



On the inflation-uncertainty hypothesis in the USA, Japan and the UK: a dual long memory approach

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Abstract

We use parametric models of long memory in both the conditional mean and the conditional variance of inflation and monthly data in the USA, Japan and the UK for the period 1962–2001 to examine the relationship between inflation and inflation-uncertainty. In all countries, inflation significantly raises inflation-uncertainty as predicted by Friedman. Increased nominal uncertainty affects inflation in Japan and the UK but not in the same manner. The results from Japan support the Cukierman–Meltzer hypothesis. In the UK uncertainty surrounding the future inflation appears to have a mixed impact on inflation.

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1. Introduction

The issue of the welfare costs of inflation has been one of the most researched topics in macroeconomics both on the theoretical and empirical front. Friedman (1977) argues that a rise in the average rate of inflation leads to more uncertainty about the future rate of inflation. The opposite type of causation between inflation and its uncertainty has also been analyzed in the theoretical macroeconomics literature. Cukierman and Meltzer (1986) argue that central banks tend to create inflation surprises in the presence of more

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inflation-uncertainty. Clarida et al. (1999) emphasize the fact that since the late 1980s a stream of empirical work has presented evidence that monetary policy may have important effects on real activity. Consequently, there has been a great resurgence of interest in the issue of how to conduct monetary policy. If an increase in the rate of inflation causes an increase in inflation-uncertainty, one can conclude that greater uncertainty—which many have found to be negatively correlated to economic activity—is part of the costs of inflation. Thus, if we hope ever to give a really satisfactory answer to the questions:

- What actions should the central bankers take?
- What is the optimal strategy for the monetary authorities to follow?

we must first develop some clear view about the temporal ordering of inflation and nominal uncertainty.

In this paper, the above issues are analyzed empirically for the USA, Japan and the UK with the use of a parametric model of long memory in both the conditional mean and the conditional variance. Our emphasis on these three countries is justified by a number of considerations. First, the USA is the best-documented case and the American experience has played an important role in setting the agenda for previous analysis of monetary policy.¹ Second, the USA and Japan are the two largest economies in the world and changes in their inflation rates (variabilities) and average growth rates have repercussions in the world economy. Third, all three countries experienced wide variations in their conduct of monetary policy in the last forty years. For example, the increase in oil prices in late 1973 was a major shock for Japan, with substantial adverse effects on inflation, economic growth, and the government's budget. In response to an increase in the inflation rate to a level above 20% in 1974 the bank of Japan, like other central banks, began to pay more attention to money growth rates.² Prior to 1978 the Bank of Japan was committed only to monitoring rather than controlling money growth. However, after 1978 there did appear to be a substantive change in policy strategy, in the direction of being more 'money-focused'. Particularly striking was the different response of monetary policy to the second oil price shock in 1979. The difference in the inflation outcome in this episode was also striking, as inflation increased only moderately with no adverse effects on the unemployment situation. Beginning 1989 asset prices came down as money growth slowed, economic activity weakened and there was a slowdown in lending by Japanese banks. In responding to these developments the Bank of Japan permitted a considerable increase in the variability of broad money growth after the late 1980s (Bernanke and Mishkin, 1992). Finally, these three countries represent 'independent observations' in the sense that, no two of them

¹As emphasized by Bernanke and Mishkin (1992), the conduct of monetary policy in the USA is conventionally divided into three regimes. During the 1970s, the Fed did not consider meeting money growth targets to be of high priority, placing greater weight on reducing unemployment while maintaining a relatively smooth path for interest rates. However, the change in Fed operating procedures in 1979 was accompanied by a decision by the Fed to place greater weight on monetary targets and to tolerate high and volatile interest rates in order to bring down inflation. The main objectives during the latter part of the 1980s were exchange rate stabilization, financial market stability (particularly after the October 1987 stock market crash) and the maintenance of low and stable inflation.

²See Bernanke and Mishkin (1992) for an excellent discussion of the monetary policy in Japan.

belonged to a common system of fixed exchange rates.³ Other countries with independent monetary policies, such as Australia, would be interesting to study but are excluded because of space.

The development of GARCH techniques allows the measurement of inflation uncertainty by the conditional variance of the inflation series and the more accurate testing of the Friedman and the Cukierman and Meltzer hypotheses. Several researchers have examined the inflation-uncertainty relationship using GARCH measures of inflation-uncertainty. Many studies on the relationship between inflation and its uncertainty used GARCH type models with a joint feedback between the conditional mean and variance of inflation (i.e., Brunner and Hess, 1993; Fountas et al., 2003). In contrast, Grier and Perry (1998), Fountas and Karanasos (2004) use the estimated conditional variance from GARCH type models and employ Granger methods to test for the direction of causality between average inflation and inflation-uncertainty for the G7.⁴ All the preceding works use traditional ARMA processes to model the conditional mean of inflation. On the other hand, Brunner and Hess (1993) argue that the US inflation rate was stationary before the 1960s, but that it has possessed a unit root since this time. Subsequently, Hassler and Wolters (1995) have found that the inflation rates of five industrial countries were well explained by different orders of integration, which varied around the stationarity border of 0.5. Baum et al. (1999) also found evidence that both CPI- and WPI-based inflation rates for many industrial as well as developing countries are fractionally integrated with a differencing parameter that is significantly different from zero or unity. Along these lines, Baillie et al. (1996b), Hwang (2001) estimate various ARFIMA–GARCH-in mean models where lagged inflation is included in the variance equation.

In a recent paper, Baillie et al. (2002) found that inflation has the rather curious property of persistence in both its first and its second conditional moments. They introduce the ARFIMA–fractionally integrated GARCH (ARFIMA–FIGARCH) model, which is sufficiently flexible to handle the dual long-memory behavior encountered in inflation. To this end, this study uses an ARFIMA–FIGARCH type model to generate a time-varying conditional variance of surprise inflation. This model has a distinct advantage for this application: it nests several alternative GARCH models of conditional heteroscedasticity. With this conditional variance as a measure of inflation-uncertainty, it then employs Granger methods to test the direction of causality between average inflation and uncertainty. The Granger-causality approach is adopted in this paper instead of the simultaneous-estimation approach because it allows us to capture the lagged effects between the variables

³ However, as Bernanke and Mishkin (1992) point out, there are some parallels between the recent histories of British and American monetary policies. As in the USA, the British introduced money targeting in the mid-1970s in response to mounting inflation concerns. During the pre-1979 period, the British monetary authorities, like their American counterparts, were not taking their money growth targets very seriously. As in the United States, the perception of an inflationary crisis led to a change in strategy in 1979. Overall, a comparison with the US and the other countries does not put British monetary policy in a favorable light. However, in the 1980s British inflation performance did improve considerably, remaining well below the 1970s level and becoming significantly less variable (Bernanke and Mishkin, 1992).

⁴ We should also mention that several empirical studies (i.e., Grier and Perry, 2000; Fountas et al., 2002) use bivariate GARCH type models to estimate simultaneously the conditional means, variances and covariances of inflation and output growth. These models make it possible to test for evidence on all the bidirectional causality relationships between inflation, output growth, and uncertainty about inflation and output growth.

of interest. In addition, the former approach minimizes the number of estimated parameters, whereas the latter approach is subject to criticism on the grounds of the potential negativity of the variance.

This article is organized as follows: [Section 2](#) considers the hypotheses about the causality between inflation and inflation-uncertainty in more detail and provides the model. [Section 3](#) discusses the data and the results and [Section 4](#) summarizes the main conclusions.

2. Theory and model

2.1. Theory

[Friedman \(1977\)](#) outlined an informal argument regarding the real effects of inflation. Friedman's point comes in two parts. In the first part, an increase in inflation may induce an erratic policy response by the monetary authority and, therefore, lead to more uncertainty about the future rate of inflation. The second part of Friedman's hypothesis predicts that inflation uncertainty causes an adverse output effect.⁵ [Ball \(1992\)](#), using an asymmetric information game, offers a formal derivation of Friedman's hypothesis that higher inflation causes more inflation-uncertainty. It is also possible that more inflation will lead to a lower level of inflation-uncertainty. The argument advanced by [Pourgerami and Maskus \(1987\)](#) is that in the presence of rising inflation agents may invest more resources in forecasting inflation, thus reducing uncertainty about inflation. A formal analysis of this effect is presented in [Ungar and Zilberfarb \(1993\)](#).

The causal effect of inflation-uncertainty on inflation has been analyzed in the theoretical macro literature by [Cukierman and Meltzer \(1986\)](#). Using the well-known Barro–Gordon model, [Cukierman and Meltzer \(1986\)](#) show that an increase in uncertainty about money growth and inflation will raise the optimal average inflation rate because it provides an incentive to the policymaker to create an inflation surprise in order to stimulate output growth. Therefore, the prediction of the Cukierman–Meltzer analysis is that higher inflation-uncertainty leads to more inflation. [Holland \(1995\)](#) has supplied a different argument based on the stabilization motive of the monetary authority, the so-called “stabilizing Fed hypothesis”. He claims that as inflation-uncertainty rises due to increasing inflation, the monetary authority responds by contracting money supply growth, in order to eliminate inflation-uncertainty and the associated negative welfare effects. Hence, Holland's argument supports a negative causal effect of inflation-uncertainty on inflation.

2.2. Model

In this section we present the ARFIMA–FIGARCH model, which generates the long-memory property in both the first and second conditional moments, and is thus sufficiently flexible to handle the dual long-memory behavior encountered in inflation.

⁵ Investigating the causal impact of nominal uncertainty on output growth is an interesting avenue for future research, but is beyond the scope of the present paper.

In the ARFIMA($p, d_m, 0$)–FIGARCH(1, $d_v, 1$) model the mean equation is defined as:

$$(1 - \phi_1 L - \phi_{12} L^{12} - \phi_{24} L^{24})(1 - L)^{d_m}(\pi_t - \mu) = \varepsilon_t, \tag{1}$$

where π_t denotes the inflation rate and $0 \leq d_m \leq 1$; ⁶ 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 $t - 1$. The structure of the conditional variance is:

$$(1 - \beta L)h_t = \omega + [(1 - \beta L) - (1 - \varphi L)(1 - L)^{d_v}] \varepsilon_t^2, \tag{2}$$

where $0 \leq d_v \leq 1$, $\omega > 0$, and $\varphi, \beta < 1$.

As Baillie et al. (2002) note, if $h_t = \omega$, a constant, the process reduces to the ARFIMA($p, d_m, 0$) model and π_t 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 $d_v \neq 0$.

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 uncertainty about inflation. 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 inflation-uncertainty. First, it recognizes the long-memory aspect of the inflation rate and provides an empirical measure of inflation-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.

3. Empirical analysis

We use monthly data on the consumer price index (CPI) obtained from the *OECD Statistical Compendium* as proxies for the price level.⁷ The data range from January

⁶ The fractional differencing operator, $(1 - L)^d$ is most conveniently expressed in terms of the hypergeometric function:

$$(1 - L)^d \equiv F(-d, 1; 1; L) = \sum_{j=0}^{\infty} \frac{\Gamma(j - d)}{\Gamma(-d)\Gamma(j + 1)} L^j = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j,$$

where

$$F(a, b; c; z) \equiv \sum_{j=0}^{\infty} \frac{(a)_j (b)_j z^j}{(c)_j j!}$$

is the Gaussian hypergeometric series, $(b)_j$ the shifted factorial defined as $(b)_j \equiv \prod_{i=0}^{j-1} (b + i)$ (with $(b)_0 \equiv 1$), and $\Gamma(\cdot)$ the gamma function.

⁷ Since most of the studies use CPI-based inflation measures (i.e., Caporale and McKierman, 1997; Grier and Perry, 1998; Baillie et al., 2002) we construct our inflation and inflation-uncertainty measures from the consumer price index. Alternatively, one can use either the producer price index (PPI) or the GNP deflator. Brunner and Hess (1993) use all three measures of inflation but they discuss only the results using CPI inflation. Grier and Perry (2000) use both (CPI and PPI) indices and find that the CPI and PPI results are virtually identical.

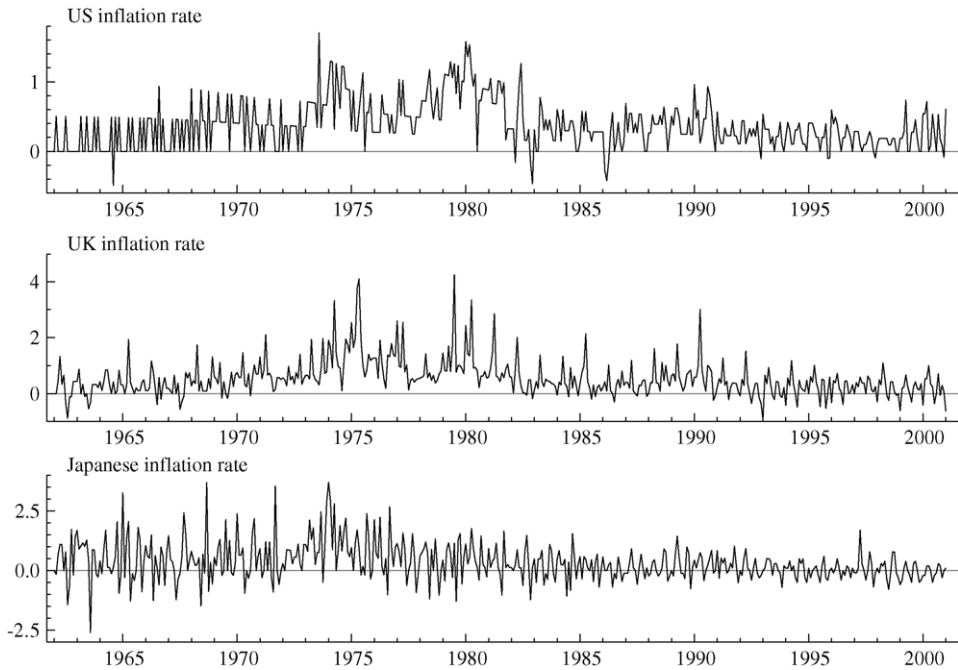


Fig. 1. Monthly inflation rates for the period January 1962–December 2000.

1962–December 2000 and cover three industrial countries, namely, the USA, the UK and Japan. Inflation is measured by the monthly difference of the log CPI [$\pi_t = 100 \log(\text{CPI}_t / \text{CPI}_{t-1})$]. Allowing for differencing leaves 469 usable observations. The inflation rates of the USA, UK and Japan are plotted in Fig. 1.

Table 1 presents summary statistics for the three inflation rates. The results indicate that the distributions of the inflation series are skewed to the right. The distribution of the British inflation rate has fat tails. 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

Table 1
Summary statistics for inflation rates

	μ	σ	K	S	JB	Q_{12}^2	LM
USA	0.376	0.341	0.77 (0.00)	0.81 (0.00)	63.47 (0.00)	1123.81 (0.00)	130.88 (0.00)
UK	0.550	0.656	6.46 (0.00)	1.90 (0.00)	1098.2 (0.00)	250.09 (0.00)	30.53 (0.00)
Japan	0.348	0.794	2.47 (0.00)	0.93 (0.00)	186.94 (0.00)	129.27 (0.00)	9.02 (0.00)

Notes: μ denotes the average inflation rate over the period January 1962–December 2000, and σ its standard deviation. K and S are the estimated kurtosis and skewness, respectively. JB is the Jarque–Bera statistic for normality. Q_{12}^2 is the 12th order Ljung–Box test for serial correlation in the squared deviations of the inflation rate from its sample mean. The Engle test for ARCH effects is denoted by LM. Numbers in parentheses are P -values.

Table 2

Tests for the order of integration of different countries' inflation series

	H ₀ : I(1)		H ₀ : I(0)	
	Z(t _μ)	Z(t _τ)	$\hat{\eta}_\mu$	$\hat{\eta}_\tau$
USA	7.11***	8.49***	0.84***	0.68***
UK	5.80***	8.33***	1.03***	0.64***
Japan	3.96***	10.47***	2.31***	0.29***

Notes: Z(t_μ) and Z(t_τ) are the Phillips–Perron 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_μ) and Z(t_τ) 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.

inflation rate from its sample mean indicate the existence of ARCH effects. This evidence is also supported by the LM statistics, which are highly significant.

The autocorrelations for CPI inflation for the three countries (not reported) exhibit the clear pattern of slow decay and persistence. The autocorrelations of the first differenced inflation series (not reported) appear to be overdifferenced with large negative autocorrelations at lag one.

Phillips and Perron (1988) (henceforth PP) and Kwiatkowski et al. (1992) (henceforth KPSS) develop two alternative approaches to testing for unit roots. Table 2 presents the results of applying the PP and KPSS tests to the three inflation series. In all cases, we reject both the KPSS and PP statistics. Hence, for all three countries there is evidence that inflation may not be generated by either an I(0) or I(1) process and this is at least indicative of fractional integration (see also Baillie et al., 1996b).

3.1. Estimated models of inflation

We proceed with the estimation of the ARFIMA(*p*, *d*_m, 0)–FIGARCH(1, *d*_v, 1) model in equations (1) and (2) in order to take into account the serial correlation observed in the levels and squares of our time series data, and to capture the possible long memory in the conditional mean and the conditional variance. We estimate the ARFIMA–FIGARCH models using the quasi maximum likelihood estimation (QMLE) method as implemented by Laurent and Peters (2002) in Ox.⁸ Table 3 reports the results for the period 1962–2001.

The seasonal autoregressive parameters ($\phi_1, \phi_{12}, \phi_{24}$) were necessary to account for the significant seasonality, which is evident for all three countries.⁹ The ϕ_1 parameter was only significant for the USA. The estimated ARCH parameters ($\hat{\alpha}, \hat{\beta}$) for the US inflation are significant and satisfy the set of conditions sufficient to guarantee the non-negativity of the

⁸The 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.

⁹Alternatively, we also estimated a moving average specification with parameters ($\theta_1, \theta_{12}, \theta_{24}$), but the AIC and SIC information criteria came out in favor of the autoregressive specification.

Table 3
ARFIMA–FIGARCH models

	$\hat{\mu}$	\hat{d}_m	$\hat{\phi}_{12}$	$\hat{\phi}_{24}$	$\hat{\omega}$	\hat{d}_v	$\hat{\phi}$	$\hat{\beta}$	Q_{12}	Q_{12}^2
USA	0.296 (2.018)	0.342 (4.787)	0.161 (3.491)	0.139 (2.719)	0.002 (1.538)	0.692 (1.892)	0.325 (1.747)	0.768 (6.863)	22.21 [0.04]	3.40 [0.99]
UK	0.100 (0.208)	0.367 (4.479)	0.407 (6.911)	0.331 (5.828)	0.022 (1.689)	0.474 (4.025)	–	–	14.01 [0.30]	12.27 [0.42]
Japan	0.098 (1.084)	0.043 (1.133)	0.361 (6.313)	0.263 (5.033)	0.026 (1.089)	0.288 (5.194)	–	–	20.05 [0.07]	12.82 [0.38]

Notes: For each of the three inflation series, Table 3 reports QMLE 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 [·] are *P*-values. For the USA we estimate a $\hat{\phi}_1$ of -0.148 (-2.024).

Table 4

Tests of fractional differencing parameters in the first and second conditional moments (USA)

	$H_0 : \text{ARMA}$ ($d_m=0$)		$H_0 : \text{GARCH}$ ($d_v=0$)		$H_0 : \text{ARMA-GARCH}$ ($d_m, d_v=0$)	
	\hat{d}_m	LR	\hat{d}_v	LR	LR	W
USA	0.342{0.071}	35.45 [0.00]	0.692 {0.365}	4.14 [0.04]	36.56 [0.00]	15.97 [0.00]

Notes: Columns 2, 4 and 5 report the value of the following likelihood ratio test: $\text{LR} = 2[\text{ML}_u - \text{ML}_r]$, where ML_u and ML_r denote the maximum log-likelihood values of the unrestricted and restricted models respectively. The last column reports the Wald statistic. The numbers in [·] are P -values. The numbers in {·} are standard errors.

conditional variance (see Baillie et al., 1996a). For the UK and Japan, the Akaike and Schwarz information criteria (hereafter, AIC and SIC respectively) come out in favor of the FIGARCH($0, d_m, 0$) model. The estimated long-memory conditional mean parameter is in the range $0 < \hat{d}_m < 0.37$. The estimated value of d_m for Japan in Table 3 is 0.043, which is significantly different from zero at the 0.26% level and implies some very mild long-memory features.¹⁰ In all countries the estimates for the fractional differencing parameter (\hat{d}_v) are relatively large and statistically significant.¹¹ Finally, with all three countries, the hypothesis of uncorrelated standardized and squared standardized residuals is well supported by the Ljung–Box test statistics, indicating that there is no statistically significant evidence of misspecification.

To test for the persistence in the first two conditional moments of the three inflation series, we examine the likelihood ratio (LR) tests for the linear constraints $d_m = 0$ ('ARMA' model) and $d_v = 0$ ('GARCH' model). We also test the joint hypothesis that $d_m = d_v = 0$ using both an LR test and a Wald (W) statistic. As seen in Table 4 for the USA the LR and W statistics clearly reject the 'ARMA' and/or the 'GARCH' null hypotheses against the ARFIMA–FIGARCH model. Similar results are obtained for the UK and Japan but are omitted for space considerations. The evidence obtained from the Wald and LR tests is reinforced by the model ranking provided by the AIC and SIC model selection criteria. In almost all cases the criteria (not reported) favor the ARFIMA–FIGARCH model over both the ARMA–FIGARCH and ARFIMA–GARCH models.¹² Hence, from the various diagnostic statistics it appears that monthly CPI inflation has long-memory behavior in both its first and its second conditional moments.

3.1.1. Predictability of higher levels of inflation

In the USA inflation accelerated two years after the augmentation of defense spending in connection with the Vietnam war which took place in the mid-1965. Moreover, the increase

¹⁰ Although d_m is insignificant for Japan it seems to improve the performance of the model since restricting d_m to being zero results in Ljung–Box test statistics which indicate serial correlation in the standardized residuals. Moreover, note that in Baillie et al. (1996b) the estimate of the d_m parameter for Japan is also insignificant.

¹¹ The estimates for d_m and d_v in Baillie et al. (2002) are quite close to the ones we obtain. In particular, for the USA and the UK they estimated a $d_m(d_v)$ of 0.414 (0.644) and 0.364 (0.633), respectively, while for Japan they estimate a d_v of 0.317.

¹² For Japan the LR test, the W statistic and the selection criteria favor the ARMA–FIGARCH model over the ARFIMA–FIGARCH model.

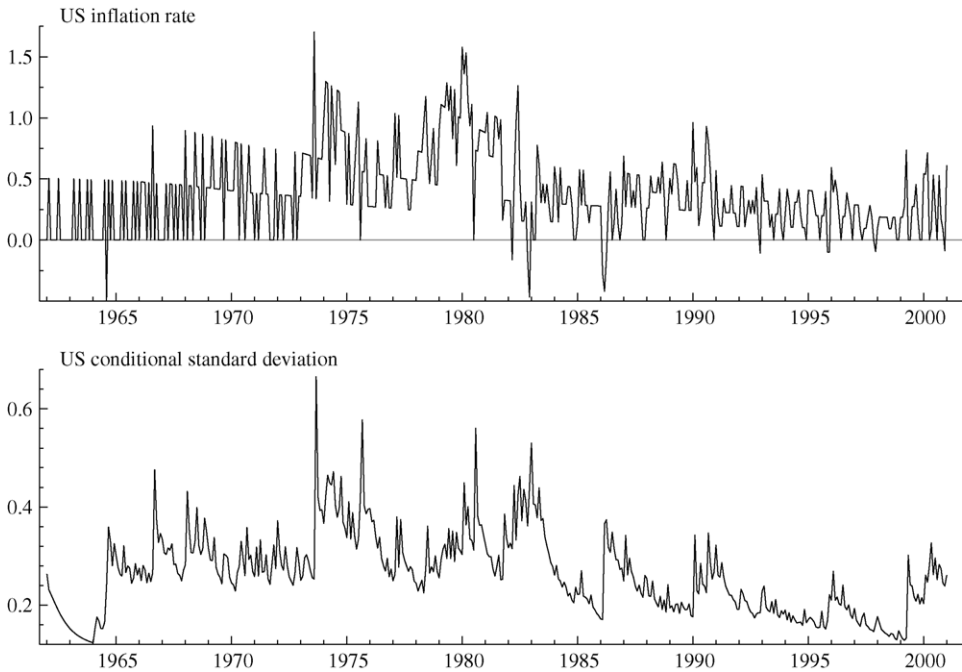


Fig. 2. US inflation rate and conditional standard deviation for the period January 1962 to December 2000.

in the oil prices by OPEC in the fourth quarter of 1973 and the progressive elimination of control on prices and wages amplify the acceleration of American inflation in 1974. Finally, the considerable fluctuation of oil prices during the period 1979–1980 led the federal reserve to implement a new restrictive monetary policy. Since the early 1980s the American economy embarked on a productivity growth phase supported by a decrease of oil prices and the reduction of the inflation rate. The USA inflation and inflation-uncertainty series are shown in Fig. 2, which plots the inflation rate and its corresponding conditional standard deviation from the ARFIMA–FIGARCH model.

Some researchers, such as [Cosimano and Jansen \(1988\)](#), have failed to find strong evidence that higher rates of inflation are less predictable. Using the dual long-memory specification, we compare our results with theirs. In contrast to the conclusion of these studies, Fig. 2 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 3.1%. In the high-inflation 1970s, the conditional standard deviation average (annual rate) is about 3.9%. Finally, in the low-inflation environment of the 1990s, the average of the conditional standard deviation is only 2.4%. Similar figures for the UK and Japan are omitted for reasons of brevity but are available from the authors on request.

3.1.2. Persistence in volatility

In order to illustrate how a shock to the conditional variance decays over time in the FIGARCH(1, d , 1) model we plot in Fig. 3 the cumulative impulse response function for

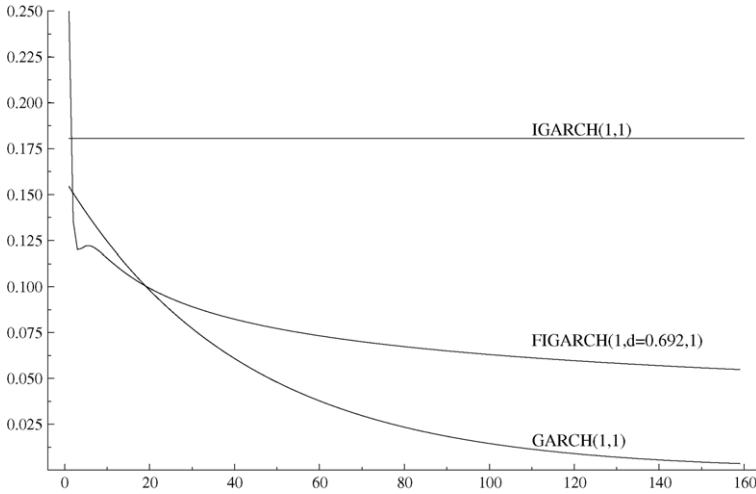


Fig. 3. Cumulative impulse response functions for the conditional variance of the US inflation rate.

the USA. Following Baillie et al. (1996a) the cumulative impulse response weights λ_k for the optimal forecast of the future conditional variance of the FIGARCH(1, d , 1) are given by:

$$\lambda_k = \sum_{j=0}^k \gamma_j, \tag{3}$$

where $\sum_{j=0}^{\infty} \gamma_j L^j \equiv (1 - \beta L)(1 - \varphi L)^{-1}(1 - L)^{1-d}$. The cumulative impulse response function of the FIGARCH model is compared with the one of the stable GARCH and integrated GARCH (IGARCH).¹³ For the stable GARCH we have $\lambda_k = (\varphi - \beta)\varphi^{k-1}$ for all $k \geq 1$ and for the IGARCH we have $\lambda_k = 1 - \beta$ for all $k \geq 1$. In the stable GARCH case a shock decays at a fast exponential rate, whereas in the IGARCH case it persists forever. In sharp contrast, the shock decays at a slow hyperbolic rate in the FIGARCH case.¹⁴

3.2. 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 average inflation and nominal uncertainty. Following Granger (1969) the following bivariate autoregression is used to test for causality between the inflation rate and its uncertainty:

$$\begin{bmatrix} \pi_t \\ h_t \end{bmatrix} = \begin{bmatrix} \alpha_\pi \\ \alpha_h \end{bmatrix} + \sum_{i=1}^k \begin{bmatrix} c_{\pi\pi,i} & c_{\pi h,i} \\ c_{h\pi,i} & c_{hh,i} \end{bmatrix} \begin{bmatrix} \pi_{t-i} \\ h_{t-i} \end{bmatrix} + \begin{bmatrix} e_{\pi t} \\ e_{h t} \end{bmatrix}, \tag{4}$$

¹³ For the stable GARCH, we estimated $\hat{\beta} = 0.822$ and $\hat{\varphi} = 0.976$. The $\hat{\beta}$ coefficient in the IGARCH was 0.819.

¹⁴ Similar plots are available for the other two countries but are omitted for reasons of brevity.

where $e_t = [e_{\pi_t}, e_{h_t}]'$ is a bivariate white noise with mean zero and non-singular covariance matrix Σ_e . The test of whether π_t (h_t) strictly Granger causes h_t (π_t) is simply a test of the joint restriction that all the $c_{h\pi,i}$ ($c_{\pi h,i}$), $i = 1, \dots, k$, are zero. In each case, the null hypothesis of no Granger-causality is rejected if the exclusion restriction is rejected. Bidirectional feedback exists if all the elements $c_{\pi h,i}$, $c_{h\pi,i}$, $i = 1, \dots, k$, are jointly significantly different from zero. However, if the variables are non-stationary, Park and Phillips (1989), Sims et al. (1990) have shown that the conventional asymptotic theory is not applicable to hypothesis testing in levels VAR's. In addition, Tsay and Chung (2000) in their analysis of spurious regression with fractionally integrated processes find that 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.

Therefore, we utilize the methodology developed by Toda and Yamamoto (1995) to test for causality between the inflation rate and its uncertainty, which leads to a χ^2 distributed test statistic despite any possible non-stationarity or cointegration between the two series.¹⁵ In other words, the advantage of this procedure is that it does not require a knowledge of cointegrated properties of the system (see Zapata and Rambaldi, 1997). The 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 information criteria. 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 Wald (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. For the USA, Japan and the UK both the AIC and SIC information criteria came out in favor of a VAR with 8, 4 and 12 lags, respectively. Since all variables are fractionally integrated with $d_m, d_v < 1$ we set $d_{\max} = 1$ and estimate VAR models with $k^* = k + 1$ lags. To ensure that our results are not sensitive to the choice of the lag length we report in Table 5 for all three countries the MW tests using four, eight and 12 lags, as well as the sums of lagged coefficients.

Panel A reports evidence on the Friedman hypothesis. Statistically significant effects are present for all countries. Panel B reports the results of the causality tests where causality runs from the nominal uncertainty to the rate of inflation. This Panel provides strong evidence in favor of the Cukierman and Meltzer hypothesis for Japan. For the USA, the effect of inflation-uncertainty on average inflation is positive but insignificant at any lag length. In contrast, we obtain mixed evidence for the UK. In particular, at eight lags uncertainty about inflation has a positive impact on inflation as predicted by Cukierman and Meltzer, whereas the value of the MW test statistic and the sign of the sum of lagged coefficients at 12 lags (optimal lag length) provide support for the Holland hypothesis.

When Grier and Perry (1998) look for institutional reasons why the inflation response to increased uncertainty varies across countries, they note that countries associated with an opportunistic response have much lower central bank independence than the countries associated with a stabilizing response. In this paper we use measures of central bank independence provided by Alesina and Summers (1993), which constructed a 1–4 (maximum independence) scale of central bank independence. The USA and Japan have relatively independent central banks with a score of 3. However, in Japan increased

¹⁵We are grateful to an anonymous referee for calling this paper to our attention.

Table 5
Granger-causality tests between inflation and nominal uncertainty

	USA	UK	Japan
(Panel A) H_0 : inflation does not Granger-cause inflation-uncertainty			
4 (5)	15.39*** (+)	14.29*** (+)	13.47*** (+)
8 (9)	19.39** (+)	35.35*** (+)	16.08** (+)
12 (13)	27.36*** (+)	36.82*** (+)	28.02*** (+)
(Panel B) H_0 : inflation-uncertainty does not Granger-cause inflation			
4 (5)	2.58 (+)	3.04 (–)	19.58*** (+)
8 (9)	7.16 (+)	31.41*** (+)	22.25*** (+)
12 (13)	11.68 (+)	63.21*** (–)	40.79*** (+)

Notes: The figures are MW statistics. The numbers in the first column give the lag structure and in parentheses the order of the VAR. The bold numbers indicate the optimal lag length chosen by AIC and SIC. A + (–) indicates that the sum of the lagged coefficients is positive (negative). ***, ** and * denote significance at the 0.01, 0.05 and 0.10 levels, respectively.

inflation-uncertainty raises inflation, while in the USA uncertainty does not Granger-cause inflation. Thus, one cannot argue that the most independent central banks are in countries where inflation falls in response to increased uncertainty. The UK has a relatively dependent central bank, with a score of 2. However, in the UK the sign (and significance) of the effect varies with the lag length. Thus, a lack of independence does not correspond to ‘opportunistic behavior’.

3.3. Comparison with other work

The GARCH time series studies that examine the link between inflation rates and inflation-uncertainty use various sample periods, frequency data sets and empirical methodologies.¹⁶ Some GARCH studies of this issue utilize the simultaneous-estimation approach. In particular, Caporale and McKierman (1997), Fountas (2001), estimate GARCH-type processes where lagged inflation is included in the conditional variance equation to test Friedman’s hypothesis. Brunner and Hess (1993) model the conditional variance as a non-linear function of lagged values of inflation. Baillie et al. (1996b), Hwang (2001), Fountas et al. (2003), employ univariate GARCH models that allow for simultaneous feedback between the conditional mean and variance of inflation while Grier and Perry (2000) use a bivariate GARCH in mean specification.¹⁷ Some researchers employ the Granger-causality approach. For example, Grier and Perry (1998), Fountas and Karanasos (2004) estimate univariate component GARCH models, Fountas et al. (2004) employ an EGARCH specification, while Fountas et al. (2002) use a bivariate constant correlation GARCH formulation.

More specifically, Baillie et al. (1996b) show that for the low-inflation countries (except the UK) there is no link between the inflation rate and its uncertainty, whereas for the high-

¹⁶ Fountas (2001) employs annual data on UK CPI for the period 1885–1998 while Fountas et al. (2004) use quarterly data from six European countries for the period 1960–1999.

¹⁷ These studies either use the conditional variance or the conditional standard deviation as a regressor in the conditional mean.

inflation countries strong evidence is provided of a bidirectional feedback between nominal uncertainty and inflation. Grier and Perry (1998) find that in all G7 countries inflation significantly raises inflation-uncertainty. Fountas et al. (2004), in five out of six European countries, and Fountas and Karanasos (2004) in six of the G7 countries also find support for Friedman's hypothesis. In sharp contrast, for Germany the study by Fountas and Karanasos finds that the effect of inflation on nominal uncertainty is negative as predicted by Pourgerami and Maskus (1987). These three studies find less robust evidence regarding the direction of the impact of a change in inflation-uncertainty on inflation. That is, they find evidence in favor of the Cukierman–Meltzer hypothesis for some countries and in favor of the Holland hypothesis for other countries.

In all three countries, our results on the Friedman hypothesis are identical to those of the above-mentioned studies. That is, in all studies which use the two-step approach and in most of the studies which use the simultaneous approach increased inflation affects nominal uncertainty positively. In sharp contrast, Baillie et al. (1996b) find, for the USA and Japan, that inflation-uncertainty is independent of changes in inflation. Hwang (2001) finds that the US inflation affected its uncertainty weakly and negatively. For the USA, our result that there is no causal effect of nominal uncertainty on inflation squares with the findings of most of the recent studies (i.e., Baillie et al., 1996b; Grier and Perry, 2000; Hwang, 2001; Fountas and Karanasos, 2004). However, Grier and Perry (1998) find that uncertainty about future inflation has a negative impact on inflation, whereas Fountas et al. (2003) find evidence for a positive effect of nominal uncertainty on inflation. Note that both studies estimate short-memory GARCH models. For Japan, we find that uncertainty about inflation has a positive effect on inflation, as predicted by Cukierman and Meltzer. This result is in agreement with the conclusion of Grier and Perry (1998), Fountas and Karanasos (2004). In sharp contrast, Fountas et al. (2002) provide strong empirical support for Holland's hypothesis. Our paper differs from Fountas et al. (2002) in the chosen econometric methodology (univariate dual long-memory GARCH-type models) and the use of CPI in measuring inflation. The authors estimate short-memory bivariate GARCH models and use PPI data. Moreover, Baillie et al. (1996b) fail to find any effect of nominal uncertainty on inflation for Japan. Our work differs from theirs in that we use more than one lag of monthly inflation and uncertainty to look for a link between the two. Finally, for the UK our result that uncertainty about future inflation appears to have a mixed impact on inflation is consistent with the findings of previous studies by Fountas et al. (2004), Fountas and Karanasos (2004). Note also that our result on the mild evidence (at lag 12) that increased nominal uncertainty lowers inflation is identical to that of Grier and Perry (1998).

4. Conclusion

We have used monthly data on inflation in the USA, Japan, and the UK to examine the possible relationship between inflation and nominal uncertainty, and hence test a number of economic theories. The results in this paper highlight the importance of modeling long memory not only in the conditional mean of inflation but in the conditional variance as well. The application of the ARFIMA–FIGARCH approach allows us to derive two

important conclusions: first, the Friedman hypothesis that inflation leads to more inflation-uncertainty applies in all countries. Since an increase in the rate of inflation causes an increase in inflation-uncertainty, we conclude that greater uncertainty—which many have found to be negatively correlated to economic activity—is part of the costs of inflation. This result may have important implications for the inflation–output relationship. Gylfason and Herbertsson (2001) argue that inflation can be detrimental to economic growth through four different channels. It would be interesting to find whether this negative effect may work also indirectly via the inflation uncertainty channel. For example, since the Japanese economy during the 1990s was plagued by a deflationary episode associated with low or zero rates of inflation and low-output growth rates it would be interesting to find out whether the low-output growth rates can be associated with the rate of inflation and the corresponding inflation-uncertainty as predicted by Friedman (1977). However, as emphasized by Gylfason and Herbertsson (2001), one can not preclude the possibility that low-inflation may be harmless to growth, perhaps even beneficial. Krugman (1998) has recommended more rapid monetary expansion and inflation in Japan in order to reduce real interest rates below 0 and thereby stimulate investment and growth.

Second, less robust evidence is found regarding the direction of the impact of a change in nominal uncertainty on inflation. No effect was present for the USA, whereas we obtained mixed evidence for the UK. At 12 lags, we find evidence against the Cukierman–Meltzer hypothesis. This evidence partially favors the ‘stabilization hypothesis’ put forward by Holland (1995). He claims that for countries where inflation leads to nominal uncertainty and real costs, we would expect the policy maker to stabilize inflation, hence a negative effect of inflation-uncertainty on inflation. This result squares with the findings of recent studies by Fountas and Karanasos (2004), Fountas et al. (2004). Both studies found that uncertainty about inflation causes negative real effects in the UK. In Japan we found strong evidence in favor of the Cukierman–Meltzer hypothesis. According to Devereux (1989) inflation-uncertainty can have a positive impact on inflation via the real uncertainty channel. If the variability of real shocks is the predominant cause of nominal uncertainty, then inflation-uncertainty and inflation are positively correlated. As real shocks become more variable the optimal degree of indexation declines. The inflation rate rises only after the degree of indexation falls. Assuming that changes in the degree of indexation take time to occur, greater inflation-uncertainty precedes higher inflation. We have also investigated whether one can find a correlation between central bank independence and inflation policy. Our conclusion is that there is no correlation.

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