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Dual Long Memory in Inflation Dynamics across Countries of the Euro Area and the Link between Inflation Uncertainty and Macroeconomic Performance

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Christian Conrad and Menelaos Karanasos

Abstract

This paper analyzes the inflation dynamics of several countries belonging to the European Monetary Union and of the UK. We estimate the two main parameters driving the degree of persistence in inflation and its uncertainty using a dual long memory process. We also investigate the possible existence of heterogeneity in inflation dynamics across Euro area countries and examine the link between nominal uncertainty and macroeconomic performance measured by the inflation and output growth rates. Strong evidence is provided for the hypothesis that increased inflation raises nominal uncertainty in all countries. However, we find that uncertainty surrounding future inflation has a mixed impact on output growth. This result brings out an important asymmetry in the transmission mechanism of monetary policy in Europe in addition to the difference in the economic sizes of the countries. We also investigate whether one can find a correlation between central bank independence and inflation policy. Our conclusion is that the most independent central banks are in countries where inflation falls in response to increased uncertainty.

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1 Introduction

An extensive body of theoretical literature examines the relationship between the rate of inflation and the nominal uncertainty. It is important to discover whether an increase in inflation precedes an increase in uncertainty, if we are to add to our understanding about the welfare costs of inflation. Different theories emphasize different channels, some pointing to a positive relationship and some to a negative one. Friedman (1977) argues that higher inflation may induce erratic policy responses to counter it, with consequent unanticipated inflation movements. In contrast, Pourgerami and Maskus (1990) point out that a negative effect may exist. The opposite direction of causality than that examined by Friedman has also been addressed in the theoretical literature. In particular, Cukierman and Meltzer (1986) contend that inflation uncertainty produces greater average inflation due to opportunistic central bank behavior, whereas according to Holland (1995) higher nominal uncertainty leads to lower average rates of inflation. Much empirical work has been done aimed at signing the effects of inflation on its uncertainty and vice versa. Contradictory empirical results are reported by various researchers. Given the theoretical ambiguity, it is not surprising that the statistical evidence is also ambiguous. Moreover, economic theory and empirical work reach a striking variety of conclusions about the responsiveness of output growth to changes in nominal uncertainty. The importance of uncertainty as a distinct channel in explaining the real effects of inflation has recently been given considerable empirical support (Grier et al., 2004, Elder, 2004, and Fountas and Karanasos, 2005). This channel was first highlighted by Friedman (1977). He argues that uncertainty about inflation causes an adverse growth effect. Dotsey and Sarte (2000) using a cash-in-advance framework obtain the opposite result: more nominal uncertainty can increase real growth.

This study has three primary objectives. First, it analyzes the inflation dynamics of several countries belonging to the European Monetary Union and of the UK. One group of countries is formed by Germany, France, Italy, the Netherlands and Spain. These five major countries represent 88 percent of the GDP of the Euro area. Given that the explicit mission of the European Central Bank (ECB) is the preservation of price stability, the analysis of the nature of inflation in the Euro area is of distinct interest. We estimate the two main parameters driving the degree of persistence in inflation and nominal uncertainty using an ARFIMA-FIGARCH process.[1] This model, developed in Baillie et al. (2002), provides a general and flexible framework with which to study a complicated process like inflation. Put differently, it is sufficiently flexible to handle the dual long memory behavior encountered in inflation.

Second, it investigates the possible existence of heterogeneity in inflation

[1]We refer to a model that is fractional integrated in both the ARMA and GARCH specifications as the ARFIMA-FIGARCH model.

dynamics across Euro area countries. Inflation differentials have important implications for the design of the optimal monetary policy. For example, as Benigno and Lòpez-Salido (2002) point out, an inflation targeting policy that assigns higher weight to countries with higher degrees of persistence benefits those countries since once the policy of the central bank is credible, it produces lower inflation rates for them simply because it cares more about those inflation rates.

The third objective of this study is to shed more light on the issue of temporal ordering of inflation and nominal uncertainty. To do this we proceed in two steps. First, we use the estimated conditional variance from the ARFIMA-FIGARCH model as our statistical measure of inflation uncertainty. Having constructed a time series of nominal uncertainty in the second part we employ Granger methods to test for evidence on the bidirectional causality relationship between inflation and uncertainty about inflation. The two-step approach has been employed among others by Grier and Perry (1998). In addition, we test for the causal effect of nominal uncertainty on output growth. The empirical evidence on this link remains scant or nonexistent, as pertains, in particular, to international data in European economies.

Our first finding is that all ten European inflation rates have the rather curious property of persistence in both their first and their second conditional moments. This empirical evidence is consistent with the evidence provided by Baillie et al. (2002) for eight industrial countries. The second result that emerges from this study is the existence of heterogeneity in inflation dynamics across Euro area countries. These countries fall into three groups in terms of the difference in the dynamics of the second moment of their inflation rates. The first group of countries includes France, Finland, the Netherlands and Sweden and is characterized by a mild long memory GARCH behavior of the inflation rate. The second includes Belgium, Italy and the UK, which are characterized by the presence of quite strong long memory in the inflation uncertainty. The third group of countries includes Portugal and Spain and is characterized by a near integrated behavior in the second conditional moment of the inflation rate. This finding is of some significance since inflation differentials are not irrelevant for monetary policy.

Third, we provide overwhelming evidence that increased inflation raises its uncertainty, confirming the theoretical predictions made by Friedman. However, we find that nominal uncertainty has a mixed impact on output growth. This result brings out an important asymmetry in the transmission mechanism of monetary policy in Europe in addition to the difference in the economic sizes of the countries. In particular, since the effects of uncertainty on output growth differ across the Euro zone, a common monetary policy that results in similar inflation rates across countries will have asymmetric real effects, provided these effects work via a change in nominal uncertainty. We also find that increased nominal uncertainty significantly affects average inflation in eight countries but not all in the same manner. These differential responses to nominal uncertainty are correlated with measures of central bank independence.

The remainder of the article is organized as follows. Section 2 summarizes several empirical studies that investigate the short-term inflation dynamics. Section 3 discusses the economic theory and the empirical testing concerning the link between inflation uncertainty and macroeconomic performance. In Section 4, we describe the time series model for inflation and nominal uncertainty and discuss its merits. The empirical results are reported in Section 5, and Section 6 draws some policy implications and proposes possible extensions of the time series model for inflation. Section 7 contains summary remarks and conclusions.

2 Inflation dynamics

This section summarizes several empirical studies that investigate the short-term inflation dynamics. The nature of the short-run inflation dynamics is a central issue in macroeconomics and one of the most fiercely debated. There is an extensive theoretical literature that attempts to develop structural models of inflation that provide a good approximation to its dynamics (see, for example, Karanassou and Snower, 2003), and an equally extensive empirical literature that attempts to document the properties of inflationary shocks. Many contradictory results can be found in the empirical literature on the persistence of inflation rates. Several studies (see, for instance, Grier and Perry, 1998) argue that inflation is I(0), whereas a large number of researchers, such as Banerjee and Russell (2001), find evidence for a unit root in inflation. Similarly, Ball and Cecchetti (1990) decompose inflation into a permanent component and a transitory component. As noted by Baillie et al. (1996) and Caporale and Gil-Alana (2003), the stationarity of real rates of interest and the Fisher relation is consistent with neoclassical models of dynamic growth, superneutrality, and capital asset pricing models. But if both inflation and nominal interest rates have a unit root then they must be cointegrated in order for the ex-post real rates to be stationary. Moreover, a nonstationary inflation process also complicates the derivation of optimal monetary policy rules (see McCallum, 1988).

Some researchers argue that inflation has become more persistent over time. In particular, Brunner and Hess (1993) model US inflation as an I(0) process before 1960 and as an I(1) process after this time. Along these lines, Evans and Watchel (1993) develop a time series model of inflation that switches from purely transitory shocks in the 1960s to purely permanent shocks in the 1970s, and back to transitory shocks in the late 1980s. They use this model to derive measures of nominal uncertainty that account for the prospects of changing inflation regimes. Generally speaking, as is often the case with post war data, one cannot say with confidence whether the two series, that is inflation and its uncertainty, are stationary or nonstationary or cointegrated if nonstationary. Accordingly, Holland (1995) performs three different tests for Granger causality between the two variables, each corresponding to one of the three different assumptions.

In sharp contrast, Backus and Zin (1993) find that a fractional root shows up very clearly in monthly US inflation. They conjecture that the long memory in inflation is the result of aggregation across agents with heterogeneous beliefs. They also demonstrate that the fractional difference process is a good descriptor of short-term interest rates and suggest that the fractional unit root in the short rates is inherited from money growth and inflation. Hassler and Wolters (1995) find that the inflation rates of five industrial countries are well explained by different orders of integration, which vary around the stationarity border of 0.5. Ooms and Hassler (1997) using data from Hassler and Wolters (1995) and a modified periodogram regression, confirm their findings. Subsequently, Baum et al. (1999) presented statistical evidence in favor of I(d) ($0 < d < 1$) behavior for both CPI-and WPI- based inflation rates for many industrial as well as developing countries.

The preceding works provide quite consistent evidence across time periods and countries that inflation exhibits long memory with an order of integration which differs significantly from 0 and 1. Overall these findings suggest that the traditional ARMA and ARIMA specifications are incapable of imparting the persistence to inflation that we find in the data. Put differently, by viewing inflation as an I(0) or I(1) process instead of an I(d) process, we bias downward or upward our estimate of inflation persistence. However, the previously mentioned articles have not explored the time-dependent heteroscedasticity in the second conditional moment of the inflation process. Along these lines, Baillie et al. (1996) utilize the ARFIMA-GARCH model to describe the inflation dynamics for ten countries. They provide strong evidence of long memory with mean reverting behavior for all countries except Japan. Hwang (2001) also estimates various ARFIMA-GARCH-type models for monthly US inflation. He finds strong evidence that inflation dynamics are well described by a fractional process with an order of integration of about 0.33.

In many applications the sum of the estimated GARCH(1,1) parameters is often close to one, which implies integrated GARCH (IGARCH) behavior. For example, Baillie et al. (1996) emphasize that for all ten countries the inflation series possesses substantial persistence in its conditional variance. In particular, the sum of the GARCH parameters was at least 0.965. Most importantly, Baillie, Bollerslev and Mikkelsen (1996), using Monte Carlo simulations, show that data generated from a process exhibiting long memory FIGARCH volatility may be easily mistaken for IGARCH behavior. Therefore recently Baillie et al. (2002) have focused their attention on the topic of long memory and persistence in terms of the second moment of the inflation process. They employ the FIGARCH specification of Baillie et al. (1996) to model the apparent long memory in the conditional variance of the inflation series. They find that the inflation rates for many industrial countries display significant fractional integration in both their first and second moments. Similarly, Conrad and Karanasos (2005) find that the ARFIMA-FIGARCH model was the preferred specification for the monthly CPI-based inflation rates for the UK and the US.

3 The link between inflation uncertainty and macroeconomic performance

3.1 Theory

In this Section, we discuss the economic theory concerning the link between nominal uncertainty and macroeconomic performance. Since Friedman (1977) stressed the harmful effects of nominal uncertainty on employment and production much research has been carried out investigating the relationship between inflation and uncertainty about inflation. The effect of inflation on its unpredictability is theoretically ambiguous. Several researchers contend that since a reduction in inflation causes an increase in the rate of unemployment, a high rate of inflation produces greater uncertainty about the future direction of government policy and the future rates of inflation. Ball's (1992) model formalizes this idea in the context of a repeated game between the monetary authority and the public.

Holland (1993) points out that in the Evans and Wachtel (1993) framework, if regime changes cause unpredictable changes in the persistence of inflation, then lagged inflation squared is positively related to nominal uncertainty. If, on the other hand, regime changes do not affect the persistence of inflation, then no relationship between the rate of inflation and its uncertainty is implied. In contrast, Pourgerami and Maskus (1990) suggest that higher inflation may induce the relevant economic agents to invest more in generating accurate predictions and hence may lead to lower nominal uncertainty. Ungar and Zilberfarb (1993) propose a mechanism that may weaken, offset, or even reverse the direction of the traditional view concerning the inflation-uncertainty relationship.

The models of Ball, Evans and Wachtel, and Holland imply that higher nominal uncertainty is part of the welfare costs of inflation because inflation causes its uncertainty. On the other hand, Cukierman and Meltzer (1986) and Cukierman (1992) using the Barro-Gordon model of Fed behavior show that greater uncertainty about money growth and inflation causes a higher mean rate of inflation by increasing the incentive for the policy-maker to create inflation surprises. In addition, Devereux (1989) emphasizes the fact that higher variability of real shocks lowers the optimal degree of indexation and increases the incentives of the policy maker to create surprise inflation. Therefore, if changes in the degree of indexation take time to occur then higher nominal uncertainty precedes greater inflation. In sharp contrast, Holland (1995) argues that due to the 'stabilization motive' higher nominal uncertainty has a negative effect on inflation.[2]

The impact of nominal uncertainty on output growth has received considerable attention in the theoretical macroeconomic literature. However, there is no consensus among macroeconomists on the direction of this effect. Theoretically speaking, the effect of uncertainty on growth is ambiguous. The second part of Friedman's hypothesis postulates that greater inflation variability, by reducing economic efficiency, has a negative impact on real growth. In particular, increased volatility in inflation rates reduces the ability of markets to convey information to market participants about relative price movements and makes long-term contracts more costly. Dotsey and Sarte (2000) analyze the effects of nominal uncertainty on economic growth in a model where money is introduced via a cash-in-advance constraint. In this setting, they find that variability increases average growth through a precautionary savings motive. Within the confines of their neoclassical growth model higher rates of inflation have negative consequences for growth, while increased inflation variability has a small positive effect on growth. In essence the offsetting growth effects of mean inflation and its uncertainty, along with the fact that these are highly correlated, provide a partial rationale for the weak and somewhat fragile relationship between growth and inflation. Finally, an alternative channel through which uncertainty about inflation might affect output growth is via the real uncertainty.[3] For example, a rising nominal uncertainty would be expected to have a positive impact on output growth via a combination of the Logue-Sweeney and Black effects.

3.2 Empirical evidence

In this Section, we discuss previous empirical testing of the link between nominal uncertainty and macroeconomic performance. The relationship between inflation and its uncertainty has been analyzed extensively in the empirical literature. Davis and Kanago (2000) survey this literature. Recent time series studies of nominal uncertainty have focused on the GARCH conditional variance of inflation as a statistical measure of its uncertainty. Some studies

[2]Under this scenario, if higher inflation raises its uncertainty, the policy maker responds by disinflating the economy in order to reduce nominal uncertainty and the associated costs.

[3]The positive association between inflation and output variability is known in the literature as the Logue-Sweeney hypothesis (see Karanasos and Kim, 2005, for details). The positive impact of output uncertainty on growth is known in the literature as the Black hypothesis (see Fountas and Karanasos, 2005, for details).

use GARCH models that include a function of the lagged inflation rate in the conditional variance equation. In particular, Brunner and Hess (1993) allow for asymmetric effects of inflation shocks on nominal uncertainty and find a weak link between US inflation and its uncertainty. Caporale and McKierman (1997) find a positive relationship between the level and variability of US inflation. Three studies use GARCH type models with a joint feedback between the conditional mean and variance of inflation. Baillie et al. (1996), for three high inflation countries and the UK, and Karanasos et al. (2004) for the US, find strong evidence in favor of a positive bidirectional relationship in accordance with the predictions of economic theory. In contrast, Hwang (2001) finds that US inflation affects its uncertainty weakly and negatively. Finally, the recent empirical literature tends to confirm the negative association between nominal uncertainty and real growth in the US. Studies by Grier and Perry (2000), Grier et al. (2004) and Elder (2004) employ bivariate GARCH in mean models and find a negative effect. However, all these studies use only US data.

Some studies examine the link between nominal uncertainty and the level of inflation using the two-step approach where an estimate of the conditional variance is first obtained from a GARCH-type model and Granger methods are then employed to test for bidirectional effects. In particular, Grier and Perry (1998) find that in all G7 countries inflation significantly raises its uncertainty. They also find evidence in favor of the Cukierman-Meltzer hypothesis for some countries and in favor of the Holland hypothesis for other countries. Fountas et al. (2004), using quarterly data and employing the EGARCH model, find that in five European countries inflation significantly raises its uncertainty. Their results regarding the direction of the impact of a change in nominal uncertainty on inflation were generally consistent with the existing rankings of central bank independence. Conrad and Karanasos (2005) for three industrial countries find strong evidence in support of both the Friedman and the Cukierman and Meltzer hypotheses.

Holland (1993) tabulates a number of empirical studies concerning the relationship between nominal uncertainty and real activity (employment or output). Studies based on surveys tend to find a negative relationship between nominal uncertainty and real activity, whereas studies based on ARCH volatility find insignificant or positive relationships. In particular, Coulson and Robins (1985) find that nominal uncertainty has a positive impact on real growth, while Jansen (1989) uses a bivariate ARCH in mean model and reports an insignificant relationship. Dotsey and Sarte (2000) also report empirical work documenting that the effect of uncertainty on growth appears to be non-negative. Grier and Tullock (1989) have been unable to verify the more conventional view that greater volatility in the inflation rate lowers growth, while McTaggart (1992) uses annual data for Australia and finds that inflation variability has a positive influence on the log of output. Levine and Renelt

(1992) use cross-country regressions to search for empirical linkages between growth rates and a variety of economic policy indicators. They find that all the results are fragile to small changes in the conditioning information set. The empirical findings in Barro (1996), for a panel of around 100 countries, support the notion that the variability of inflation has no significant relation with growth.

The empirical evidence on the relationship between nominal uncertainty and output growth remains scant or nonexistent, as pertains, in particular, to international data in industrialized economies. An exception is Fountas et al., (2004) and Fountas and Karanasos (2005). They employ the two-step approach in a univariate GARCH framework using data for six European and the G7 countries respectively and find significant evidence in favor of the Friedman hypothesis for some countries and in favor of the Dotsey-Sarte hypothesis for other countries. That is, the evidence regarding the direction of the impact of a change in nominal uncertainty on real growth found in these two studies is not robust across countries.

There are a limited number of studies using international data that are based on GARCH measures of inflation uncertainty. These are Baillie et al. (1996), Grier and Perry (1998), Fountas et al. (2004), Fountas and Karanasos (2005), and Conrad and Karanasos (2005). Only Fountas et al. (2004) investigate the relationship between inflation and its uncertainty for six European countries and only Conrad and Karanasos (2005) focus on a statistical measure of nominal uncertainty that captures the dual long memory aspect of inflation, namely the ARFIMA-FIGARCH (conditional) variance of inflation. This study aims to fill the gaps arising from the lack of interest in the European case, where the results would have interesting implications for the successful implementation of a common European monetary policy, and from the methodological shortcomings of the previous studies.

4 Methodology

4.1 The ARFIMA-FIGARCH process

It appears from the study of Baillie et al. (2002) that the apparent long memory in the inflation rate is also present in nominal uncertainty. Hence, there seems to be a need to have a joint model which incorporates long memory in both the conditional mean and the conditional variance of inflation. In other words, the time series features of inflation seem to require the use of fractional integrated models from two different classes, namely the ARMA and the GARCH.

Along these lines, in this Section we describe the time series model for inflation and nominal uncertainty and discuss its merits. First, we denote the inflation rate by π_t and we define its mean equation as

$$(1-L)^{d_m}(1-\phi_6 L^6-\phi_9 L^9-\phi_{12}L^{12}-\phi_{24}L^{24})(\pi_t-\mu)=\varepsilon_t, \qquad (1)$$

where $(1-L)^{d_m}$ is the fractional differencing operator with $d_m \leq 1$. That is the inflation rate follows an ARFIMA(24,d_m,0) specification. Second, let us suppose that ε_t is conditionally normal with mean zero and variance h_t. That is $\varepsilon_t|\Omega_{t-1} \sim N(0,h_t)$, where Ω_{t-1} is the information set up to time $t-1$. Finally, we assume that the structure of the conditional variance is given by a FIGARCH$(1,d,1)$, i.e.

$$(1-\beta L)h_t=\omega+\left[(1-\beta L)-(1-\alpha L)(1-L)^{d_v}\right]\varepsilon_t^2, \qquad (2)$$

where $0 \leq d_v \leq 1$, $\omega > 0$, and $\alpha,\beta < 1$. For necessary and sufficient conditions on the parameters (α,β,d) guaranteeing the nonnegativity of the conditional variance in the FIGARCH$(1,d,1)$ model see Conrad and Haag (2005). The FIGARCH specification reduces to a GARCH model for $d_v = 0$ and to an IGARCH model for $d_v = 1$. If $h_t = \omega$, a constant, the process reduces to the ARFIMA $(24,d_m,0)$ model. Then the inflation rate will be covariance stationary and invertible for $-0.5 < d_m < 0.5$ and will be mean reverting for $d_m < 1$. Although the ARFIMA-FIGARCH process is strictly stationary and ergodic for $0 \leq d_v \leq 1$, it will have an infinite unconditional variance for all d_m given a $d_v \neq 0$. Clearly, the unit root corresponds to the null hypothesis $H_0 : d_m = 1$.

The ARFIMA-FIGARCH model in (1) and (2) has a distinctive feature. It allows us simultaneously to estimate the degree of persistence in both inflation and nominal uncertainty. It also has the advantage of keeping the analytical elegance of the ARMA-GARCH model while enhancing its dynamics. Put differently, the ARFIMA-FIGARCH model has at least two important implications for our understanding of inflation and nominal uncertainty. First, it recognizes the long memory aspect of the inflation rate and provides an empirical measure of its uncertainty that accounts for long memory in the second conditional moment of the inflation process. Second, it allows for a more systematic comparison of many possible models that can capture the features of the inflation series.

4.2 Two-step strategy

To test for the relationship between inflation and nominal uncertainty, one can use either the two-step or the simultaneous approach. Under the latter one, an ARFIMA-FIGARCH-in-mean model is estimated with the conditional variance equation incorporating lags of the inflation series, thus allowing simultaneous estimation and testing of the bidirectional causality between the two variables. The two-step approach is performed by first obtaining an estimate of the conditional variance from the ARFIMA-FIGARCH model and then Granger methods are employed to test for bidirectional effects. We prefer the two-step strategy for the following reasons (see Grier and Perry, 1998). First, it allows us to capture the lagged effects between the variables of interest. In particular, the in-mean model suffers from the disadvantage that it does not allow the testing of a lagged effect of nominal uncertainty on inflation, which would be expected in a study that employs monthly 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. Second, the simultaneous approach is subject to the criticism of the potential negativity of the conditional variance. This is because there is no way of guaranteeing the non-negativity of the conditional variance by imposing constraints on the parameters of the conditional variance specification since the sign of the inflation series is time-varying. Third, the two-step approach minimizes the number of estimated parameters.

It is also interesting to note the similarities between the Lagrange Multiplier (LM) test for ARCH effects and the Granger causality methodology. In the LM statistic the first step is to estimate the conventional regression model for inflation by OLS (i.e., assuming independent errors) and obtain the fitted residuals ($\widehat{e}_t$). The second step is to regress $\widehat{e}_t^2$ on a constant and lags of $\widehat{e}_t^2$. If ARCH effects are present (i.e., if the squared errors are linearly related), the estimated parameters should be statistically significant. In our two-step strategy an estimate of the conditional variance is first obtained from an ARFIMA-FIGARCH model (i.e., assuming that inflation and its uncertainty are uncorrelated) and then causality tests are run to test for bidirectional effects between the two variables.

5 Empirical analysis

5.1 Data

In this Section we look at some of the time series characteristics of inflation. Monthly data, obtained from the *OECD Statistical Compendium*, are used to provide a reasonable number of observations. The inflation and output growth series are calculated as the monthly difference in the natural log of the Consumer Price Index and Industrial Production Index respectively. The data range from 1962:01 to 2004:01 and cover ten European countries, namely, Belgium, Finland, France, Germany, the Netherlands, Italy, Portugal, Spain, Sweden and the UK. Allowing for differencing this implies 504 usable observations.[4]

The summary statics (not reported) for the ten inflation rates show that the German (Portuguese) inflation rate has the lowest (highest) mean and standard deviation. Furthermore, the summary statistics indicate that the distributions of all ten inflation series are skewed to the right. The large values of the Jarque-Bera (JB) statistic imply a deviation from normality, and the significant Q-statistics of the squared deviations of the inflation rate from its sample mean indicate the existence of ARCH effects. This evidence is also supported by the Lagrange Multiplier (LM) test statistics, which are highly significant.

Next, we employ the PP and KPSS unit root tests, suggested by Phillips and Perron (1988) and Kwiatkowski et al. (1992) respectively. The results are presented in Table 1 and can be summarized as follows. For all the inflation series shown, based on the PP test we are able to reject the unit root hypothesis, whereas based on the KPSS test the null hypothesis of stationarity is rejected.[5] In other words, the application of these tests yields contradictory results. With all inflation series we find evidence against the unit root as well as against the stationarity hypothesis. Thus fractional integration allowing for long memory is a plausible model.[6] Finally, the results of the unit root tests applied to the output growth series (not reported) imply that we can treat these series as stationary processes.

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[4]The only exceptions are Belgium and Spain for which output data was available only from January 1965 onwards. For all countries the industrial production series are seasonally adjusted.

[5]We used a Bartlett kernel for the PP test and chose five as truncation lag, the number of lags included in the KPSS test was set to four. Alternatively, the Augmented Dickey-Fuller (ADF) statistic can be applied to test the unit root hypothesis. The evidence obtained from the PP test statistic is reinforced by the results (not reported) provided by the ADF tests.

[6]Of course, these unit root tests are merely suggestive. For example, Lee and Amsler (1997) show that the KPSS statistic cannot distinguish consistently between nonstationary long memory and unit root. We also examine the characteristics of inflation graphically by presenting the autocorrelation function of inflation and changes in inflation. Among other things, the figures (not reported) make clear the long memory property of inflation, that is the inflation series itself show significant positive and slowly decaying autocorrelations while the differenced series appear to follow an MA(1) process. Finally, we plot the autocorrelations of the squared and absolute values of the residuals from an estimated ARFIMA$(24, d_m, 0)$ model. Interestingly, these autocorrelations are extremely persistent, which is suggestive of long memory behavior in the conditional variance.

Table 1: Tests for order of integration of different countries' inflation series.

	PP		KPSS	
Country	$Z(t_\mu)$	$Z(t_\tau)$	$\hat{\eta}_\mu$	$\hat{\eta}_\tau$
Belgium	-14.24***	-14.93***	2.45***	0.83***
Finland	-17.00***	-18.31***	3.20***	0.81***
France	-10.29***	-11.48***	3.31***	1.46***
Germany	-16.24***	-16.55***	1.26***	0.34***
Italy	-8.30***	-8.71***	2.19***	1.41***
Netherlands	-21.23***	-22.07***	2.45***	0.34***
Portugal	-18.21***	-18.32***	1.63***	1.27***
Spain	-15.30***	-16.15***	2.80***	1.09***
Sweden	-17.66***	-18.12***	2.20***	1.06***
UK	-13.95***	-14.37***	1.88***	0.90***

Notes: $Z(t_\mu)$ and $Z(t_\tau)$ are the PP adjusted t-statistics of the lagged dependent variable in a regression with intercept only, and intercept and time trend included, respectively. The 0.01 critical values for $Z(t_\mu)$ and $Z(t_\tau)$ are -3.43 and -3.96. $\hat{\eta}_\mu$ and $\hat{\eta}_\tau$ are the KPSS test statistics based on residuals from regressions with an intercept and intercept and time trend, respectively. The 0.01 critical values for $\hat{\eta}_\mu$ and $\hat{\eta}_\tau$ are 0.739 and 0.216. *** denotes significance at the 0.01 level.

5.2 Model estimates

The analysis in Baillie et al. (2002) suggests that the ARFIMA-FIGARCH specification describes the inflation series of several industrial countries well. Within this framework we will analyze the dynamic adjustments of both the conditional mean and the conditional variance of inflation for several European countries, as well as the implications of these dynamics for the direction of causality between nominal uncertainty and macroeconomic performance. Estimates of the ARFIMA-FIGARCH model are shown in Table 2. These were obtained by quasi-maximum likelihood estimation (QMLE) as implemented by Laurent and Peters (2002) in Ox. The truncation lag for the fractional differencing operator was chosen to be 500. For a detailed description of the estimation procedure see Baillie, Bollerslev and Mikkelsen (1996). The consistency and asymptotic normality of the QMLE has been established only for specific special cases of the ARFIMA and/or FIGARCH model. However, a detailed Monte Carlo study, where ARFIMA-FIGARCH type models were simulated, was performed by Baillie et al. (2002) and it was found that the quality of the application of the QMLE is generally very satisfactory. To check for the robustness of our estimates we used a range of starting values and hence ensured that the estimation procedure converged to a global maximum.

Several findings emerge from Table 2. The estimated long memory conditional mean parameter is in the range $0.141 \leq \widehat{d}_m \leq 0.353$. The value of the coefficient for Portugal (0.141) is markedly lower than the corresponding value for Italy (0.353). However, although the estimated value of d_m for Portugal is relatively small it is significantly different from zero. Furthermore, once long memory in the conditional mean has been accounted for, an AR(24) specification appears to capture the serial correlation in all ten inflation series. That is, all the $\widehat{\phi}_{12}$ and $\widehat{\phi}_{24}$ parameters are much larger than their standard errors.

The estimation of a FIGARCH model for Portugal and Spain realized an estimated value of d_v close to 0.9 (0.874 and 0.866 respectively), whereas in sharp contrast, for France and Sweden it realized a value close to 0.1 (0.130 and 0.133 respectively). In other words, the estimates of d_v that govern the dynamics of the conditional heteroscedasticity indicate that the conditional variances of the Portuguese and Spanish inflation are characterized by a near integrated GARCH behavior, whereas the conditional variances of the French and Swedish inflation are characterized by a very mild long memory GARCH behavior. For the other six countries, the values of d_v vary from 0.195 (Netherlands), 0.209 (Finland), and 0.269 (Germany) to 0.330 (Belgium), 0.457 (UK), and 0.529 (Italy). For Finland, France, Germany, the Netherlands, Sweden and the UK the Akaike and Schwarz information criteria (AIC and SIC respectively) come out in favor of the FIGARCH$(0, d_v, 0)$ model, while for Italy, Portugal and Spain (the three countries with the highest $\widehat{d}_v$) the FIGARCH$(1, d_v, 0)$ is the preferred specification. In addition, note that the estimated GARCH parameters for these three countries and for Belgium satisfy the set of conditions which are necessary and sufficient to guarantee the nonnegativity of the conditional variance (see Conrad and Haag, 2005).

The ten European countries fall into three groups in terms of the differences in the sum of the two fractional differencing parameters $(d_m + d_v)$. The first group of countries includes Finland, France, Germany, the Netherlands and Sweden: $0.310 < \widehat{d}_m + \widehat{d}_v < 0.480$. In all these countries, except France, the estimated value of d_m is very close to the estimate of d_v. The second includes Belgium, Italy and the UK: $0.540 < \widehat{d}_m + \widehat{d}_v < 0.890$. The third group of countries includes Portugal and Spain: $\widehat{d}_m + \widehat{d}_v \simeq 1$. Portugal and Spain have very similar estimated mean (0.141, 0.181) and variance (0.874, 0.866) fractional differencing parameters.[7] Interestingly, these are the two countries with the lowest (highest) long memory mean (variance) parameters. Whether the sum of the two estimated fractional differencing parameters is below or above 0.5 will become of importance when analyzing causal relationships between inflation and its uncertainty in the next section. In seven out of the ten countries the estimates of d_m are smaller than the estimates of d_v.

[7]The Portugese and Spanish inflation series are also the two series with the highest sample means and standard deviations.

Table 2: ARFIMA-FIGARCH models.

	Belgium	Finland	France	Germany	Italy
$\widehat{\mu}$	0.282	0.125	0.314	0.254	0.186
	(1.984)	(1.258)	(1.440)	(2.510)	(1.186)
$\widehat{d}_m$	0.214	0.189	0.305	0.203	0.353
	(3.481)	(6.169)	(6.712)	(5.049)	(4.124)
$\widehat{\phi}_{12}$	0.249	0.283	0.212	0.376	0.300
	(4.988)	(4.071)	(3.415)	(7.914)	(6.769)
$\widehat{\phi}_{24}$	0.092	0.156	0.185	0.253	0.191
	(1.882)	(2.700)	(4.064)	(4.925)	(4.646)
$\widehat{\omega}$	0.013	0.038	0.021	0.014	0.001
	(2.013)	(1.893)	(1.873)	(2.454)	(0.517)
$\widehat{d}_v$	0.330	0.209	0.130	0.269	0.529
	(2.311)	(4.810)	(1.678)	(2.077)	(3.371)
Q_{12}	15.65	8.74	13.94	13.78	21.26
	[0.21]	[0.72]	[0.30]	[0.31]	[0.05]
Q_{12}^2	10.54	5.28	12.03	5.07	14.27
	[0.57]	[0.95]	[0.44]	[0.96]	[0.28]

	Netherlands	Portugal	Spain	Sweden	UK
$\widehat{\mu}$	0.293	0.163	0.361	0.413	0.339
	(1.932)	(1.050)	(1.836)	(2.885)	(0.831)
$\widehat{d}_m$	0.203	0.141	0.181	0.185	0.340
	(4.093)	(3.413)	(3.185)	(4.202)	(4.414)
$\widehat{\phi}_{12}$	0.507	0.389	0.334	0.261	0.401
	(6.584)	(7.578)	(5.027)	(4.547)	(7.471)
$\widehat{\phi}_{24}$	0.173	0.180	0.263	0.216	0.330
	(3.545)	(3.681)	(4.485)	(3.713)	(6.310)
$\widehat{\omega}$	0.021	0.002	0.003	0.106	0.019
	(1.788)	(0.420)	(1.702)	(1.939)	(1.666)
$\widehat{d}_v$	0.195	0.874	0.866	0.133	0.457
	(3.685)	(10.82)	(5.854)	(1.760)	(4.056)
Q_{12}	16.24	17.93	20.01	15.92	14.36
	[0.18]	[0.12]	[0.07]	[0.19]	[0.28]
Q_{12}^2	14.55	6.04	9.66	18.74	13.29
	[0.27]	[0.91]	[0.65]	[0.09]	[0.35]

Notes: For each of the ten inflation series, Table 2 reports QML parameter estimates for the ARFIMA-FIGARCH model. The numbers in parentheses are t-statistics. Q_{12} and Q_{12}^2 are the 12th order Ljung-Box tests for serial correlation in the standardized and squared standardized residuals respectively. The numbers in brackets are p values. The ϕ_6 and ϕ_9 coefficients are significant only in Belgium and France. For Italy, Portugal and Spain we estimate a β of 0.266(1.010), 0.772(9.878) and 0.724(6.129) respectively. The α coefficient is significant only in Belgium: $\widehat{\alpha} = -0.280(-1.933)$.

Generally speaking, the parameter estimates support the idea that dual long memory effects are present in the inflation process for all ten European countries, suggesting that the dual long memory is an important characteristic of the inflation data. Finally, with all countries, the hypothesis of uncorrelated standardized and squared standardized residuals is well supported, indicating that there is no statistically significant evidence of misspecification.

To test for the persistence of the conditional heteroscedasticity models, we examine the likelihood ratio (LR) tests and the Wald (W) statistics for the linear constraints $d_m = d_v = 0$ (ARMA-GARCH model). As seen in Table 3 the LR tests and W statistics clearly reject the ARMA-GARCH null hypotheses against the ARFIMA-FIGARCH model for all ten inflation series. Thus, purely from the perspective of searching for a model that best describes the degree of persistence in both the mean and the variance of the inflation series, the ARFIMA-FIGARCH model appears to be the most satisfactory representation.

Table 3: LR and W test statistics.

	Be	Fi	Fr	Ge	It	Ne	Po	Sp	Sw	UK
LR	16.95	47.73	39.64	42.88	113.88	58.22	18.62	17.69	8.03	33.98
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.02]	[0.00]
W	6.72	29.92	29.12	12.93	9.10	12.31	71.22	19.35	8.89	15.52
	[0.03]	[0.00]	[0.00]	[0.00]	[0.01]	[0.00]	[0.00]	[0.00]	[0.01]	[0.00]

Notes: For each of the ten inflation series the first row of Table 3 reports the value of the following LR test: LR=2[ML$_u$-ML$_r$], where ML$_u$ and ML$_r$ denote the maximum log-likelihood values of the unrestricted (ARFIMA-FIGARCH) and restricted (ARMA-GARCH) models respectively. Row 2 reports the corresponding W statistics. The numbers in brackets are p-values. Be: Belgium, Fi: Finland Fr: France, Ge: Germany, It: Italy, Ne: Netherlands, Po: Portugal, Sp: Spain, Sw: Sweden.

Following the work of Grier and Perry (1998) among others, the LR test can be used for model selection. Alternatively, the AIC, SIC and Hannan-Quinn or Shibata information criteria (HQIC, SHIC respectively) can be applied to rank the various ARMA-GARCH type models. These model selection criteria check the robustness of the LR and W testing results discussed above.[8] According to the four information criteria, in all ten cases the optimal GARCH type model is the ARFIMA-FIGARCH.[9] Hence, the model selection criteria are in accordance with the LR and W testing results. Furthermore, we should

[8]The analysis in Caporin (2003) focuses on the identification problem of FIGARCH models. Caporin performs a detailed Monte Carlo simulation study and shows that the four information criteria can clearly distinguish between long and short memory data generating processes. Finally, Caporin's results show that when LR tests are applied to time series for which the true data generating process is FIGARCH with a d_v parameter being close to one, the LR test has no power to distinguish between the fractionally integrated and the IGARCH model.

[9]We do not report the AIC, SIC, HQIC or SHIC values for space considerations.

also mention that although the estimated d_v parameter for Portugal and Spain is not significantly different from unity, it appears that the volatility dynamics in these two countries are better modelled by the fractional differencing parameter since both the LR and W statistics (not reported) clearly reject the 'IGARCH' hypothesis against the FIGARCH model. In addition, the information criteria favor the FIGARCH model over the IGARCH model.

Finally, we test for the similarity of the optimal mean fractional differencing parameters estimated for each of the ten inflation series using a pairwise Wald test:

$$\mathrm{W}_m = \frac{(\widehat{d}_{m,1} - \widehat{d}_{m,2})^2}{(\mathrm{SE}_{m,1})^2 + (\mathrm{SE}_{m,2})^2},$$

where $\widehat{d}_{m,i}$, $(i = 1, 2)$ is the mean fractional differencing parameter from the ARFIMA-FIGARCH model estimated for the inflation series for country i and $\mathrm{SE}_{m,i}$ is the standard error associated with the estimated model for country i. The above W statistic tests whether the mean fractional differencing parameters of the two countries are equal $(\widehat{d}_{m,1} = \widehat{d}_{m,2})$, and is distributed as $\chi^2_{(1)}$. In the majority of the cases the results (not reported) of this pairwise testing procedure provide support for the null hypothesis that the estimated fractional parameters are not significantly different from one another.[10]

5.3 Granger causality tests

In this section we report results of Granger causality tests to provide some statistical evidence on the nature of the relationship between nominal uncertainty and macroeconomic performance. Tsay and Chung (2000) in their analysis of spurious regression with independent, fractionally integrated processes find that in bivariate regressions no matter whether the dependent variable and the regressor are stationary or not, as long as their fractional orders of integration sum up to a value greater than 0.5, the t ratios become divergent. Recall that in five countries (Belgium, Italy, Portugal, Spain and the UK) the estimated sum of the two long memory parameters $(d_m + d_v)$ is greater than 0.5.

Consequently, we utilize the methodology developed by Toda and Yamamoto (1995) to test for causality between nominal uncertainty and either inflation or output growth, which leads to a χ^2 distributed test statistic despite any possible nonstationarity or cointegration between the series.[11] The

[10] It should be noted that the mean fractional differencing parameters are related to the other parameters in the ARFIMA-FIGARCH model. In particular, the information matrix between the AR parameters and the fractional parameter is not block diagonal. Hence, comparison of estimated d_m parameters, specially between countries with different model specifications, should be taken with a pinch of salt.

[11] Note, that this procedure also avoids the problem of unbalanced regression, which could occur in regressions involving the I(0) output series and the near-integrated conditional

test is performed in two steps. In the first step, the optimal lag length (k) of the system is determined by utilizing the AIC and SIC. In the second step a VAR of order $k^* = k + d_{\max}$ is estimated (where $d_{\max}$ is the maximal (integer) order of integration suspected to occur in the system) and a modified W (MW) test is applied to the first k VAR coefficient matrices to make Granger causal inference. This MW test statistic has an asymptotic χ^2 distribution with k degrees of freedom. Since inflation and its uncertainty are fractionally integrated with $d_m, d_v < 1$ we set $d_{\max} = 1$ and estimate VAR models with $k^* = k + 1$ lags.[12] The optimal lag length turned out to be either 4, 8 or 12 for all countries. To ensure that our results are not sensitive to the choice of the lag length we report in Table 4 for all ten countries the MW tests using 4, 8 and 12 lags, as well as the sign of the sums of lagged coefficients in case of significance.

Panel A reports results on the impact of changes in inflation on its uncertainty. We apply the MW tests and use the Newey-West heteroscedasticity and autocorrelation consistent standard errors. Statistically significant effects are present for all countries. There is strong evidence that inflation affects its uncertainty positively, as predicted by Friedman (1977) and Ball (1992).

We then perform Granger causality tests in order to examine the causal effect of nominal uncertainty on macroeconomic performance. The tests are performed under the assumption that the conditional variances follow GARCH-type processes.[13] Panel B reports the results of the causality tests where causality runs from nominal uncertainty to the rate of inflation. This Panel shows a significant positive effect of uncertainty on inflation for three out of the ten countries. The evidence is strong for France and Spain and weaker for Portugal, where it applies for only one of the chosen lags. The results from these three countries support the Cukierman-Meltzer hypothesis that Grier and Perry (1998) label as the 'opportunistic Fed'. Increases in nominal uncertainty raise the optimal average inflation by increasing the incentive for the policy-maker to create inflation surprises. For Sweden, we find strong evidence

variance series of Portugal and Spain.

[12]Of course, the Toda and Yamamoto (1995) procedure is inefficient and suffers some loss of power since one intentionally over-fits the model. However, if – as in our case – the VAR system has only two variables and long lag length, the inefficiency caused by adding only one more lag is expected to be relatively small.

[13]In the presence of conditional heteroskedasticity Vilasuso (2001) investigates the reliability of causality tests based on least squares. He demonstrates that when conditional heteroskedasticity is ignored, least squares causality tests exhibit considerable size distortion if the conditional variances are correlated. In addition, inference based on a heteroskedasticity and autocorrelation consistent covariance matrix constructed under the least squares framework offers only slight improvement. Therefore, he suggests that causality tests be carried out in the context of an empirical specification that models both the conditional means and conditional variances. However, if the conditional variances are unrelated, then there is only slight size distortion associated with least-squares tests, and the inconsistency of the least squares standard errors is unlikely to be problematic.

for a negative effect of nominal uncertainty on inflation, which along with the growth effect of nominal uncertainty squares with Holland's stabilization hypothesis. In other words, this result suggests that the 'stabilizing Fed' notion is plausible. Increased inflation raises uncertainty, which creates real welfare losses and then leads to monetary tightening to lower inflation and thus also uncertainty (see Grier and Perry, 1998). For Finland there is evidence for Holland's argument at lag 8 only. We also obtain mixed evidence for Germany, the Netherlands and the UK. In particular, at eight lags uncertainty has a positive impact on inflation, whereas the value of the MW test statistic and the sign of the sum of lagged coefficients at 12 lags (optimal lag length) implies a negative relationship. We view this as support for Holland's stabilization hypothesis. Since monetary policy takes time to materialize, it is not surprising that a negative effect is found at 12 lags, but not at 4 or 8 lags. A time horizon of 3 to 4 quarters is what one would usually expect for monetary policy to effect the economy. However, neither of the two theories is supported in Belgium and Italy, where inflation is independent from changes in its uncertainty. Thus, increased uncertainty significantly affects future inflation in most of the countries in the sample, but not all in the same manner.

The Granger causality test results of uncertainty on real growth are given in Panel C. As we show above high-inflation countries are also likely to experience highly volatile inflation rates. If only uncertainty is included in the estimated regression equations, it is impossible to determine whether it is the inflation rate or its uncertainty that is affecting output growth. Hence, in order to control for possible effects of uncertainty on growth that take place via changes in inflation Panel C reports the MW statistics when the regressions include in addition lagged inflation rates. Nominal uncertainty has a mixed impact on output growth. Friedman's hypothesis regarding the negative real effects of uncertainty receives support in five out of the ten countries. The evidence is strong in Belgium, Sweden and the UK, mild in Italy, and weaker in Germany where it applies for only one of the chosen lags. In contrast, in the other five countries we find that uncertainty has a positive impact on real growth, supporting the Dotsey-Sarte hypothesis. The evidence is strong in Finland, France and the Netherlands, and mild in Portugal and Spain. The fact that many other factors are likely to be related to output growth-either causally or because both are influenced by a third factor makes it more difficult to gauge the significance and magnitude of the impact of uncertainty on growth. Therefore, one should be wary of putting too much faith in the uncertainty-growth relationship. But at the broadest level, the available evidence supports the Friedman hypothesis in some countries and is in favor of the Dotsey-Sarte hypothesis for other countries.

Moreover, note that for France, Portugal and Spain we find evidence for a positive effect of nominal uncertainty on inflation, which along with the output effect of inflation uncertainty squares with the Cukierman-Meltzer ('opportunistic') hypothesis. In other words, the central banks dislike inflation but value the higher employment that results from surprise inflation. Therefore, increases in nominal uncertainty raise the average inflation rate by increasing the incentive for the policy-makers to create inflation surprises (Grier and Perry, 1998).

Table 4: Granger-causality tests between inflation, output growth and inflation uncertainty.

	Belgium	Finland	France	Germany	Italy
Panel A. H_0: Inflation does not Granger-cause inflation uncertainty.					
4	**11.51**[0.02](+)	**17.37**[0.00](+)	17.04[0.00](+)	**15.65**[0.00](+)	**21.62**[0.00](+)
8	17.14[0.03](+)	25.81[0.00](+)	28.79[0.00](+)	17.57[0.02](+)	22.57[0.00](+)
12	20.61[0.06](+)	27.89[0.00](+)	**35.32**[0.00](+)	21.28[0.05](+)	24.90[0.01](+)
Panel B. H_0: Inflation uncertainty does not Granger-cause inflation.					
4	3.12[0.54]	**4.00**[0.40]	**7.42**[0.11](+)	5.89[0.21]	**4.93**[0.29]
8	**4.06**[0.85]	17.07[0.03](-)	26.46[0.00](+)	22.22[0.00](+)	10.90[0.21]
12	6.72[0.87]	15.96[0.19]	28.50[0.00](+)	**25.68**[0.01](-)	12.96[0.37]
Panel C. H_0: Inflation uncertainty does not Granger-cause output growth.					
4	**7.22**[0.12](-)	**7.27**[0.12](+)	**15.43**[0.00](+)	**0.86**[0.93]	**8.10**[0.09](-)
8	37.47[0.00](-)	20.44[0.01](+)	20.47[0.00](+)	5.70[0.68]	10.53[0.23]
12	25.22[0.01](-)	36.56[0.00](+)	70.42[0.00](+)	59.81[0.00](-)	22.34[0.03](-)

	Netherlands	Portugal	Spain	Sweden	UK
Panel A. H_0: Inflation does not Granger-cause inflation uncertainty.					
4	8.75[0.07](+)	**16.09**[0.00](+)	**13.02**[0.01](+)	**43.83**[0.00](+)	**54.45**[0.00](+)
8	19.01[0.01](+)	27.28[0.00](+)	18.51[0.02](+)	57.96[0.00](+)	17.87[0.02](+)
12	**13.38**[0.34]	35.02[0.00](+)	24.71[0.02](+)	52.58[0.00](+)	36.51[0.00](+)
Panel B. H_0: Inflation uncertainty does not Granger-cause inflation.					
4	21.84[0.00](+)	5.63[0.23]	6.73[0.15](+)	**17.11**[0.00](-)	4.82[0.31]
8	20.65[0.01](+)	**8.96**[0.34]	17.83[0.02](+)	21.74[0.00](-)	54.37[0.00](+)
12	**22.05**[0.04](-)	17.14[0.14](+)	**39.64**[0.00](+)	26.08[0.01](-)	**67.12**[0.00](-)
Panel C. H_0: Inflation uncertainty does not Granger-cause output growth.					
4	**8.86**[0.06](+)	**1.37**[0.84]	**8.06**[0.09](+)	**11.53**[0.02](-)	**38.77**[0.00](-)
8	16.71[0.03](+)	13.67[0.09](+)	8.87[0.35]	28.30[0.00](-)	48.69[0.00](-)
12	34.96[0.00](+)	40.96[0.00](+)	19.00[0.09](+)	32.49[0.00](-)	68.56[0.00](-)

Notes: The figures are MW statistics. The numbers in the first column give the lag structure for the MW tests, i.e. the orders of the VAR's are 5, 9 and 13. The bold (underlined) numbers indicate the optimal lag length chosen by SIC(AIC). The numbers in [·] are p values. A +(-) indicates that the sum of the lagged coefficients is positive (negative).

The three Figures in the Appendix plot (for Germany, the Netherlands and the UK) (i) the time profiles of inflation and its uncertainty due to shocks in nominal uncertainty and inflation respectively and (ii) the time profile of output growth due to shocks in nominal uncertainty.[14] The maximum effect of inflation on its uncertainty takes place after three (two) months for the Netherlands (Germany and the UK). The negative impact of nominal uncertainty on output growth reaches its peak after nine and twelve months in Germany and the UK respectively. In contrast, in the Netherlands the maximum (positive) effect takes place after five months. Finally, the sign of the effect of nominal uncertainty on inflation varies considerably over time. In all three countries the negative impact reaches its peak after twelve months. In Germany and the Netherlands the effect also seems much smaller in size than the effect of inflation uncertainty on real growth.

To summarize, the results in this section confirm that inflation affects its uncertainty positively. Uncertainty surrounding future inflation appears to have a mixed impact on both inflation and output growth.

5.4 Robustness

5.4.1 Monte-Carlo study

To check the sensitivity of our results to the orders of integration of inflation (d_m) and its uncertainty (d_v), we are also using the inflation series filtered by $(1-L)^{\widehat{d}_m}$ and the series of the estimated conditional variances filtered by $(1-L)^{\widehat{d}_v}$. We carry out the conventional Granger causality tests using both sets of data, i.e., our original set of data and the one with the two filtered series. If significant effects are obtained for the original series, but not when applying the Toda and Yamamoto (1995) procedure (or using the appropriately differenced series), this could be viewed as evidence for spurious regression in the simple Granger causality tests. The results (not reported) are very similar to those obtained using the methodology developed in Toda and Yamamoto (1995).[15] In particular, when the original data are used the primary difference lies in the stronger evidence on the Cukierman-Meltzer hypothesis for Portugal at 4 lags. The main difference when the filtered series are used is that now no evidence appears for the Dotsey-Sarte hypothesis in Spain.

Since the results from the simple Granger-causality tests and those obtained by the Toda and Yamamoto (1995) procedure are basically identical, it seems that hardly any spurious effect due to the fractionally integrated variables occurs in our setting. At first sight this result seems to be at odds with the findings of Tsay and Chung (2000) who have shown that regressions involving fractionally integrated regressors can lead to spurious results. In

[14] Generalized impulse response functions are calculated as suggested in Pesaran and Shin (1998). We do not report Figures for the other countries for space considerations.

[15] Since we also apply the Toda and Yamamoto (1995) procedure to inflation series for which $\widehat{d}_m + \widehat{d}_v < 0.5$, we should mention that in all these cases the results from the two methodologies were qualitatively identical.

particular, analyzing the bivariate regression of y_t on a constant and x_t where $y_t \sim \mathrm{I}(d_y)$ and $x_t \sim \mathrm{I}(d_x)$ they show that the corresponding t-statistic will be divergent provided $d_y + d_x > 0.5$.

We illustrate that their result does not apply to our setting in a small Monte-Carlo study by simulating the critical values of causality tests which are performed for two independent series having the same orders of fractional integration as the estimated ones for the UK and Portugal.[16] Recall that both countries satisfy $\widehat{d}_m + \widehat{d}_v > 0.5$. The simulation is performed in the following way:

Step 1. We generate two independent series

$$\begin{aligned} \tilde{\pi}_t &= (1-L)^{-\widehat{d}_m}\tilde{\varepsilon}_t^{\pi} = \sum_{j=0}^{500} \psi_j^{\pi}\tilde{\varepsilon}_{t-j}^{\pi} \qquad t = 1, \ldots, 504, \\ \tilde{h}_t &= (1-L)^{-\widehat{d}_v}\tilde{\varepsilon}_t^{h} = \sum_{j=0}^{500} \psi_j^{h}\tilde{\varepsilon}_{t-j}^{h} \qquad t = 1, \ldots, 504, \end{aligned}$$

where $\widehat{d}_m$ and $\widehat{d}_v$ are the estimated orders of fractional integration and $\tilde{\varepsilon}_t^{\pi}$ and $\tilde{\varepsilon}_t^{h}$ are $iid \sim N(0,1)$. Hence, $\tilde{\pi}_t$ and $\tilde{h}_t$ are integrated of order $\widehat{d}_m$ and $\widehat{d}_v$, respectively, and satisfy the assumptions made in Tsay and Chung (2000).

Step 2. For the generated sample $\{\tilde{\pi}_t, \tilde{h}_t\}$ we run the following regressions:

$$\tilde{\pi}_t = \beta_0 + \beta_1^h \tilde{h}_t + \eta_t^{\pi}, \tag{3}$$

$$\tilde{\pi}_t = \beta_0 + \sum_{j=1}^{k} \beta_j^h \tilde{h}_{t-j} + \sum_{j=1}^{k} \beta_j^{\pi} \tilde{\pi}_{t-j} + \eta_t^h \qquad \text{for} \quad k = 4, 8, 12, \tag{4}$$

and calculate the corresponding value of the test-statistic ($H_0 : \beta_j^h = 0,\ j = 1, \ldots, k$). Equation (3) corresponds to the setting described in Tsay and Chung (2000), while equation (4) is our setting from table 4, panel B. Repeating step 1 and 2 for M=10000 times we approximate the distribution of the test-statistic. From the simulated distribution we calculate the 5% and 1% critical values.

The theoretical results derived in Tsay and Chung (2000) suggest that spurious regression occurs in equation (3), but what about equation (4) ? The simulation results presented in table 5 show that spurious regression is a much more severe problem in the bivariate case considered by Tsay and Chung (2000) than in regressions including lagged dependent and independent variables with 4 or more lags. In the bivariate case the 5% critical value according to the F distribution would be 3.86, while the simulated critical values are 17.91

[16] Note, that the intention of the simulation is to show that the results of Tsay and Chung (2000) do not carry over to our setting and not to obtain the correct critical values for the Granger causality tests using the original series. For this one would have to generate two series according to the equations (1) and (2) by using the parameter estimates from table 2 and drawing innovations from the estimated standardized residuals.

($\widehat{d}_m = 0.34, \widehat{d}_v = 0.46$) and 10.91 ($\widehat{d}_m = 0.14, \widehat{d}_v = 0.87$). Hence, relying on the F critical value in the bivariate regression would lead to rejection of the null hypothesis of no Granger causality in by far too many cases. However, the more lags are included, the less different are the critical values of the F-distribution and the simulated distribution. In particular, the 5% critical value for 12 lags is 1.77 for the F-distribution, while the simulated ones are 1.82 ($\widehat{d}_m = 0.34, \widehat{d}_v = 0.46$) and 1.83 ($\widehat{d}_m = 0.14, \widehat{d}_v = 0.87$). Since the difference between the critical values for the F and the simulated distributions are very small in our setting the influence of spurious regression does not seem to play an important role. This explains why the simple Granger causality results are very similar to those obtained by using the Toda-Yamamoto methodology.

Table 5: Simulated critical values.

	$d_m = 0, d_v = 0$ $F(k, 504-k)$		$\widehat{d}_m = 0.34, \widehat{d}_v = 0.46$ $F^\star$		$\widehat{d}_m = 0.14, \widehat{d}_v = 0.87$ $F^\star$	
	5%	1%	5%	1%	5%	1%
Bivariate	3.86	6.69	17.91	34.42	10.91	18.46
$k = 4$	2.39	3.35	2.73	3.75	2.81	3.85
$k = 8$	1.96	2.54	2.05	2.65	2.10	2.71
$k = 12$	1.77	2.22	1.82	2.26	1.83	2.30

Notes: $F^\star$ is the simulated distribution of the test-statistic using M=10000 generated samples. The figures are 5% and 1% critical values. k denotes the lag length in the Granger causality test.

5.4.2 Simultaneous approach

This section reports the estimation results of an ARFIMA-FIGARCH-in-mean model with lagged inflation included in the variance specification (the so called level effect). We estimate a system of equations that allows only the current value of the conditional variance to affect average inflation and up to the fifth lag of average inflation to influence the conditional variance. In other words, the model includes the mean equation which adds the variance of inflation (δh_t) to the expressions reported in Table 2, and the variance equation augmented by the term $\gamma_i \pi_{t-i}$. In the expressions for the conditional variances reported in Table 2, various lags of inflation (from 1 to 5) were considered with the best model chosen on the basis of the minimum value of the AIC. Table 6 reports only the two estimated parameters of interest.

Table 6: ARFIMA-FIGARCH-in-mean-level models.

	Belgium	Finland	France	Germany	Italy
$\widehat{\gamma}_i$	–*	0.097 [4]	0.028 [5]	–	0.003 [2]
		(3.066)	(1.811)		(1.830)
$\widehat{\delta}$	–	0.143	-1.203	–	0.299
		(0.689)	(1.405)		(1.017)

	Netherlands	Portugal	Spain	Sweden	UK
$\widehat{\gamma}_i$	0.058 [4]	–	–	0.153 [4]	0.048 [5]
	(3.825)			(6.160)	(2.868)
$\widehat{\delta}$	0.015	–	–	-0.901	-0.097
	(0.048)			(3.982)	(0.514)

Notes: For each of the ten inflation series, Table 6 reports QML estimates of the in-mean and level parameters for the ARFIMA-FIGARCH-in-mean-level model. The numbers in parentheses are absolute t-statistics. *A – indicates that there was no convergence. The numbers in [$\cdot$] indicate the lag of inflation in the variance equation.

In four out of the ten countries (Belgium, Germany, Portugal and Spain) there is no convergence when we employ the model with the simultaneous feedback. In all other six countries we find a positive association between lagged inflation and nominal uncertainty similar to that found with the two-step approach. However, another disadvantage of the simultaneous methodology is that in some cases the estimates of the conditional variances are negative.

In Finland, France, Italy and the UK the in-mean coefficient is insignificant, a result which is identical to that of the causality tests at lag 4. Similarly, in Sweden, as with the two-step approach, we find evidence for the Holland hypothesis. Moreover, in the four countries where there is no convergence, we estimate the model without the level effect ($\gamma_i = 0$) and the results (not reported) square with the findings of the two-step strategy at lag 4. That is, we do not find a significant effect of uncertainty on inflation. Hence, we generally find the two approaches to be in agreement. The only exception is the case of the Netherlands, where we estimate an insignificant δ, but find significant evidence for the Cukierman-Meltzer hypothesis at lag 4 in the two-step approach. However, it should be re-emphasized that such a result is plausible, since 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.

5.4.3 European Monetary System

Hyung and Franses (2004) point out that inflation rates may perhaps show long memory because of the presence of neglected occasional breaks in the series rather than being really I(d). Our sample period includes various exchange rate and monetary policy regimes. For example, the Bundesbank set a monetary target in 1975, after the break up of Bretton Woods. Originally, a fixed money target was announced but after two years this was changed to a fixed range. Like many other central banks, the Bundesbank translated its main policy goals (e.g., controlling inflation) into near term interest rate objectives. It in turn supplied bank reserves to meet these objectives. After 1985 the Bundesbank supplied banks with reserves mainly via repurchase agreements. Reunification of course introduced new complexities for monetary management. The British also introduced money targeting in the mid-1970s in response to mounting inflation concerns. Although inflation fell subsequent to the 1973 oil price shock, beginning in 1978 prices in the United Kingdom began to accelerate again, with inflation ultimately reaching nearly 20% by 1980. The perception of an inflationary crisis led to a change in strategy in 1979. A comparison with Germany does not portray British monetary policy in a favorable light. Not only has British inflation had higher mean and greater volatility, but the unemployment rate has also been high and variable. However, in the 1980s British inflation performance did improve considerably, remaining well below the 1970s level and becoming less variable.

Overall, the four decades under investigation are characterized by persistently high inflation, as was the case from the late 1960s through the early 1980s, followed by the relatively shock-free 1990s. Since the early 1980s, there has been a tremendous improvement in macroeconomic performance in European countries. This was the case for two reasons. First, the global reduction in inflationary pressures. Second, some countries 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. Furthermore, both inflation and output growth have become more stable. In what follows we examine whether the transition from the high inflation of the sixties and seventies to an era of low inflation during the 1980s and 1990s affects the dynamic interaction between nominal uncertainty and either inflation or output growth by examining the period that starts in 1980 and continues till to the end of the sample. The choice of this period is also based on the widely accepted notion that with the introduction of the exchange rate mechanism (ERM) in March 1979 monetary stability was achieved in Europe.

Table 7 presents QML estimates of d_m and d_v. For all countries the estimated long memory conditional mean parameter is in the range $0.134 \leq \widehat{d}_m \leq 0.379$. The value of the coefficient for the Netherlands (0.134) is markedly lower than the corresponding value for Spain (0.379). However, although the estimated value of d_m for the Netherlands is relatively small it is significantly different from zero. The estimation of a FIGARCH model (not reported) for the Netherlands realized an estimated value of d_v close to and not significantly different from zero. In other words, the conditional variance of this inflation series is characterized by a stable GARCH behavior. For the other nine countries, the values of d_v vary from 0.203 (Sweden) to 0.339 (Germany).

It is noteworthy that for the majority of the countries the estimates of d_m and d_v are similar to the ones for the entire period. Moreover, in Portugal and Spain (the two countries which were characterized by a near integrated GARCH behavior) the estimated values of d_v (d_m) are lower (higher) than the corresponding values for the whole sample. Thus, for these two inflation series there appears to be a trade off between the degree of persistence in the first two conditional moments. In sharp contrast, for Belgium, Italy and the UK which were characterized by the presence of quite strong long memory in the inflation uncertainty, the estimates of both d_m and d_v are lower than the ones for the 1962-2003 period.

Generally speaking, in the majority of the cases the estimated values of d_m and d_v are lower than the corresponding values for the entire sample. This result is in agreement with the conclusion of Caporale and Gil-Alana (2003). They investigate the stochastic behavior of the inflation series in three hyper-inflation countries. They test for fractional integration and find that when allowing for structural breaks the order of integration of the series decreases considerably. However, the parameter estimates still support the idea that dual long memory effects are present in the inflation process for nine out of the ten European countries. This result is consistent with the findings of a previous study by Bos et al. (1999). They find that the apparent long memory in monthly G7 inflation rates is quite resistant to mean shifts.

Table 7: ARFIMA-FIGARCH models 1980-2004.

	Belgium	Finland	France	Germany	Italy
$\widehat{d}_m$	0.146	0.136	0.190	0.209	0.289
	(2.102)	(2.973)	(2.333)	(2.571)	(4.718)
$\widehat{d}_v$	0.216	0.233	0.208	0.339	0.277
	(2.309)	(2.166)	(2.119)	(1.927)	(3.179)

	Netherlands	Portugal	Spain	Sweden	UK
$\widehat{d}_m$	0.134	0.218	0.379	0.160	0.275
	(2.407)	(1.983)	(5.257)	(2.122)	(2.339)
$\widehat{d}_v$	-	0.221	0.313	0.203	0.273
		(4.256)	(4.951)	(1.836)	(2.132)

Notes: For each of the ten inflation series, Table 7 reports QML estimates of the two long memory parameters for the ARFIMA-(FI)GARCH model. The numbers in parentheses are t-statistics.

Table 8 reports the results of causality between inflation, nominal uncertainty and real growth for the various ARFIMA-FIGARCH models for the post-1979 period.[17]

[17] Table 8 reports (in case of significance) only the sign of the sum of lagged coefficients

Panel A considers Granger causality from inflation to uncertainty about inflation. For this subperiod we find evidence that increased inflation raises its uncertainty in nine countries. For the Netherlands inflation has no impact on its uncertainty. Moreover, in this low-inflation period the evidence is mild(weak) for France(Germany) where it applies for two(one) of the chosen lags. Hence, the picture for the post-1979 period is similar to that of the entire period.

Panel B reports the results of the causality tests where causality runs from the nominal uncertainty to inflation. The findings for this subperiod provide support for the Cukierman-Meltzer hypothesis in some countries and for the Holland hypothesis in other countries. For three countries, uncertainty about inflation has a positive impact on inflation. Strong evidence in favour of the Cukierman-Meltzer hypothesis applies for Portugal and Spain. Relatively weak evidence applies for France (12 lags). Holland's hypothesis receives support in five countries, namely, in Germany, Italy, the Netherlands, the UK (optimal lag length) and Sweden (for all lags). None of the two theories is supported in Belgium and Finland where inflation is independent from changes in nominal uncertainty. The results are qualitatively similar to the analogous results from the entire sample. However, in Italy a negative effect begins to exist after 1979, whereas in France the evidence for the Cukierman-Meltzer hypothesis is weaker for the low-inflation period.

For completeness Panel C reports the results of causality from nominal uncertainty to output growth. For the post-1979 period, as for the entire sample period, we find evidence supporting the negative welfare effects of nominal uncertainty in Germany, Italy, Sweden and the UK. For both periods Dotsey-Sarte's hypothesis regarding the positive growth effects of uncertainty receives support in Finland, the Netherlands, Portugal and Spain. In the post-1979 period the evidence is weaker for the Netherlands and Portugal where it applies for only one of the chosen lags. In France the effect is positive during the entire period but turns to negative in the post-1979 period. In Belgium the effect is negative in the period 1965-2004 but it disappears in the low-inflation period.

Comparing the results of the post-1979 period with those of the entire period, we note that for the majority of the countries the three effects for the low-inflation period are very similar to those for the entire period. For those countries where we found changes in the effects either the impact of nominal uncertainty on inflation or output growth became less significant which is not surprising since the inflation series are less volatile in the low-inflation period or are more in line with Holland's stabilization hypothesis.

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for the optimal lag length. The figures of the MW statistics for the three different lags (4, 8 and 12) and the corresponding p-values are omitted for reasons of brevity.

Table 8: Granger-causality tests between inflation, output growth and inflation uncertainty (1980-2004).

	Be	Fi	Fr	Ge	It	Ne	Po	Sp	Sw	UK
Panel A:	+	+	+	+	+	x	+	+	+	+
Panel B:	x	x	+	−	−	−	+	+	−	−
Panel C:	x	+	−	−	−	+	+	+	−	−

Notes: The countries are as in Table 3. Panels A, B and C are as in Table 4. A +(−) indicates that the sign of the effect is positive (negative). An x indicates that the effect is insignificant.

6 Discussion

6.1 European monetary policy

The link between the inflation rate and its uncertainty acquires significant importance for the member countries of the Euro zone. The evidence that in all ten countries higher inflation causes greater uncertainty which then has negative output effects in five out of the ten countries strengthens the case for the choice of price stability as one of the objectives of monetary policy. Moreover, since the effects of nominal uncertainty on economic growth differ across the Euro zone, a common monetary policy that results in similar inflation rates across countries will have asymmetric real effects, provided these effects work via a change in nominal uncertainty. In other words, a reduction in inflation arising from a contractionary monetary policy applied by the ECB will increase growth in Belgium, Germany, Italy, Sweden, and the UK (where the Friedman hypothesis holds) but reduce it in Finland, France, the Netherlands, Portugal and Spain, where there is a positive effect of uncertainty. Therefore, the lack of uniform evidence supporting the second part of the Friedman hypothesis across the Euro zone countries has important policy implications as it makes a common monetary policy a less effective stabilization policy tool. It is noteworthy that evidence for the Dotsey-Sarte hypothesis obtains for the majority of the countries in the group which is characterized by a mild long memory in the conditional variance, and also for the two countries which exhibit near integrated GARCH behavior. In sharp contrast, evidence for the Friedman hypothesis applies in the three countries which are characterized by the presence of quite strong long memory GARCH behavior.

Moreover, less robust evidence is found regarding the direction of the impact of a change in nominal uncertainty on inflation. Countries like France, Portugal and Spain, for which we find evidence in favor of the Cukierman-Meltzer hypothesis, 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. When Grier and Perry (1998) looked for institutional reasons why the inflation response to increased uncertainty varies across countries, they noted that the countries associated with an opportunistic response have much lower central bank independence ratings than the countries associated with a stabilizing response. In this article we have used measures of central bank independence provided by Alesina and Summers (1993), which constructed a 1-4 (maximum independence) scale of central bank independence. Germany is rated as highly independent, with a score of 4. Netherlands also has a relatively independent central bank with a score of 2.5. In both countries increased inflation uncertainty lowers inflation as the sign at lag 12 (optimal lag length) is negative. Thus, one can argue that the most independent central banks are in countries where inflation falls in response to increased uncertainty. France has a relatively dependent central bank, with a score of 2. On the low side of the independence spectrum, Spain's rating is only 1. A lack of independence does seem to correspond to 'opportunistic behavior' because both countries show a highly significant positive effect of uncertainty on inflation. It is worth noting that evidence for the Cukierman-Meltzer hypothesis obtains for the two countries which exhibit near integrated GARCH behavior. In sharp contrast, evidence for the Holland hypothesis obtains for the majority of the countries which are characterized by the presence of mild long memory in nominal uncertainty. Finally, inflation is independent from changes in its uncertainty for two countries which are characterized by the presence of quite strong long memory in the conditional variance of the inflation rate.

6.2 Possible extensions

The main goal of this article has been to investigate the link between nominal uncertainty and macroeconomic performance, and to estimate the two main parameters driving the degree of their persistence, for ten European countries. In that respect the article has achieved its goal. As Hassler and Wolters (1995) point out, a likely explanation of the significant persistence in the inflation rate series is the aggregation argument, which states that persistence can arise from aggregation of constituent processes, each of which has short memory. Alternatively, Baum et al. (1999) conjecture that the long memory property of monetary aggregates will be transmitted to inflation, given the dependence of long-run inflation on the growth rate of money. However, one might also ask why it is necessary to allow for long memory in the conditional variance of inflation. To answer this we must enquire into the possible theoretical sources of heteroscedasticity in the inflation shocks. It will be very useful to provide a theoretical rationale for the dynamics of inflation. Here the choice of the FIGARCH model is justified solely on empirical grounds.

There is substantial evidence that European inflation rates have long memory, a feature which can be captured by a fractional integrated I(d) model. Hyung and Franses (2002) put forward a joint model which incorporates both long memory and occasional level shifts. Overall, however, they find that the dominant feature in 23 US inflation rates is long memory and that the level shifts are less important. This result suggests several avenues for further research. One promising avenue would be to adapt the ARFIMA-FIGARCH model in a way that incorporates occasional level shifts in both the conditional mean and the conditional variance.

Bos et al. (2002) have emphasized that the introduction of two macroeconomic leading indicators namely, the unemployment rate and the short term interest rate, in the ARFIMA model lower the estimate of the fractional parameter and thus account partly for the persistence in inflation. More importantly, they argue that the multi-step forecast intervals of the ARFIMAX model are more realistic than of the ARIMAX model.[18] In the context of our analysis, incorporating macroeconomic variables either in the ARFIMA or in the FIGARCH specification or in both could be at work. We look forward to sorting this out in future work.

Finally, Morana (2002) suggests that long memory in inflation is due to the output growth process. His model implies that inflation and output growth must share a common long memory component. Using a bivariate ARFIMA-FIGARCH model, which allows the measurement of uncertainty about inflation and output growth by the respective conditional variances, one can test for the empirical relevance of several theories that have been advanced on the relationship between inflation, output growth, real and nominal uncertainty.

7 Conclusions

In this article we have used ARFIMA-FIGARCH models to generate the conditional variance of inflation as a proxy of its uncertainty. We then performed Granger causality tests to examine the bidirectional relationship between the two variables. We provided overwhelming evidence that increased inflation raises nominal uncertainty, confirming the theoretical predictions made by Friedman. Uncertainty surrounding future inflation appeared to have a mixed impact on inflation. The division of countries by how their inflation rates respond to inflation uncertainty appears to be closely related to existing rankings of central bank independence. We also found that increased nominal uncertainty significantly affects output growth in the ten European countries but not all in the same manner. The lack of uniform evidence supporting the second leg of the Friedman hypothesis across the Euro zone countries has important implications as it makes a common monetary policy a less effective stabilization policy tool.

The results in this article highlight the importance of modeling long mem-

[18]ARFIMAX denotes an ARFIMA model with explanatory variables in the mean equation.

ory not only in the conditional mean of inflation but in its conditional variance as well. We find that in all the cases there is a need to consider the joint ARFIMA-FIGARCH model, as in no case does one of its nested versions yield a better fit. Overall, these findings suggest that much more attention needs to be paid to the consequences of dual long memory when estimates of nominal uncertainty are used in applied research. In other words, as our results indicate, estimates of uncertainty that ignore the effects of dual long memory may seriously underestimate both the degree of persistence of uncertainty and its consequences for the inflation-uncertainty hypothesis.

Possible extensions of this article could go in different directions. One could provide an enrichment of the dual long memory model by allowing lagged values of the conditional variance to affect the inflation. Finally, it is worth pointing to an important issue which we have not addressed. The dual long memory model used in this article ignores the possibility of structural instability caused by changing regimes. One could develop a dual long memory Markov switching model that explains the changing time series behavior of inflation in the post war era. This is undoubtedly a challenging yet worthwhile task.

Appendix

A Impulse response functions

Figures 1 - 3 plot the effects of a one-time one-standard-deviation increase in inflation on nominal uncertainty (left), in nominal uncertainty on inflation (middle) and in nominal uncertainty on output (right) for Germany, the Netherlands and the UK.

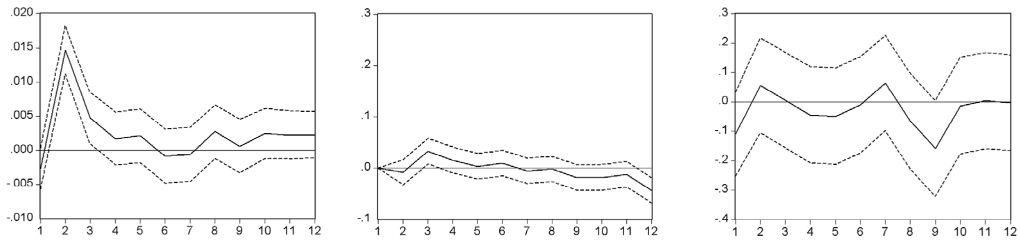

Figure 1: Germany.

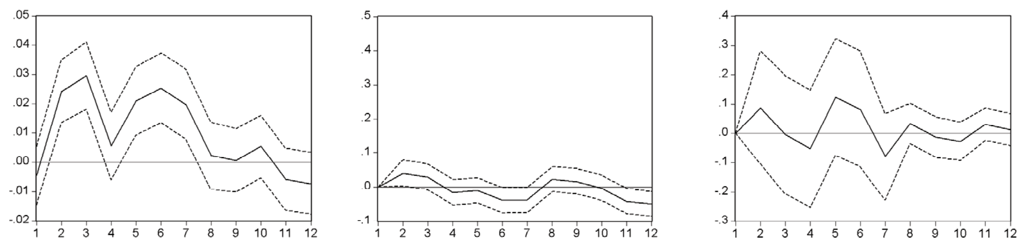

Figure 2: Netherlands.

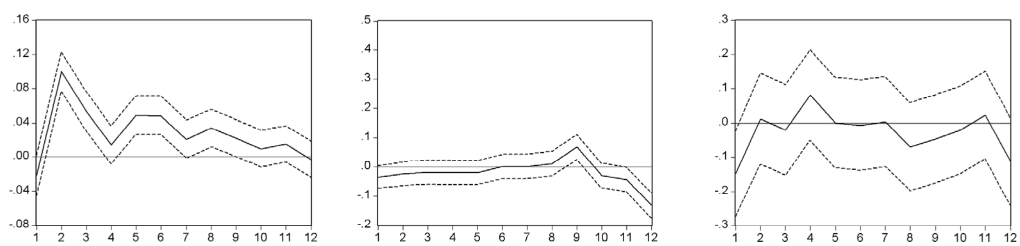

Figure 3: UK.

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

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