	2.213 [0.331]	2.316 [0.51]	3.796 [0.434]	11.798 [0.038]	12.338 [0.055]
Portugal	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_t + \beta_{2,t+1}\pi_{t-2} + e_{t+1}$ $\sigma_0 = 0.039^{***}$ $\sigma_1 = 0.031^{***}$ $\sigma_2 = 0.038^{***}$	0.006 <sup>*</sup>	0.121 <sup>***</sup>	0.863 <sup>***</sup>	3.518 [0.061]	3.862 [0.145]	4.358 [0.225]	6.394 [0.172]	10.18 [0.07]	13.418 [0.037]
Greece	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_{t-1} + \beta_{2,t+1}\pi_{t-2} + \beta_{2,t+1}\pi_{t-6} + \beta_{2,t+1}\pi_{t-7} + e_{t+1}$ $\sigma_0 = 0.038^{***}$ $\sigma_1 = 0.033^{***}$ $\sigma_2 = 0.00$ $\sigma_3 = 0.005$ $\sigma_4 = 0.00$	0.001	0.062 <sup>***</sup>	0.9378 <sup>***</sup>	0.636 [0.425]	5.624 [0.06]	6.327 [0.097]	7.033 [0.134]	8.688 [0.122]	11.096 [0.086]
Ireland	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_t + \beta_{2,t+1}\pi_{t-3} + e_{t+1}$ $\sigma_0 = 0.087^{***}$ $\sigma_1 = 0.019$ $\sigma_2 = 0.018$	0.307 <sup>***</sup>	0.989 <sup>***</sup>	0.000	1.606 [0.205]	1.920 [0.383]	2.216 [0.529]	2.880 [0.578]	3.322 [0.651]	12.036 [0.061]
Finland	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_{t-5} + \beta_{2,t+1}\pi_{t-8} + e_{t+1}$ $\sigma_0 = 0.021^{***}$ $\sigma_1 = 0.022^*$ $\sigma_2 = 0.025^*$	0.001 <sup>***</sup>	0.0625 <sup>**</sup>	0.9366 <sup>***</sup>	0.078 [0.78]	0.522 [0.77]	0.953 [0.813]	0.966 [0.915]	6.292 [0.279]	9.898 [0.129]
Belgium	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_t + \beta_{2,t+1}\pi_{t-3} + \beta_{2,t+1}\pi_{t-5} + \beta_{2,t+1}\pi_{t-14} + e_{t+1}$ $\sigma_0 = 0.008$ $\sigma_1 = 0.012$ $\sigma_2 = 0.021^{**}$ $\sigma_3 = 0.002$ $\sigma_4 = 0.014$	0.004	0.079 <sup>**</sup>	0.874 <sup>***</sup>	1.232 [0.267]	5.010 [0.082]	6.667 [0.083]	8.311 [0.081]	8.312 [0.14]	9.905 [0.129]
Netherlands	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_{t-3} + \beta_{2,t+1}\pi_{t-11} + e_{t+1}$ $\sigma_0 = 0.022^{***}$ $\sigma_1 = 0.00$ $\sigma_2 = 0.012$	0.011 <sup>**</sup>	0.169 <sup>**</sup>	0.666 <sup>***</sup>	0.823 [0.364]	1.458 [0.483]	2.448 [0.485]	3.836 [0.429]	4.134 [0.53]	15.056 [0.02]
Luxembourg	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_t + \beta_{2,t+1}\pi_{t-4} + \beta_{2,t+1}\pi_{t-5} + \beta_{2,t+1}\pi_{t-10} + e_{t+1}$ $\sigma_0 = 0.025^{**}$ $\sigma_1 = 0.011$ $\sigma_2 = 0.00$ $\sigma_3 = 0.031^{**}$ $\sigma_4 = 0.00$	0.014 <sup>**</sup>	0.081 <sup>**</sup>	0.751 <sup>***</sup>	0.147 [0.701]	0.274 [0.872]	0.354 [0.95]	0.413 [0.981]	0.422 [0.995]	1.026 [0.985]
Austria	$\pi_{t+1} = \beta_{0,t+1} + \beta_{1,t+1}\pi_{t-5} + \beta_{2,t+1}\pi_{t-11} + e_{t+1}$ $\sigma_0 = 0.019^{***}$ $\sigma_1 = 0.015$ $\sigma_2 = 0.044$	0.012 <sup>**</sup>	0.066	0.774 <sup>***</sup>	0.183 [0.669]	0.692 [0.708]	0.697 [0.874]	0.859 [0.93]	1.036 [0.96]	2.74 [0.841]

Note: the reported  $\sigma_i$ 's are estimates of the parameters' variance (main diagonal elements of the Q matrix). LB(k) indicates the Ljung–Box test statistic for squared standardized residuals' ( $\hat{\epsilon}_t/h_t$ ) serial correlation at lag k, and the figures in square brackets represent the associated p-values. The following asterisks \*\*\*, \*\*, \* indicate the 1, 5, 10% level of significance, respectively.

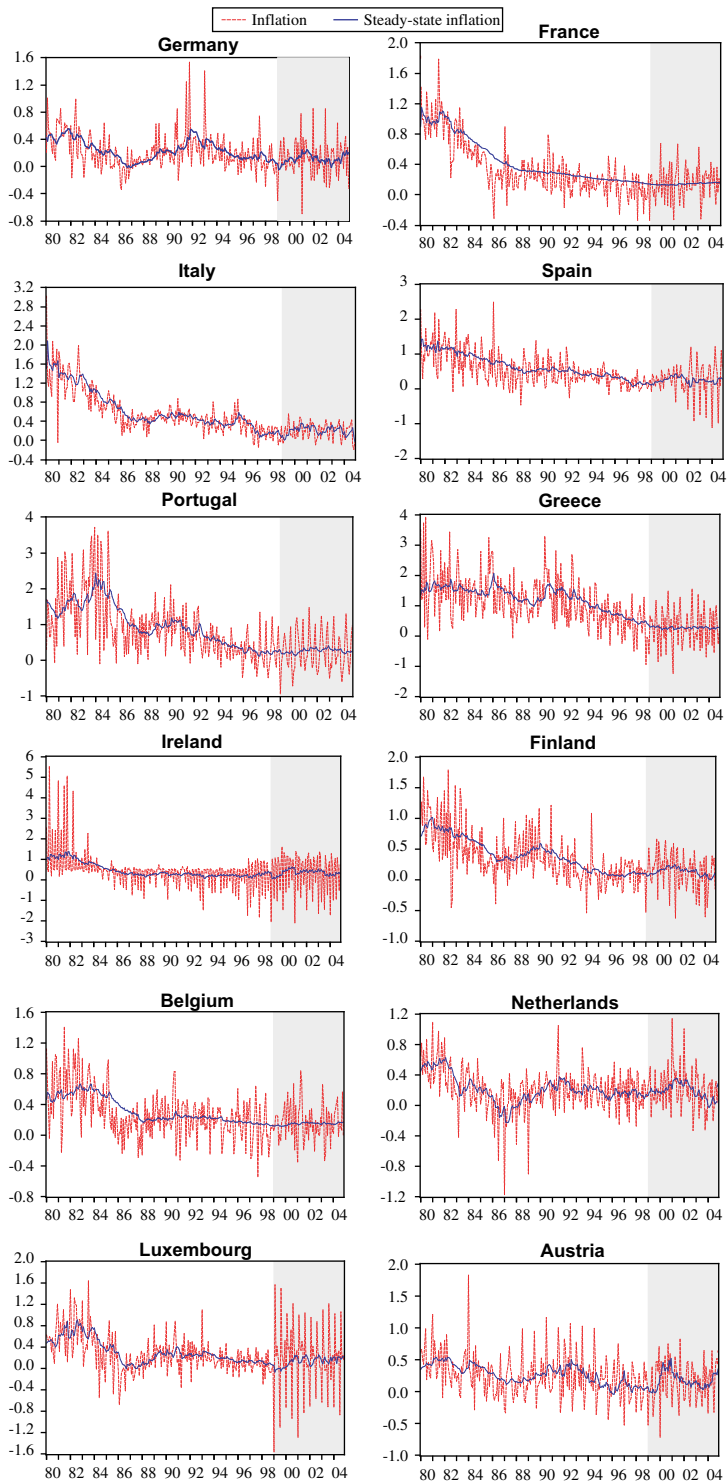


Fig. 1. Actual inflation and steady-state inflation, 1980–2004.

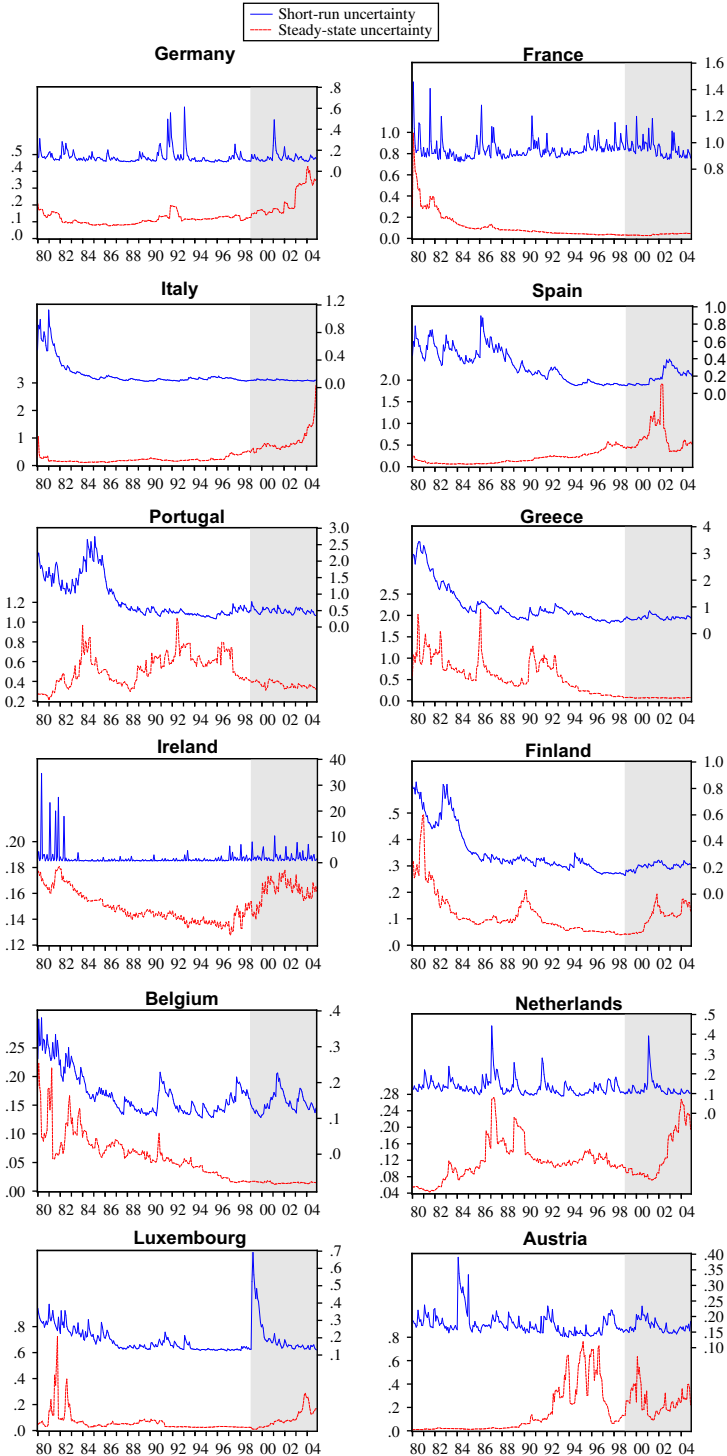


Fig. 2. Short-run and steady-state inflation uncertainty, 1980–2004.

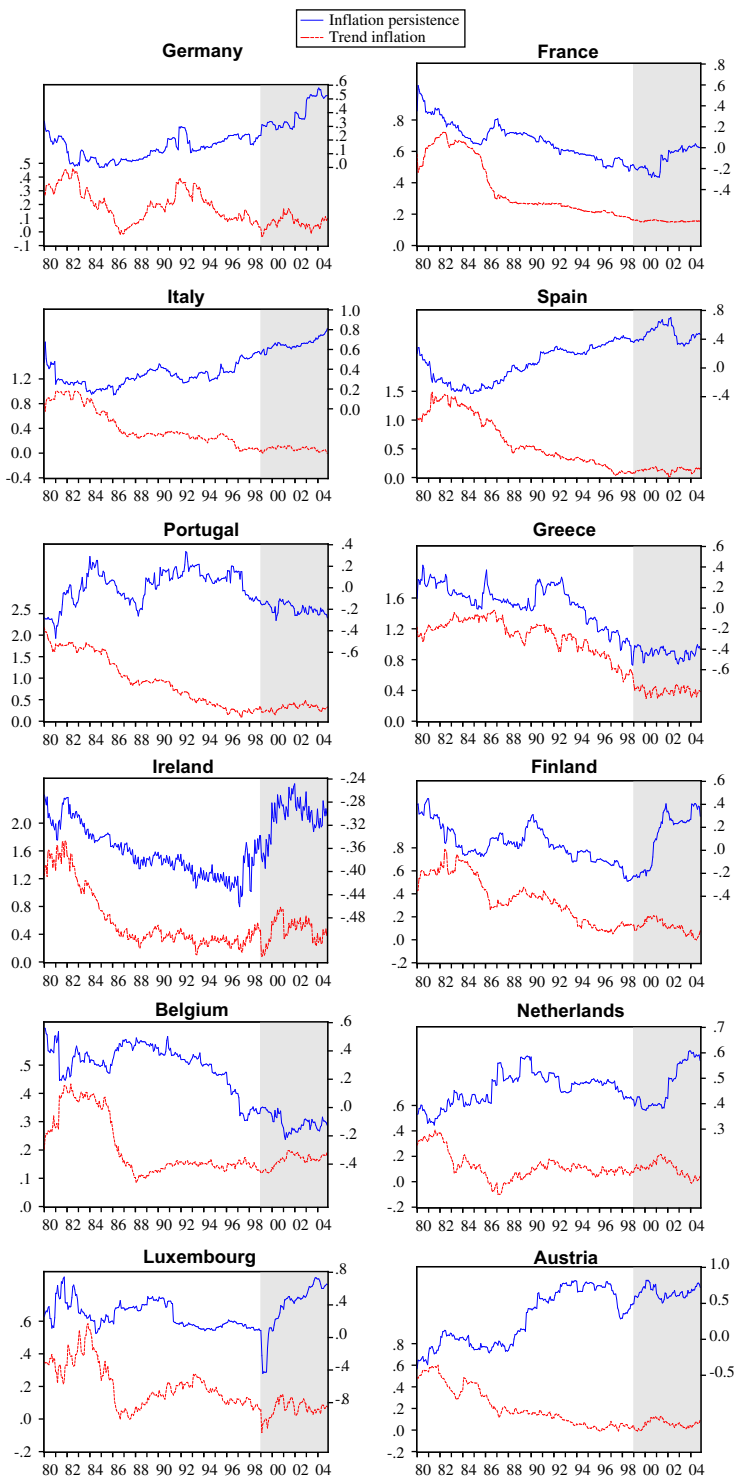


Fig. 3. Inflation persistence and trend inflation, 1980–2004.

temporary jump in 1999. From a policy point of view, it should be pointed out that some short-term volatility in inflation is inevitable given the fact that monetary policy can only affect prices with long and uncertain lags – hence the focus of the ECB on medium-term price stabilisation. Regarding the uncertainty associated with long-run inflation, it appears again that a uniform experience did not occur. Steady state uncertainty seems to increase over time (especially towards the end of the sample period) in Germany, Italy, Spain, Netherlands and Austria, while in France, Greece and Belgium it decreases over time. In Ireland and Finland a negative trend in steady state uncertainty throughout most of the sample period is replaced by a positive trend towards the end of the sample. Clearly, the presence of such significant differentials across the countries of the Euro area in terms of long-run (as opposed to short-run) uncertainty has important policy implications, given the focus of the ECB on maintaining price stability in the Euro area over longer periods of time.

Fig. 3 plots inflation persistence (the sum of the estimated autoregressive coefficients in the inflation specification) and the trend component of inflation (the estimated constant in the inflation process). The former increases over time in Germany, Italy, Spain, Netherlands and Austria. Sharp increases in inflation persistence during the Euro era can be observed in Germany, Finland and Luxembourg. These results are in line with previous work by Angeloni et al. (2006) finding that inflation persistence in the Euro area did not decline after the introduction of the Euro. Batini (2002) also shows that inflation in the Euro area is inertial using the autocorrelation function of inflation and the lag in the inflation response to monetary policy shocks from VAR's to measure inflation persistence. Our results show that in some cases inflation persistence becomes negative. This can be interpreted in terms of an error-correction mechanism in inflation: as inflation grows large, the central bank adopts tougher anti-inflationary policies. Trend inflation decreases over time in the majority of the sample countries, reflecting the general move towards lower inflation after the highly inflationary 1970s.

#### 4.4. Estimates of the inflation-uncertainty relationship

Table 3 reports robust estimates of the parameters of Eq. (11) (see Newey and West, 1994). Consistently with the hypothesis put forward by Friedman (1977) and formalised by Ball (1992), the coefficient of past inflation,  $\gamma_2$ , is positive and significant in seven out of our 12 sample countries in the steady-state uncertainty regressions, i.e. in the case of Germany, France, Greece, Ireland, Finland, Belgium and Luxembourg. In contrast to the Friedman–Ball hypothesis and in line with the arguments put forth by Pourgerami and Maskus (1987) and Ungar and Zilberfarb (1993) the relationship is negative and significant in Spain, Netherlands and Austria. When short-run uncertainty is employed as a dependent variable,  $\gamma_2$  is significantly positive in nine instances, i.e. in Germany, Italy, Spain, Portugal, Greece, Ireland, Finland, Belgium and Luxembourg.<sup>15</sup> This suggests that monetary authorities can reduce the negative consequences of inflation uncertainty by lowering average inflation. Our results for the Euro area establish a link between inflation and both short-run and long-run uncertainty in most countries, and therefore differ from those of Evans (1991) and Ball and Cecchetti (1990) for the US, as these authors documented a strong link between inflation and long-term uncertainty but reported little evidence of a link between inflation and short-term uncertainty.

As for the impact of the Euro and common monetary policy on the level of inflation uncertainty, we find that the coefficient of the intercept dummy,  $\gamma_1$ , is positive and statistically significant in eight out of 12 cases in the steady-state regressions, indicating that steady-state uncertainty has increased in the Euro period in these countries (Germany, France, Italy, Spain, Ireland, Finland, Luxembourg and Austria). Steady-state uncertainty decreased only in Portugal and Belgium. There are 50% less cases of a statistically significant intercept dummy when short-run uncertainty is considered, since only in 4/12 cases  $\gamma_1$  is positive and significant (Germany, Italy, Ireland and Luxembourg) and in 1/12 cases negative and significant (Netherlands). Hence, it appears that it was steady-state, rather than short-run, uncertainty that largely shifted after 1999.

<sup>15</sup> This is in line with previous evidence for the UK (see Kontonikas, 2004).

**Table 3**

Robust estimates of Eq. (11), 1980–2004.

Parameter	Germany		France		Italy		Spain	
	Steady-state uncertainty	Short-run uncertainty	Steady-state uncertainty	Short-run uncertainty	Steady-state uncertainty	Short-run uncertainty	Steady-state uncertainty	Short-run uncertainty
$\gamma_0$	0.104***	0.107***	0.016	0.94***	0.224***	0.036***	0.258***	0.23***
$\gamma_1$	0.115***	0.027*	0.021*	0.011	0.742***	0.07***	0.407***	-0.035
$\gamma_2$	0.039***	0.127***	0.28***	0.023	-0.013	0.268***	-0.117***	0.175***
$\gamma_3$	-0.009	-0.135**	-0.273***	-0.01	-0.51	-0.265***	-0.128	-0.179***
$R^2$	0.51	0.22	0.66	0.02	0.63	0.495	0.52	0.27
$\sigma_\theta$	0.047	0.056	0.069	0.07	0.213	0.118	0.201	0.157
Wald $F$ -stat $\gamma_2 + \gamma_3 = 0$	-	0.238	0.046	-	-	0.235	-	0.099
	Portugal		Greece		Ireland		Finland	
$\gamma_0$	0.537***	0.545***	0.232***	0.644***	0.145***	0.498***	0.071***	0.236***
$\gamma_1$	-0.171***	-0.051	-0.16	-0.031	0.015***	2.099***	0.034***	-0.023
$\gamma_2$	-0.009	0.376***	0.303***	0.352***	0.006***	2.667***	0.115***	0.256***
$\gamma_3$	0.011	-0.389***	-0.301***	-0.336***	-0.005***	-4.056***	-0.136***	-0.266***
$R^2$	0.19	0.32	0.53	0.23	0.363	0.45	0.33	0.38
$\sigma_\theta$	0.141	0.476	0.282	0.582	0.01	2.434	0.06	0.134
Wald $F$ -stat $\gamma_2 + \gamma_3 = 0$	-	0.243	8.942***	1.281	12.624***	46.5***	1.483	1.253
	Belgium		Netherlands		Luxembourg		Austria	
$\gamma_0$	0.048***	0.148***	0.132***	0.14***	0.036***	0.16***	0.166***	0.171***
$\gamma_1$	-0.033***	-0.005	-0.013	-0.028***	0.04***	0.05***	0.088*	-0.004
$\gamma_2$	0.061***	0.079***	-0.077***	-0.025	0.087***	0.0831***	-0.079*	0.019
$\gamma_3$	-0.06***	-0.072*	0.013	0.096*	-0.0874***	-0.0835**	0.045	0.006
$R^2$	0.50	0.21	0.21	0.07	0.15	0.12	0.07	0.04
$\sigma_\theta$	0.028	0.049	0.044	0.041	0.06	0.073	0.18	0.032
Wald $F$ -stat $\gamma_2 + \gamma_3 = 0$	1.545	0.11	-	-	0.0001	0.0001	-	-

Note:  $\sigma_\theta$  represents the standard deviation of the regression's residuals. The following asterisks \*\*\*, \*\*, \* indicate the 1, 5, 10% level of significance, respectively.

The coefficient of the slope dummy,  $\gamma_3$ , is negative and statistically significant in six countries (France, Greece, Ireland, Finland, Belgium and Luxembourg) in the steady-state regressions and in nine countries (Germany, Italy, Spain, Portugal, Greece, Ireland, Finland, Belgium and Luxembourg) in the short-run regressions. This indicates an important change in the underlying relationship between inflation and uncertainty occurring in these countries as a result of the introduction of the Euro, since a negative and significant  $\gamma_3$  implies that in the Euro period further reductions in average inflation tend to increase, rather than reduce, uncertainty. The Wald  $F$ -statistic for the null hypothesis:  $\gamma_2 + \gamma_3 = 0$ , indicates that after the introduction of the Euro the relationship between past inflation and current short-run uncertainty breaks down since the null hypothesis is not rejected in 8/12 cases, while remaining overall positive only in 3/12 steady-state uncertainty regressions. Thus, the Friedman–Ball link that calls for policies aiming at low inflation in order to reduce the corresponding uncertainty appears to have weakened considerably during the Euro period.

Finally, we allow for the possible structural breaks in the relationship between inflation and inflation uncertainty to be determined endogenously using the Bai and Perron (1998, 2003) procedure outlined in Section 3. The estimated break dates are reported in Tables 4 and 5 for short-run and steady-state uncertainty, respectively.<sup>16</sup> As can be seen, for most countries two or three breaks are found. Concerning the dates of the breaks, a break in relationship between inflation and

<sup>16</sup> The corresponding estimated coefficients for the implied subperiods are not included to save space, but are available from the authors upon request.

**Table 4**

Bai–Perron endogenous break test, short-run uncertainty, 1980–2004.

Countries	Number of breaks	Break dates	BIC
Germany	2	1991.07, 1996.11	-5.891
France	1	1985.07	-5.258
Italy	1	1984.12	-4.448
Spain	3	1988.09, 1993.09, 1999.11	-4.825
Portugal	2	1986.03, 1991.03	-2.723
Greece	2	1984.12, 1995.04	-1.986
Ireland	2	1984.12, 1997.01	0.77
Finland	2	1984.12, 1991.11	-5.404
Belgium	1	1984.12	-6.691
Netherlands	2	1986.02, 1991.02	-6.463
Luxembourg	3	1986.05, 1993.08, 1999.02	-5.466
Austria	3	1985.01, 1992.09, 1997.09	-6.997

Note: BIC denotes the Bayesian Information Criterion. The following specification is assumed in the Bai–Perron test:  $H_{t+1} = \delta_0 + \delta_1 \pi_t + u_{t+1}$ .

short-run uncertainty appears to occur around 1985–1986 in all countries apart from Germany and Spain. The most important policy event taking place in the then called European Community around the time of the break detected in most countries was the adoption by the Committee of Central Bank Governors of some changes in the operation of the EMS and in the rules governing the activities of the European Monetary Cooperation Fund (EMCF).<sup>17</sup> These rules entered into force on 1 July 1985.<sup>18</sup> In seven cases (Germany, Spain, Portugal, Finland, Netherlands, Luxembourg and Austria) a break is estimated in the period 1991–1993, that is, during the first stage of progress towards EMU.<sup>19</sup> Finally, in two cases (Spain, and Luxembourg) a break is estimated around 1999, the year when the third, and final, stage of EMU started with the adoption of a common currency and monetary policy. In the case of the relationship between inflation and steady-state uncertainty, seven countries exhibit a break around 1985–1986 (France, Greece, Ireland, Finland, Belgium, Netherlands and Luxembourg), eight countries during the first stage of EMU (Germany, France, Portugal, Ireland, Finland, Belgium, Netherlands and Austria), and eight countries in 1999 (Germany, Italy, Spain, Ireland, Finland, Netherlands and Luxembourg).

In general, it is clear that breaks in the relationship between the different types of inflation uncertainty and inflation itself occurred in many cases well before the introduction of the Euro and a common monetary policy on 1 January 1999, consistently with the theoretical literature that the mere announcement of a regime switching from floating to fixed rates at a given future date determines changes in the behaviour of rational agents prior to the fixing. The majority of the pre-Euro breaks in the inflation-uncertainty relationship occur around 1985 (when some changes in rules of the EMS occurred), and during the first stage of the progress towards EMU.

<sup>17</sup> In particular, there were improvements in certain aspects of the use of the ECU by the central banks: more representative ECU interest rate, change in ECU holdings against foreign currencies, ECU for “other holders”, 100% acceptability of the ECU for a creditor central bank with holdings lower than the volume allocated.

<sup>18</sup> For a chronology of relevant policy events, see “EMU: A Historical Documentation”, [http://ec.europa.eu/economy\\_finance/emu\\_history/legalaspects/part\\_c\\_1.htm](http://ec.europa.eu/economy_finance/emu_history/legalaspects/part_c_1.htm).

<sup>19</sup> In June 1989 the European Council decided that the first stage towards European Monetary Union (EMU) would begin in July 1990. The Treaty of Maastricht was agreed by the heads of state of the European Union (EU) in December 1991 setting out the framework for stages two and three of progress towards EMU. In the first stage, the members of the European Monetary System (EMS) abolished all remaining capital controls. Also, there was an increase in the degree of co-operation among the EMS central banks, while exchange rate realignments remained possible. The second stage started on 1/1/1994. During that stage, the European Monetary Institute, the precursor of the European Central Bank, was created. In order to participate in the third stage, which started on 1/1/1999 (apart from Greece where it started on 1/1/2001) countries had to satisfy five convergence criteria.

**Table 5**  
Bai–Perron endogenous break test, steady-state uncertainty, 1980–2004.

Countries	Number of breaks	Break dates	BIC
Germany	2	1990.11, 1999.11	–6.172
France	2	1984.12, 1991.01	–5.645
Italy	2	1994.11, 1999.11	–3.166
Spain	3	1989.11, 1994.11, 1999.11	–3.382
Portugal	2	1990.02, 1997.05	–4.491
Greece	2	1984.12, 1994.02	–3.069
Ireland	3	1984.12, 1990.05, 1999.09	–10.213
Finland	3	1984.12, 1992.09, 1999.11	–5.925
Belgium	3	1984.12, 1991.01, 1996.01	–7.919
Netherlands	3	1986.07, 1991.08, 1999.11	–6.613
Luxembourg	2	1984.12, 1999.11	–5.368
Austria	2	1992.07, 1997.07	–4.445

Note: BIC denotes the Bayesian Information Criterion. The following specification is assumed in the Bai–Perron test:  
 $\sigma_t^2(\pi_{t+1}^*) = \delta_0 + \delta_1 \pi_t + v_{t+1}$ .

## 5. Conclusions

In this paper, we have investigated empirically the relationship between inflation and inflation uncertainty in 12 EMU countries. Following Evans (1991) and Berument et al. (2005), we have adopted a time-varying GARCH specification to model the conditional volatility of inflation in order to be able to distinguish between short-run (structural and impulse) and steady-state uncertainty. We have also analysed the impact on the links between inflation and inflation uncertainty of the policy regime shift which occurred in 1999, when the Euro was introduced and the ECB was given the task of setting a common monetary policy for all EMU countries. First, we have imposed exogenously a break date corresponding to the actual introduction of the Euro on 1 January 1999; second, we have allowed for the possibility of an earlier adjustment in the behaviour of rational agents knowing in advance (and with certainty) that such a regime change would take place, and have therefore used a procedure for determining endogenously the timing of the breaks (see Bai and Perron, 1998, 2003).

Our empirical findings can be summarised as follows. The inflation performance of the EMU member states has been very different over the whole period starting at the beginning of the 1980s, in terms of both actual and steady-state inflation. Similarly, no consistent pattern can be found for the degree of persistence of inflation. By contrast, as one would expect given the less inflationary environment prevailing after the inflation hike of the 1970s, trend inflation has generally become much lower. Concerning short-run and steady-state uncertainty, again the EMU countries appear to have had rather different experiences, with no clear picture emerging. There is clear evidence that the Euro has had a significant impact on the relationship between inflation uncertainty and inflation, and that this has happened well before the 1st of January 1999, as agents already knew that this regime change would take place. Moreover, the Friedman–Ball link between inflation and inflation uncertainty appears to have become weaker and even to have broken down in a number of countries. Clearly, the new environment poses a huge challenge to the ECB: remaining inflation differentials mean that a common monetary policy will have asymmetric effects across member countries, as real interest rates will not be equalized. Although heterogeneity in short-run uncertainty is not surprising given the typical lags of monetary policy, differences in steady-state uncertainty are undoubtedly a major concern in view of the goal of long-term price stability. Moreover, a less effective monetary policy after 1999, as implied by a weaker Friedman–Ball link, makes the task of the ECB even more difficult.

Overall, in view of our findings, one could conclude that the ECB up to now has performed rather well in keeping inflation under control, and agree broadly with Mr Trichet's assessment. However, one might ask, could the ECB become even more successful? In particular, it could be argued that conflicting monetary signals, specifically some lack of transparency in the two-pillar strategy employed by the ECB, could be the reason why in the new economic environment monetary policy appears to have become less effective in lowering inflation uncertainty. As Bofinger (2002, p. 11) argues, "In sum, while



the first pillar is too narrowly focused on the money stock M3...the second pillar is much too broad to provide any guidance for the ECB's internal decisions or its dialogue with public". Rudebusch and Svensson (1999) also point out that emphasis on using movements in the stock of money as a rationale for policy is undesirable since it may result in higher inflation and output variability. Hence, it appears that improvements could be made to the ECB's analytical framework with a view to lowering long-run uncertainty in individual member countries – for instance, a more explicit focus on the inflation forecast might be useful in this respect.

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