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INFLATION, INFLATION UNCERTAINTY AND A COMMON EUROPEAN MONETARY POLICY*

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The relationship between inflation and inflation uncertainty is investigated in six European Union countries for the period 1960–99. Exponential generalized autoregressive conditional heteroscedasticity models are used to generate a measure of inflation uncertainty and then Granger methods are employed to test for causality between average inflation and inflation uncertainty. In all the European countries except Germany, inflation significantly raises inflation uncertainty as predicted by Friedman. However, in all countries except the UK, inflation uncertainty does not cause negative output effects, implying that a common European monetary policy applied by the European Central Bank might lead to asymmetric real effects via the inflation uncertainty channel. Less robust evidence is found regarding the direction of the impact of a change in inflation uncertainty on inflation. In Germany and the Netherlands, increased inflation uncertainty lowers inflation, while in Italy, Spain and, to a lesser extent, France increased inflation uncertainty raises inflation. These results are generally consistent with the existing rankings of central bank independence.

1 Introduction

The importance of inflation uncertainty as a distinct channel in explaining the real effects of inflation has recently been given considerable empirical support (Grier and Tullock, 1989; Grier and Perry, 2000; Judson and Orphanides, 1999). This channel was first highlighted in Friedman's (1977) Nobel Lecture. Friedman supplied an informal argument that an increase in the average inflation rate would lead to more inflation uncertainty, thus creating distortions in the workings of the price mechanism in allocating

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resources efficiently. Subsequent theoretical research focused on the opposite type of causation, which runs from inflation uncertainty to inflation. For example, Cukierman and Meltzer (1986) employ the Barro–Gordon set-up and show that an increase in uncertainty about money growth and inflation will increase the optimal average inflation rate because it provides an incentive to the policymaker to create an inflation surprise in order to stimulate output growth. Holland (1995) argues that more inflation uncertainty can lead to a lower average inflation rate if the central bank tries to minimize the welfare losses arising from more inflation uncertainty. In addition, the evidence on the direction of the effect of inflation uncertainty on inflation can be compared with the existing measures of central bank independence (Grier and Perry, 1998). These authors find that the most independent central banks are in countries where inflation declines as inflation uncertainty rises, thus contradicting the Cukierman–Meltzer hypothesis.

The issue of the relationship between inflation, inflation uncertainty and output growth acquires great importance for the member countries of the Euro-zone. First, evidence that higher inflation causes more inflation uncertainty and therefore possible negative output effects would strengthen the case for the choice of price stability by the European Central Bank (ECB) as one of the primary objectives of monetary policy. Second, if the effects of inflation on output that take place via changes in inflation uncertainty differ across the Euro-zone, it is possible that a common monetary policy that results in similar inflation rates across countries will have asymmetric real effects. In other words, a reduction in inflation arising from a contractionary monetary policy applied by the ECB could reduce output in some countries but increase output in others, depending on the combination of two effects: (a) the Friedman hypothesis, i.e. the effect of inflation on inflation uncertainty; and (b) the effect of inflation uncertainty on output growth. Therefore, lack of uniform evidence supporting the effect of inflation on output via the inflation uncertainty channel across the Euro-zone countries would have important policy implications as it would make a common monetary policy a less effective stabilization policy tool in dealing with national disparities.

Autoregressive conditional heteroscedasticity (ARCH) and generalized ARCH (GARCH) techniques represent a commonly used approach to proxy uncertainty using the conditional variance of unpredictable shocks to the inflation rate.[1] These techniques have recently been employed by Grier and Perry (1998) to investigate the direction of causality in the inflation–inflation uncertainty relationship for the G7.[2] Similarly, Grier and Perry (2000) aim to

[1]Alternative measures of uncertainty include survey-based forecasts and a moving standard deviation of inflation.

[2]We differ from this study in several respects: the GARCH model employed, the sample period, the data frequency, the country group, and the consideration of output growth and its relationship with inflation uncertainty.

examine the inflation–output uncertainty nexus in the USA. The empirical evidence to date on the Friedman and the Cukierman–Meltzer hypotheses provided by Grier and Perry (1998) and a few other recent studies summarized below is rather mixed. Grier and Perry (1998)[3] use a GARCH model to estimate inflation uncertainty and run Granger-causality tests. We employ an exponential GARCH (EGARCH) model for two reasons: first, we find evidence for asymmetries in the inflation uncertainty–inflation relationship and, second, we follow Brunner and Hess (1993) in testing Friedman's hypothesis.

Our paper contributes to the empirical relationship between inflation and inflation uncertainty in several ways. First, we use an EGARCH model instead of a GARCH model, as discussed above. Second, we examine the relationship between inflation and inflation uncertainty for several EU countries in order to examine whether a case could be made against a common monetary policy, along the lines discussed above. Third, we examine whether inflation is costly, a much-debated issue in monetary economics. Our approach allows us to distinguish between the direct costs of inflation and those that arise via the inflation uncertainty channel, as predicted by Friedman (1977). The rest of the paper is structured as follows. In Section 2 our theoretical econometric model is presented. In Section 3 we summarize our empirical results. In Section 4 we interpret these results and relate them to the predictions of economic theory and other recent empirical studies. Finally, Section 5 concludes.

2 The EGARCH Model

2.1 The AR(p)–EGARCH(1, 1) Process

One of the principal empirical tools used to model inflation uncertainty has been the ARCH class of models. Following Engle's (1982) pathbreaking idea, several formulations of conditionally heteroscedastic models (e.g. GARCH, fractionally integrated GARCH, switching GARCH, component GARCH) have been introduced in the literature, forming an immense ARCH family. However, as Brunner and Hess (1993, p. 187) argue, 'The GARCH model places a symmetric restriction on the conditional variance. Since the variance is a function of squared residuals, agents become more uncertain about future inflation whether inflation unexpectedly falls or unexpectedly rises. The essence of Friedman's hypothesis is inconsistent with such a symmetry restriction, since new information suggesting that inflation is lower should reduce, rather than raise, uncertainty about future inflation.'

[3]The authors estimate both asymmetric and symmetric GARCH models. However, they cannot reject the null hypothesis of symmetry. They therefore proceed to perform the Granger-causality tests using the estimated conditional variance from the GARCH model of each country.

Many of the proposed GARCH models include a term that can capture correlation between the inflation rate and inflation uncertainty. Models with this feature are often termed asymmetric or leverage volatility models. One of the earliest asymmetric GARCH models is the EGARCH model of Nelson (1991). In contrast to the conventional GARCH specification which requires non-negative coefficients, the EGARCH model, by modelling the logarithm of the conditional variance, does not impose the non-negativity constraints on the parameter space. Of the many different functional forms, the EGARCH model has become perhaps the most common. In particular, various cases of the EGARCH model have been applied by many researchers. For example, Brunner and Hess (1993), using EGARCH models, find that estimates of the conditional variance of US inflation are very similar to those obtained using state-dependent models.

We model the conditional mean of inflation as

$$\Phi(L)\pi_t = \phi + \varepsilon_t \tag{1a}$$

with

$$\Phi(L) \equiv \prod_{l=1}^{p}(1-\phi_l L) \tag{1b}$$

where π_t denotes the rate of inflation. Equation (1) is simply an AR(p) process.

In addition, we model the time-varying residual variance as an EGARCH(1, 1) process. This can be written as

$$\varepsilon_t = e_t h_{\pi t}^{1/2} \tag{2a}$$

$$(1-\beta L)\ln(h_{\pi t}) = \omega + d\frac{\varepsilon_{t-1}}{(h_{\pi,t-1})^{1/2}} + c\left|\frac{\varepsilon_{t-1}}{(h_{\pi,t-1})^{1/2}}\right| \tag{2b}$$

where $\{e_t\}$ is a sequence of independent, normally distributed random variables with mean zero and variance 1. In the empirical work reported below, we estimate AR(p)–EGARCH(1, 1) models for inflation and then use the conditional variance $h_{\pi t}$ as a measure of inflation uncertainty.

3 Empirical Analysis

3.1 Methodological Issues

The relationship between inflation and inflation uncertainty could be estimated in a simultaneous approach as in a GARCH-in-mean (GARCH-M) model that includes a function of the lagged inflation rate in the conditional variance equation or in a two-step approach where an estimate of the conditional variance is first obtained from a GARCH-type model and then causality tests are run to test for bidirectional effects. Examples of the former approach include Brunner and Hess (1993), Grier and Perry (1998), Baillie *et al.* (1996) and Fountas *et al.* (2000).[4] The latter approach was followed in Grier and Perry (1998).

The simultaneous approach suffers from the disadvantage that it does not allow the testing of a lagged effect of inflation uncertainty on inflation, which would be expected in a study that employs monthly or quarterly data. As Grier and Perry (1998) mention, the impact of a change in inflation uncertainty on average inflation, via a change in the stabilization policy of the monetary authority, takes time to materialize and cannot be fairly tested in a model that restricts the effect to being contemporaneous.

3.2 The Empirical Evidence to Date

The inflation–inflation uncertainty relationship has been analysed extensively in the empirical literature. Holland (1993) and Davis and Kanago (2000) survey this literature. Inflation uncertainty is measured either using survey-based forecasts of inflation or the GARCH approach. In the recent literature that employs the GARCH approach, the US evidence in favour of the Friedman hypothesis is mixed. Brunner and Hess (1993), Grier and Perry (1998, 2000) and Fountas *et al.* (2000) find evidence in favour, whereas Baillie *et al.* (1996) find evidence against it. The US evidence on the Cukierman–Meltzer hypothesis is rather negative. Only Fountas *et al.* (2000) find evidence in favour of the hypothesis. There are a limited number of studies using international data that employ the GARCH approach. They are Baillie *et al.* (1996) and Grier and Perry (1998). Grier and Perry (1998) find evidence supporting the Friedman hypothesis in the rest of the G7 countries but Baillie *et al.* (1996) find mixed evidence. Grier and Perry (1998) find evidence supporting the Cukierman–Meltzer hypothesis in Japan and France and Baillie *et al.* (1996) in the UK and three high-inflation economies, Argentina, Brazil and Israel.

This study aims to fill the gaps arising from the methodological shortcomings of the previous studies and the lack of interest in the European case, where the results would have interesting implications for the successful implementation of common European monetary policy.

[4]Grier and Perry (1998) use a component GARCH-M model of US inflation that includes lagged inflation in the conditional variance, whereas Brunner and Hess (1993) use a state-dependent model where the standard deviation of inflation is included in the mean equation and the lagged value of the squared deviation of inflation from a parameter is included in the variance equation. Baillie *et al.* (1996) model inflation as a fractionally integrated process and include lagged inflation in the conditional variance equation and the standard deviation in the mean equation. Fountas *et al.* (2000) use a GARCH-M model that includes the lagged inflation rate in the variance equation.

3.3 UK Results

3.3.1 Description of the UK Data and Estimation Results. We first test for the relationship between inflation and inflation uncertainty using UK data. Even though the UK is not at present a member of the Euro-zone, it is likely that it will participate in the European monetary union (EMU) in the future. In our empirical application we use non-seasonally adjusted time series data on the consumer price index (CPI) obtained from the OECD Main Economic Indicators Database. Our sample includes quarterly data from 1960:Q1 through 1999:Q2. Figure 1 plots the inflation rate (π_t) series constructed as the first difference of the log of CPI. To establish that the inflation data series is stationary we use both the augmented Dickey–Fuller (ADF) and Phillips–Perron tests presented in Table 1, part (a). Using the second lagged difference terms in the ADF test and setting the truncation lag at 4 in the Phillips–Perron test, we find that both tests reject the null hypothesis of a unit root at the 0.01 significance level. Hence, we have evidence in this sample that the UK inflation rate is stationary. We choose an AR(6) plus two seasonal

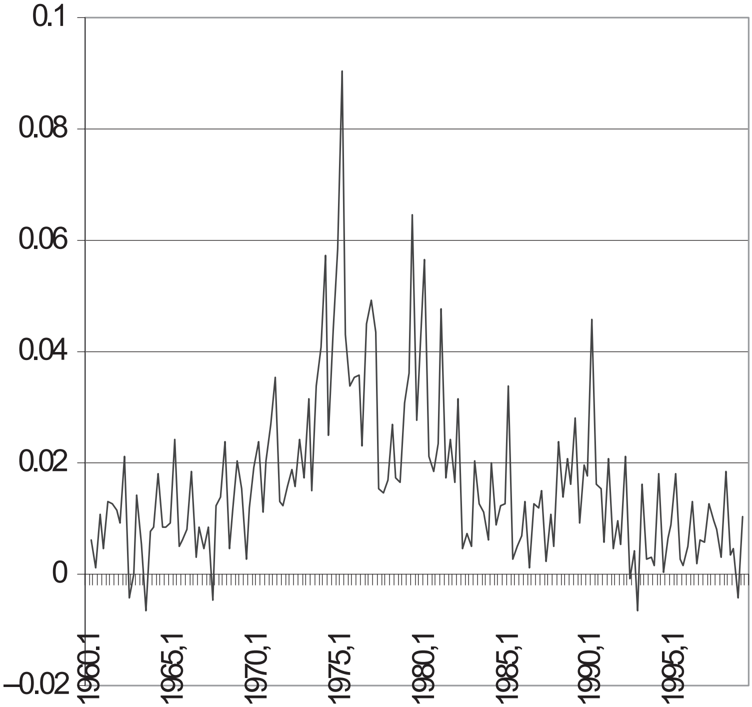

FIG. 1 UK Inflation Rate, 1960:Q1–1999:Q2

TABLE 1

(a) Inflation unit root tests

Country	ADF t statistic	Phillips–Perron t statistic
UK	−2.620***	−8.040***

(b) The estimated AR(6)–EGARCH(1, 1) model for the UK inflation rate

$$\underset{(0.00)}{\pi_t = 0.003} + \underset{(0.00)}{0.566}\pi_{t-1} + \underset{(0.00)}{0.178}\pi_{t-4} - \underset{(0.01)}{0.169}\pi_{t-5} + \underset{(0.00)}{0.188}\pi_{t-6} + \underset{(0.00)}{0.009}d_{2t} - \underset{(0.00)}{0.007}d_{3t} + \varepsilon_t$$

$$\ln(h_{\pi t}) = -\underset{(0.01)}{3.196} + \underset{(0.00)}{0.697}\ln(h_{\pi,t-1}) + \underset{(0.18)}{0.242}|e_{t-1}| + \underset{(0.02)}{0.437}e_{t-1}$$

$Q(12) = 8.789$ [0.721], $Q(24) = 24.810$ [0.416], $Q(36) = 38.650$ [0.351]
$Q^2(12) = 12.957$ [0.372], $Q^2(24) = 22.579$ [0.545], $Q^2(36) = 26.794$ [0.867]

Notes: In the ADF tests in (a), we use two lagged differenced terms. In the Phillips–Perron tests, the truncation lag is set at 4.
The first equation in (b) represents the estimated conditional mean of the autoregressive model. d_{2t} and d_{3t} are seasonal dummies. The figures in parentheses under the coefficients and inside the square brackets show the probability values.
*** Rejection of the unit root null at the 0.01 level.

dummy variables[5] model for the mean inflation rate and an EGARCH(1, 1) model for the variance equation, according to the minimum Akaike information criterion (AIC) and Schwarz criterion (SC).

Table 1, part (b), presents the estimates of an AR(6)–EGARCH(1, 1) model for the UK inflation rate with two seasonal dummies. The model was estimated under quasi-maximum likelihood estimation using the consistent variance–covariance estimator of Bollerslev and Wooldridge (1992). Residual diagnostics for this model are also reported in Table 1, part (b), and include Ljung–Box (Q) tests for residual correlation and Ljung–Box diagnostics for serial dependence in the squared residuals. As reported, the Ljung–Box tests for serial correlation in the levels and squares of the standardized residuals do not reject the hypothesis of no autocorrelation. Thus, the Ljung–Box tests indicate that the estimated model fits the data very well. The persistence of volatility implied by the EGARCH equation is measured by the size of β, which is highly significant. Asymmetry in inflation uncertainty is conveniently quantified by examining the sign of d. In the present case, the positive and significant value of the coefficient implies that periods of positive inflation shocks are accompanied by high inflation uncertainty and periods of negative inflation shocks are accompanied by lower uncertainty about inflation. In summary, the AR(6)–EGARCH(1, 1) model seems to fit both the mean and variance of the UK inflation rate quite well.

[5]The seasonal dummy variables are included to seasonally adjust the inflation series. We find that two of these dummies are jointly statistically significant.

TABLE 2
GRANGER-CAUSALITY TESTS: UK (1960–99)

	H_0: $\pi_t \to h_{\pi t}$	H_0: $h_{\pi t} \to \pi_t$	H_0: $h_{\pi t} \to y_t$	H_0: $h_{\pi t} \to y_t^{a}$
Two lags	4.741*** (+)	3.070** (+)	8.775*** (–)	6.770*** (–)
Four lags	4.721*** (+)	8.343*** (–)	6.410*** (–)	3.589*** (–)
Six lags	9.202*** (+)	6.864*** (+)	5.410*** (–)	1.610
Eight lags	5.294*** (+)	5.568*** (–)	3.345*** (–)	0.966

Notes: (1) The figures are F statistics.
(2) $\pi_t \to h_{\pi t}$, inflation does not Granger-cause inflation uncertainty; $h_{\pi t} \to \pi_t$, inflation uncertainty does not Granger-cause inflation; $h_{\pi t} \to y_t$, inflation uncertainty does not Granger-cause output growth.
(3) The positive or negative sign in parentheses indicates the sign of the sum of the lagged coefficients in the respective equation.
[a] Lagged inflation has been added to the regression.
*** Rejection of the null hypothesis at the 0.01 level of significance.
** Rejection of the null hypothesis at the 0.05 level of significance.
* Rejection of the null hypothesis at the 0.10 level of significance.

3.3.2 Granger-causality Tests. Next we employ Granger methods to test for bidirectional causality between inflation and inflation uncertainty. In particular, we test the null hypotheses that inflation does not Granger-cause inflation uncertainty and that inflation uncertainty does not Granger-cause inflation, using two, four, six and eight lags.[6] The F statistics are reported in Table 2. These statistics have been obtained following correction for serial correlation and/or heteroscedasticity in the unrestricted regression in each case. The first null hypothesis is rejected at the 0.01 level for all lags, while the second is also rejected at 0.10 or better. The sums of the coefficients on lagged uncertainty in the inflation equation (at lags 2 and 6) and on lagged inflation in the inflation uncertainty equation are positive. We thus provide strong empirical confirmation of Friedman's hypothesis. We also find some evidence that increased inflation uncertainty increases inflation, confirming the theoretical predictions made by Cukierman and Meltzer (1986).

Inflation uncertainty has real effects only if it leads to output losses. To test for such effects we have used the index of industrial production to construct the growth rate of output. Our Granger-causality results in Table 2 (fourth column) indicate that higher inflation uncertainty causes a negative output growth effect, thus supporting the argument that higher inflation uncertainty is part of the welfare costs of inflation.[7] Finally, in the last column of Table 2, we report the F statistics on the causal effect of inflation uncertainty on output growth, where the regression includes in addition lagged inflation rates. The rationale for this choice is to control for possible effects of inflation uncertainty on output that take place via changes in

[6]It is possible to test for the relationship between inflation and its uncertainty simultaneously, as argued in the Appendix.
[7]This result is identical if we use GDP to measure real output.

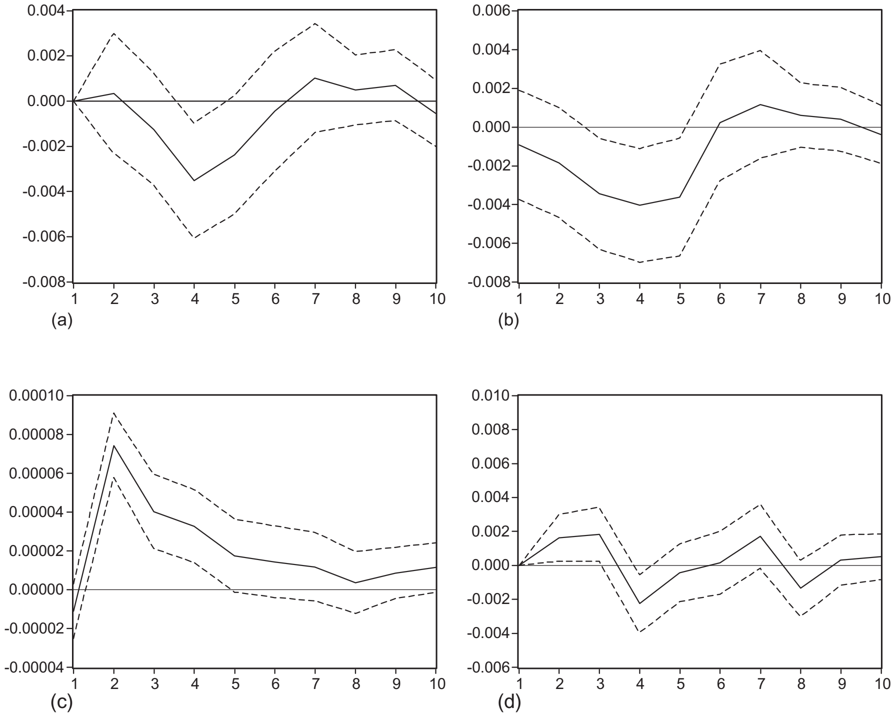

FIG. 2 Effects of a One-time One-standard-deviation Increase (UK): (a) Effects of a Change in Uncertainty on Output; (b) Effects of a Change in Uncertainty on Output (Bivariate VAR); (c) Effects of a Change in Inflation on Uncertainty; (d) Effects of a Change in Uncertainty on Inflation

Note: The dotted lines indicate ± two standard deviation bands computed by the asymptotic standard errors.

inflation.[8] The reported results indicate that inflation uncertainty still affects output negatively, even though the effect is perhaps somewhat weaker (i.e. it applies for two and four lags only).

Figure 2 plots (i) the time profile of output growth due to shocks in inflation uncertainty and (ii) the time profiles of inflation uncertainty and inflation due to shocks in inflation and inflation uncertainty, respectively. Parts (a) and (b) indicate that the negative impact of inflation uncertainty on output reaches its peak after four quarters and equals about 0.5 per cent. Part (c) shows that the maximum effect of inflation on its uncertainty takes place after two quarters. Finally, according to part (d), the sign of the effect of inflation uncertainty on inflation varies considerably over time. The effect

[8]We are grateful to an anonymous referee for this point.

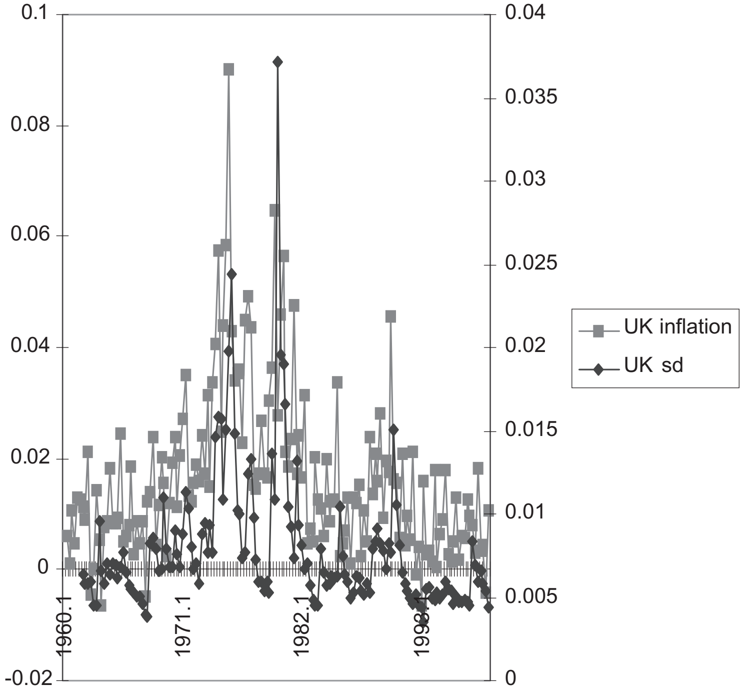

FIG. 3 UK Inflation and Conditional Standard Deviation, AR(6)–EGARCH(1, 1) Model

also seems to be much smaller in size (about 0.2 per cent) than the effect of inflation uncertainty on output.

3.3.3 Predictability of Higher Levels of UK Inflation. Several researchers, such as Engle (1983) and Cosimano and Jansen (1988), have failed to find strong evidence that higher rates of inflation are less predictable. Using the EGARCH model, we compare our results with theirs. The inflation and inflation uncertainty series for the AR(6)–EGARCH(1, 1) model are shown in Fig. 3, which plots the inflation rate and its corresponding conditional standard deviation in dual scale.

In contrast to the conclusion of the above-mentioned studies, Fig. 3 provides evidence that higher levels of inflation are less predictable. According to our estimates, the conditional standard deviation average (annual rate) in the low-inflation 1960s is about 2.4 per cent. In the high-inflation 1970s, the conditional standard deviation average (annual rate) is about 4.3 per cent. Finally, in the low-inflation environment of the 1990s, the average of the conditional standard deviation is only 2.4 per cent. Brunner and Hess (1993) argue that it is the relaxation of the symmetry restriction in condi-

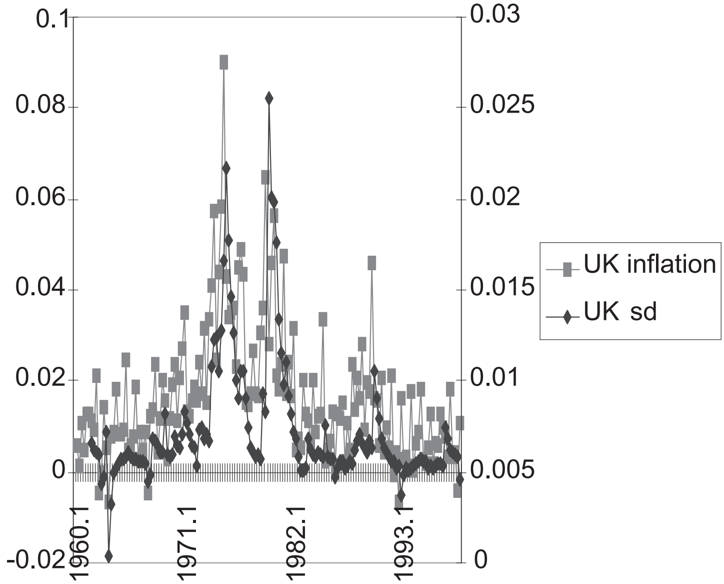

FIG. 4 Inflation Rate and Conditional Standard Deviation, AR(6)–Asymmetric GARCH(1, 1) Model

tional volatility models which enables them to find that higher levels of inflation are less predictable. We reach the same conclusion by using an EGARCH model. To compare our results with theirs we also use an asymmetric GARCH process. The AIC and SC were −6.979647 and −6.759845 respectively, much worse than those of the EGARCH model. The estimates of the conditional standard deviation were quite unsatisfactory as well. Figure 4 shows that the volatility of inflation for the GARCH model is unreasonably high during the relatively low and stable inflation years of the late 1980s and 1990s.

3.4 *Evidence for the Euro-zone Countries*

3.4.1 Description of the Data and Estimation Results. We apply the above empirical approach to five European countries (France, Germany, Italy, the Netherlands and Spain) that are at present members of the Euro-zone. Our group of countries includes the four largest EMU countries. We use quarterly non-seasonally adjusted time series on CPI obtained from the OECD Main Economic Indicators Database from 1960:Q1 to 1999:Q3.[9] To adjust

[9]The time series for France ends in 1999:Q2, for Germany in 1999:Q2, for Italy in 1999:Q3, for the Netherlands in 1999:Q2 and for Spain in 1999:Q2.

Table 3
Inflation Unit Root Tests

Country	*ADF t statistic*	*Phillips–Perron t statistic*
France	−1.910	−3.110**
Germany	−2.150	−5.420***
Italy	−2.340	−3.880***
Netherlands	−4.840***	−11.150***
Spain	−2.720*	−6.550***

Notes: In the ADF tests we use two lagged differenced terms. In the Phillips–Perron tests the truncation lag is set at 4.
*** Rejection of the unit root null at the 0.01 level of significance.
** Rejection of the unit root null at the 0.05 level of significance.
* Rejection of the unit root null at the 0.10 level of significance.

the time series for seasonality, we use three seasonal dummy variables in each country, provided they are jointly significant.[10]

Table 3 presents ADF and Phillips–Perron tests of the unit root hypothesis for each country. The Phillips–Perron tests reject the null hypothesis of a unit root for all six countries at the 0.01 (0.05 for France) significance level. The ADF tests for France, Germany and Italy fail to reject the null hypothesis of a unit root, but we will consider their inflation series stationary in our analysis, taking into consideration the Phillips–Perron results.

The best fitted model is chosen according to the minimum values of the AIC and SC. We choose an EGARCH(1, 1) specification for the conditional variance and an AR(3) model for France, an AR(7) for Germany, an AR(4) for Italy and an AR(8) for the Netherlands and Spain. Table 4 shows the estimated results for each country for the models specified above. In all countries except Germany, the estimated coefficient d is statistically significant and positive, indicating evidence of asymmetry in the conditional variance. This implies that negative and positive shocks to the inflation process have a different impact on inflation uncertainty. More specifically, positive (negative) inflation surprises lead to more (less) inflation uncertainty. For Germany, the estimated coefficient of asymmetry is negative, implying that a positive inflation shock leads to less uncertainty about inflation. This finding can be attributed to the strong commitment of the German monetary authority towards anti-inflationary policies. We also perform the same specification tests for the adequacy of the models as we did for the UK above. For all the estimated models, residuals diagnostics (not reported) yield no evidence of mis-specification.

3.4.2 Granger-causality Tests. Table 5 reports the Granger-causality test results for the above five countries. The null hypothesis that inflation does not

[10]The seasonal dummy variables are jointly significant in all the countries examined except for the Netherlands and Spain.

TABLE 4
THE ESTIMATED AR(p)–EGARCH(1, 1) MODELS

	Country				
Parameter	*France*	*Germany*	*Italy*	*Netherlands*	*Spain*
π_{t-1}	0.645 [0.000]	0.351 [0.000]	0.754 [0.000]	0.471 [0.000]	0.182 [0.006]
π_{t-2}	–0.016 [0.837]		0.114 [0.240]		0.247 [0.001]
π_{t-3}	0.296 [0.000]	0.279 [0.000]	0.063 [0.408]		0.191 [0.022]
π_{t-4}		0.361 [0.000]	0.039 [0.420]	0.337 [0.000]	0.232 [0.000]
π_{t-5}				–0.195 [0.000]	
π_{t-6}				0.115 [0.010]	
π_{t-7}		–0.205 [0.003]			–0.193 [0.003]
π_{t-8}				0.218 [0.000]	0.253 [0.000]
d	0.223 [0.000]	–0.032 [0.736]	0.422 [0.000]	0.347 [0.050]	0.126 [0.001]
c	–0.215 [0.013]	0.319 [0.070]	–0.194 [0.000]	0.912 [0.000]	–0.163 [0.048]
β	0.942 [0.000]	0.680 [0.011]	0.942 [0.000]	0.745 [0.000]	0.969 [0.000]

Notes: (1) The estimated conditional variance equation has the form

$\ln(h_{\pi t}) = \omega + \beta \ln(h_{\pi,t-1}) + c\,|e_{t-1}| + de_{t-1}$

(2) A constant term and seasonal dummies were included but are not reported.
(3) Probability values are given in square brackets.

Granger-cause inflation uncertainty is rejected for all the countries examined at the 0.05 level and for each lag length, except Germany. These results are similar to those for the UK, supporting the Friedman hypothesis. The null hypothesis that inflation uncertainty does not Granger-cause inflation is rejected in all countries. However, only in the case of Italy, France (two lags) and Spain (six and eight lags) is the effect positive, supporting the Cukierman–Meltzer hypothesis. For Germany and the Netherlands, where the effect is negative, we find evidence in favour of Holland's (1995) stabilization hypothesis discussed below.

Finally, the fourth column of Table 5 indicates that inflation uncertainty does not Granger-cause output growth in all countries, except perhaps Italy, where we find a significant and negative impact on real output growth at the 10 per cent level, and the Netherlands and Spain, where the effect is positive. Somewhat similar results apply in the last column of Table 5, which adds the inflation rate in the right-hand side of the regression. The only differences are the insignificance of inflation uncertainty in Italy and the slight evidence for a positive impact in France. These results are discussed further below.[11]

3.4.3 Predictability of Inflation. As in the case of UK inflation, there is evidence that higher rates of inflation are less predictable for each of the other

[11]The choice of the industrial production index in measuring real output is dictated by the unavailability of quarterly national accounts (and hence GDP) for the full sample period for several countries in our sample.

Table 5
Granger-causality Tests for Inflation and Inflation Uncertainty (1960–99)

	H$_0$: $\pi_t \rightarrow h_{\pi t}$	H$_0$: $h_{\pi t} \rightarrow \pi_t$	H$_0$: $h_{\pi t} \rightarrow y_t$	H$_0$: $h_{\pi t} \rightarrow y_t$[a]
France				
Two lags	46.381*** (+)	6.319*** (+)	0.274	1.425
Four lags	28.570*** (+)	0.002	0.443	0.657
Six lags	22.838*** (+)	0.240	1.128	1.902* (+)
Eight lags	17.676*** (+)	0.918	1.003	1.110
Germany				
Two lags	1.200	0.946	0.850	0.728
Four lags	0.669	2.780** (–)	0.517	0.528
Six lags	0.350	3.235*** (–)	1.151	0.535
Eight lags	0.407	3.520*** (–)	0.803	0.409
Italy				
Two lags	42.552*** (+)	0.039	2.540* (–)	0.023
Four lags	39.568*** (+)	4.559*** (+)	1.277	0.389
Six lags	34.108*** (+)	6.458*** (+)	0.856	0.853
Eight lags	38.773*** (+)	4.334*** (+)	1.754* (–)	1.016
Netherlands				
Two lags	20.394*** (+)	5.188*** (–)	8.106*** (+)	5.978*** (+)
Four lags	9.382*** (+)	5.256*** (–)	6.889*** (+)	3.318** (+)
Six lags	10.114*** (+)	3.526*** (–)	3.734*** (+)	1.806* (+)
Eight lags	8.280*** (+)	2.820*** (–)	7.823*** (+)	0.606
Spain				
Two lags	10.021*** (+)	1.081	3.254** (+)	6.643*** (+)
Four lags	8.221*** (+)	0.454	1.766	3.256*** (+)
Six lags	7.142*** (+)	2.227** (+)	2.113** (+)	4.026*** (+)
Eight lags	4.952*** (+)	2.212** (+)	1.346	3.677*** (+)

Notes: (1) The figures are *F* statistics.
(2) $\pi_t \rightarrow h_{\pi t}$, inflation does not Granger-cause inflation uncertainty; $h_{\pi t} \rightarrow \pi_t$, inflation uncertainty does not Granger-cause inflation; $h_{\pi t} \rightarrow y_t$, inflation uncertainty does not Granger-cause output growth.
(3) The positive or negative sign in parentheses indicates the sign of the sum of the lagged coefficients in the respective equations.
[a] Lagged inflation has been added to the regression.
*** Rejection of the null at the 0.01 level of significance.
** Rejection of the null at the 0.05 level of significance.
* Rejection of the null at the 0.10 level of significance.

European countries. This conclusion is derived from an examination of the plots (not reported) of the inflation rate and its corresponding conditional standard deviation for each country. This result is in agreement with the conclusion of Brunner and Hess (1993) for the USA. According to our estimated model for France, the average of the conditional standard deviation (annual rate) in the high-inflation 1970s is 2 per cent and in the low-inflation environment of the 1990s only 0.9 per cent. Similarly, according to our estimated model for Spain, the average value of the conditional standard deviation (annual rate) in the 1970s is 4 per cent whereas in the stable inflationary environment of the 1990s the average figure is 1.6 per cent. Similar results apply for the rest of the countries in our sample.

TABLE 6
GRANGER-CAUSALITY TESTS: UK

	H_0: $\pi_t \to h_{\pi t}$	H_0: $h_{\pi t} \to \pi_t$	H_0: $h_{\pi t} \to y_t$	H_0: $h_{\pi t} \to y_t$[a]
1979–90				
Two lags	0.350	14.867*** (–)	0.873	0.360
Four lags	3.150** (+)	6.233*** (–)	1.375	0.554
Six lags	2.585** (+)	2.785** (–)	1.030	0.428
Eight lags	3.495*** (–)	2.615** (–)	0.909	0.560
1973–99				
Two lags	38.047*** (+)	0.722	2.514* (–)	1.061
Four lags	38.830*** (+)	2.286* (–)	2.518** (–)	0.815
Six lags	55.507*** (+)	1.050	5.604*** (–)	0.896
Eight lags	39.675*** (+)	1.118	3.700*** (–)	1.113

Notes: (1) The figures are F statistics.
(2) $\pi_t \to h_{\pi t}$, inflation does not Granger-cause inflation uncertainty; $h_{\pi t} \to \pi_t$, inflation uncertainty does not Granger-cause inflation; $h_{\pi t} \to y_t$, inflation uncertainty does not Granger-cause output growth.
(3) The positive or negative sign in parentheses indicates the sign of the sum of the lagged coefficients in the respective equations.
[a] Lagged inflation has been added to the regression.
*** Rejection of the null at the 0.01 level of significance.
** Rejection of the null at the 0.05 level of significance.
* Rejection of the null at the 0.10 level of significance.

3.5 Robustness

Our sample period 1960–99 includes various exchange rate and monetary policy regimes. For example, the UK operated under a managed float regime, following the collapse of the Bretton Woods system, for most of our sample, except for the brief period of exchange rate mechanism participation. In addition, from 1979 to 1990, Thatcher's government emphasized a strong anti-inflation objective. Hence, we repeat the above analysis for the UK for two periods: 1973–99 (the managed float regime) and 1979–90 (the Thatcher years). The results are presented in Table 6. Overall, these results are in broad agreement with those reported in Table 2. It is interesting to note that during the Thatcher years inflation uncertainty had no impact on output growth, a result that is very robust to the presence or absence of lagged inflation in the regression equation. Moreover, there is some evidence, in both periods under consideration, that inflation uncertainty lowers inflation, in agreement with the stabilization hypothesis.

For the rest of our sample, countries that were exchange rate mechanism members for most of our original sample period 1960–99, we repeat the above analysis for the period 1983–99.[12] The choice of this period is based on the widely accepted notion that the exchange rate mechanism entered a calmer phase in 1983 following the turbulent early years (Gros and Thygesen, 1992).

[12]For all six countries, we have re-estimated the GARCH model using the new sample periods and obtained new values for the conditional variances. These values have then been used in performing the Granger-causality tests.

Following the estimation of GARCH models for each country (results not reported), we perform Granger-causality tests as previously and report the results in Table 7. These results support those reported for the full sample period (see Table 5) in many respects. First, we find strong support for the Friedman hypothesis regarding the positive impact of inflation on inflation uncertainty in most countries (see the second column). Second, as was the case in the analysis of the full period, we find that inflation uncertainty does not seem to lead to lower output, with a single exception (last two columns of Table 7). Finally, significant difference obtains between the full sample and the post-1983 period on the significance of the causal effect of inflation uncertainty on inflation. We find (the third column) that in most countries

Table 7
Granger-causality Tests for Inflation and Inflation Uncertainty (1983–99)

	H_0: $\pi_t \to h_{\pi t}$	H_0: $h_{\pi t} \to \pi_t$	H_0: $h_{\pi t} \to y_t$	H_0: $h_{\pi t} \to y_t^{a}$
France				
Two lags	4.724** (+)	0.001	2.015	2.373* (+)
Four lags	3.432** (+)	0.169	1.905	1.766
Six lags	3.031** (+)	0.237	0.552	1.018
Eight lags	3.111*** (+)	0.396	0.604	1.269
Germany				
Two lags	2.762* (+)	0.651	1.566	1.410
Four lags	4.026*** (+)	0.943	1.670	1.689
Six lags	2.927** (+)	0.845	2.479** (–)	1.643
Eight lags	2.293** (+)	0.691	1.937* (–)	1.374
Italy				
Two lags	5.898*** (+)	3.605** (–)	1.127	2.079
Four lags	5.576*** (+)	2.038* (–)	0.780	1.535
Six lags	3.015** (+)	2.933** (–)	1.018	1.778
Eight lags	1.947* (+)	2.608** (–)	1.033	0.844
Netherlands				
Two lags	3.829** (–)	4.492** (–)	0.546	0.299
Four lags	8.285*** (–)	1.442	2.068* (+)	0.124
Six lags	14.995*** (–)	0.971	1.585	0.297
Eight lags	9.717*** (–)	1.069	0.539	0.686
Spain				
Two lags	0.890	1.629	0.081	0.074
Four lags	1.567	1.659	2.594** (+)	2.923** (+)
Six lags	0.754	1.483	2.004* (+)	1.549
Eight lags	0.561	1.772	1.626	1.245

Notes: (1) The figures are F statistics.
(2) $\pi_t \to h_{\pi t}$, inflation does not Granger-cause inflation uncertainty; $h_{\pi t} \to \pi_t$, inflation uncertainty does not Granger-cause inflation; $h_{\pi t} \to y_t$, inflation uncertainty does not Granger-cause output growth.
(3) The positive or negative sign in parentheses indicates the sign of the sum of the lagged coefficients in the respective equations.
[a] Lagged inflation has been added to the regression.
*** Rejection of the null at the 0.01 level of significance.
** Rejection of the null at the 0.05 level of significance.
* Rejection of the null at the 0.10 level of significance.

there is no causal effect. This result squares with the loss of monetary policy independence in the exchange rate mechanism period, as monetary policy was constrained by the exchange rate peg objective.

4 Discussion

Our full sample period includes considerable inflation diversity both across countries and across time. The high-inflation 1970s was followed by the low-inflation 1980s and 1990s. This was the case for two reasons: first, the global reduction in inflationary pressures; second, some European countries, France, Italy and Spain in our sample, joined the European Monetary System (EMS) in 1979 in order to borrow Germany's anti-inflation reputation. This is less so for the Netherlands, which has traditionally aligned its monetary policy stance to Germany's. The reduction in inflation for France, Italy and Spain was more prevalent during the last stage of the EMS, starting in 1987. During most of the 1990s, inflation remained low and relatively stable. The significant variability in the level of inflation and the uncertainty about it during our sample period provides the testing ground to examine the bidirectional relationship between inflation and inflation uncertainty.

We discuss first the Granger-causality results on the effect of inflation on inflation uncertainty. Our results indicate strong evidence in support of the Friedman hypothesis for all countries except Germany.[13] The lack of evidence for Germany is not surprising as it is consistent with Ball's (1992) theory, which formalized Friedman's prediction.[14] Using Ball's (1992) argument, an increase in German inflation would not lead to more inflation uncertainty as the Bundesbank had a strong anti-inflation reputation and therefore was willing to bear the costs of disinflation. Our result on the Friedman hypothesis for France, Italy and the UK is consistent with the Grier and Perry (1998) study of the G7. However, in contrast to our study, Grier and Perry (1998), using a different methodology (GARCH model), sample size (1948–93) and data frequency (monthly, as opposed to quarterly), find support for Friedman's hypothesis for Germany. Our finding of a non-causal effect of inflation on inflation uncertainty for this country indicates that inflation uncertainty in Germany is not caused by rising inflation rates.

Regarding the causality from inflation uncertainty to output growth, our Granger-causality tests indicate that only in the UK does inflation uncertainty have a negative effect on output growth (when the full period is used). In the EMU countries during the more relevant 1983–99 period such

[13]No evidence in favour of the Friedman hypothesis applies for the Netherlands and Spain when the 1983–99 period is used.

[14]Ball (1992) uses an asymmetric information game where two policymakers with different preferences towards inflation alternate stochastically in office. Therefore, a higher current inflation rate raises inflation uncertainty as it is not known which policymaker will be in office in the next period.

an effect does not apply. This is according to the last column of Table 7 which allows us to separate the effects of inflation uncertainty on output. Hence, we conclude that the welfare costs of inflation do not seem to be significant, with the exception of the UK.[15] This finding has important implications for the ECB's policymaking strategy. In particular, it supports those claiming that the objective of price stability has been overemphasized by the ECB. The second implication of these output growth Granger-causality results concerns the application of a common monetary policy by the ECB following the launch of the Euro-zone in 1999. As we saw earlier, inflation uncertainty does not seem to cause negative real effects across all countries in our sample, except in the UK. Hence, a common European monetary policy would have relatively asymmetrical real effects, which work through the inflation uncertainty channel, across the EMU countries.[16]

Our evidence on the Cukierman–Meltzer hypothesis is rather mixed. For Germany and the Netherlands, we find evidence against this hypothesis. This evidence partially favours the 'stabilization hypothesis' put forward by Holland (1995). He claims that, for countries where inflation leads to inflation uncertainty and real costs, we would expect the central bank to stabilize inflation—hence a negative effect of inflation uncertainty on inflation. Our evidence is in part[17] consistent with this argument for Germany and the Netherlands. In contrast, for Italy, France (two lags) and Spain (not robust across the various lags considered) we find evidence in favour of the Cukierman–Meltzer hypothesis. Hence, these countries would be expected to gain significantly from EMU as the surrender of their monetary policy to the ECB would eliminate the policymakers' incentive to create inflation surprises. Finally, our evidence for the UK is rather mixed. At two and six lags we find evidence supporting the Cukierman–Meltzer hypothesis, and at four and eight lags evidence against the hypothesis. Our evidence for France (two lags) and the UK (two and six lags) squares with the findings of Grier and Perry (1998) and Baillie *et al.* (1996), respectively.

More independent central banks would have stronger anti-inflation preferences than the government and hence lead to a lower optimal inflation rate (Rogoff, 1985). Moreover, if inflation uncertainty is costly, i.e. it implies real output effects, and inflation can affect inflation uncertainty (the Fried-

[15]For the UK, the welfare cost of inflation was contained by the adoption of inflation targeting in 1992, following the country's brief participation in the EMS.

[16]For the Netherlands, Spain and possibly France, we find evidence that inflation uncertainty raises output growth, in particular when the full sample period 1960–99 is used. This seemingly surprising result may arise under the assumption of risk-averse agents and a precautionary motive for savings, as Dotsey and Sarte (2000) have shown in their theoretical model. According to their argument, when inflation uncertainty rises, savings increase and this boosts investment and growth.

[17]Our partial support arises from a lack of evidence for a negative impact of inflation uncertainty on output growth for these two countries and a lack of evidence for the Friedman hypothesis for Germany.

man hypothesis), an independent central bank will have a greater incentive (and freedom) to reduce inflation in response to more uncertainty. This is because in doing so (and hence keeping inflation uncertainty lower) the central bank can attain both lower inflation and higher output, i.e. a higher welfare level. The predictions of this analysis are borne out by the empirical evidence. Alesina and Summers (1993) show that more independent central banks are indeed associated with both lower inflation and inflation uncertainty.

Our results for the impact of inflation uncertainty on inflation are generally consistent with the existing literature on the rankings of central bank independence (see Alesina and Summers, 1993). Countries like France, Italy and Spain have less independent central banks than Germany and the Netherlands, at least using the measures of central bank independence that refer to the pre-1990 period, which is more in line with our sample period. Hence, we would expect that less independent central banks would be more likely to cause inflation surprises in response to higher inflation uncertainty, a result consistent with the Cukierman–Meltzer hypothesis. Our empirical analysis generally supports this prediction. Our conclusion on France and Germany also agrees with Grier and Perry (1998).

5 Conclusions

The relationship between inflation and inflation uncertainty has been investigated in six EU countries for the period 1960–99. EGARCH models were used to generate a measure of inflation uncertainty and then Granger methods were employed to test for causality between average inflation and inflation uncertainty. In all the European countries of our sample except Germany, inflation significantly raises inflation uncertainty, as predicted by Friedman. However, in all countries except the UK, inflation uncertainty does not cause negative output effects, implying that a common European monetary policy applied by the ECB might lead to asymmetric real effects via the inflation uncertainty channel.

Less robust evidence is found regarding the direction of the impact of a change in inflation uncertainty on inflation. In Germany and the Netherlands, increased inflation uncertainty lowers inflation, while in Italy, Spain and to a lesser extent France, increased inflation uncertainty raises inflation. These results are generally consistent with the existing rankings of central bank independence.

The reported differences in the results between this study and related studies, such as Grier and Perry (1998), can be attributed to the different methodologies, sample periods and data frequency. These differences highlight the need for further empirical work in search of more robust evidence on the relationship between inflation, inflation uncertainty and output growth. This work will provide an additional testing ground for the empirical relevance of economic theories and at the same time will be rather informative for the authorities in charge of monetary policymaking.

Appendix

This Appendix reports the estimation results of an EGARCH-M model of inflation in six countries with lagged inflation included in the conditional variance. As in the text, the estimation period is 1960–99 and the data frequency quarterly. We simultaneously estimate a system of equations that allows only the current value of either the conditional variance or the standard deviation of inflation[18] to affect average inflation and also allows up to the twelfth lag of average inflation to influence the conditional variance. The model includes the inflation equation which adds the inflation variance to the equation reported in the text

$$\pi_t = \tilde{\pi}_{t-1} + \delta h_{\pi t} + \varepsilon_t$$

and the conditional variance equation:

$$(1-\beta L)\ln(h_{\pi t}) = \omega + c|e_{t-1}| + de_{t-1} + k_i \pi_{t-i}$$

In the mean equation, $\tilde{\pi}_{t-1}$ stands for the part of the regression that includes the intercept and lagged inflation rates. In the variance equation, various lags of inflation (from 1 to 12) were considered with the best model chosen on the basis of the minimum value of the AIC.

Table A1 reports only the estimated parameters of interest. Note that when we estimate the model for the UK without the in-mean effect ($\delta = 0$), the coefficient for

Table A1

	EGARCH *level* k_i	*EGARCH* *in-mean* δ	*EGARCH* *level, in-mean* k_i	δ
UK	$k_4 = 23.67$	17.28	$k_4 = 25.20$	7.92
	(0.03)	(0.32)	(0.03)	(0.72)
France	$k_6 = 6.26$	177.18	$k_4 = 15.82$	80.39
	(0.00)	(0.00)	(0.05)	(0.00)
Germany	$k_4 = 18.63$	–46.85	$k_2 = 8.05$	–46.21
	(0.16)	(0.74)	(0.36)	(0.65)
Italy	$k_4 = 8.42$	–15.28	$k_6 = 3.40$	–12.94
	(0.06)	(0.40)	(0.08)	(0.44)
Netherlands	$k_6 = 3.21$	2.50	$k_{12} = 0.37$	–3.20
	(0.76)	(0.79)	(0.97)	(0.81)
Spain	$k_6 = 20.74$	13.87	$k_6 = 20.92$	–3.99
	(0.04)	(0.60)	(0.03)	(0.60)

Notes: (1) Probability values are given in parentheses.
(2) k_i indicates the estimated coefficient that corresponds to the ith lag in the inflation rate.

[18]According to the information criteria, the models with the variance of inflation were preferred to those with the standard deviation.

the effect of the fourth lag of inflation is 23.67 and is statistically significant. When we estimate the model without the level effect ($k_i = 0$), the in-mean coefficient is insignificant (the probability value is 0.32). When we estimate the model with the simultaneous feedback between the conditional variance and the conditional mean (last two columns in the table), the above results imply a positive association between lagged inflation and uncertainty similar to that found using the two-step method in the text. We do not find a significant effect of uncertainty on average inflation. However, as we emphasize in the text, such a result is plausible, as any relationship where uncertainty influences average inflation takes time to materialize and cannot be fairly tested in a model that restricts the effect to being contemporaneous.

A comparison of the results of the simultaneous estimation (last two columns in the table) with the Granger-causality results reported in the text for the rest of the countries indicates that, in general, there is consistency between the two approaches. In particular, as far as the Friedman hypothesis (significance of k_i) is concerned, we find the two approaches to be in agreement except in the case of the Netherlands. However, a comparison of the significance of the causal effect of inflation uncertainty on the inflation rate is not valid due to the contemporaneous nature of the effect under the simultaneous approach. As expected, the simultaneous approach does not detect such an effect in the majority of the countries considered.

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