The Effect of Inflation on Inflation Uncertainty in the G7 Countries: A Double Threshold GARCH Model

Kushal Banik Chowdhury[®] **and Nityananda Sarkar**

Indian Statistical Institute and Indian Statistical Institute

ABSTRACT

This paper studies the impact of inflation on inflation uncertainty in a modelling framework where both the conditional mean and conditional variance of inflation are regime specific, and the GARCH model for inflation uncertainty is extended by including a lagged inflation term in each regime. Applying this model to the G7 countries with monthly data from 1970 till 2013, it is found that the impact of inflation on inflation uncertainty differs over the regimes in most of the G7 countries. The findings also provide strong empirical support to the well-known Friedman-Ball hypothesis of positive impact of inflation on inflation uncertainty, but only for the high-inflation regime.

Key words: *Inflation; Inflation Uncertainty; Regime Switching Model*
JEL Classifications: C22; E31

1. INTRODUCTION

The impact of inflation on inflation uncertainty is well documented in theory. The empirical literature on this topic is also sizeable. However, the findings on the nature of the relationship and the empirical support for the link involving these variables vary with mixed results. Consequently, the Friedman-Ball hypothesis (Friedman, 1977; and Ball, 1992), which states that the effect of inflation on its uncertainty is positive, has found empirical support in varying degrees[1]. While differences in the sample periods and frequencies of the data sets used could account for mixed results, more importantly, it is the methodology or modelling approach applied that would explain the varied findings.

It is worth mentioning that in many such empirical studies, a standard auxiliary assumption typically made is that the parameters depicting the relationship are constant over the entire sample period. This means that the causal links are assumed to be stable over time. However, this assumption is far from innocuous and may often not hold in practice. Furthermore, there is considerable evidence supporting the view that the economic time series are best thought of as being generated by processes with time dependent parameters. According to several studies, including Fischer (1993), Stock and Watson (1996), Caglayan and Filiztekin (2003), Arghyrou et al. (2005) and Lanne (2006), social, economic and political changes are likely to

® Kushal Banik Chowdhury, Economic Research Unit, Indian Statistical Unit, 203 B. T. Road, Kolkata-700108, India, (email: kush.kolkata@gmail.com), Tel: +919051906848.
Nityananda Sarkar, Indian Statistical Institute, 203 B.T. Road, Kolkata-700108, India, (email: tanya@isical.ac.in).

[1] See, for details on the evidence of positive effect of inflation on inflation uncertainty, Grier and Perry (1998), Fountas (2001), Fountas et al. (2002), Kontonikas (2004), Daal et al. (2005), and Fountas and Karanasos (2007), and for negative effect, Engle (1983), Bollerslev (1986), Cosimano and Jansen (1988), Hwang (2001), and Wilson (2006).

make the macroeconomic relationship with constant parameter linear models quite untenable, especially when the length of the time series is long enough.

In fact, a recent trend in the literature on the relationship between inflation and inflation uncertainty is that this relationship is assumed to be neither linear nor stable over time. While there are a number of modelling procedures that incorporate these important aspects, an important class of such models is the regime switching models. In case of inflation, regime switching models are quite relevant and appropriate since inflation, by its very nature as a macro variable, may change for a period of time before reverting back to its original behaviour or switching to yet another style of behaviour, often due to intervention by government and/ or regularity authorities. The issue of regime switching behaviour of inflation while dealing with inflation uncertainty, was first raised by Evans and Wachtel (1993). They concluded that adequate attention needs to be paid to the consequences of regime switches; otherwise, existing models without any consideration to regime changes would seriously underestimate both the degree of uncertainty of inflation and its impact. Following their lead, some researchers including Kim (1993), Kim and Nelson (1999), Bhar and Hamori (2004), Chang and He (2010), and Chang (2012) have studied the relationship between inflation and its uncertainty where the regime specific behaviour have been duly incorporated in the model. A point worth noting is that all the above-mentioned studies have used the concept of unobserved regime and accordingly applied the well-known Markov switching regression model[2].

There is another class of regime switching models where regimes can be characterized by a threshold variable that is observable (see, Tong and Lim, 1980; Chan and Tong, 1986; Granger and Teräsvirta, 1993; Teräsvirta, 1998; Li and Li, 1996; and Brooks, 2001; for details on such models). This work, which essentially studies the impact of inflation on inflation uncertainty, takes be the past level of inflation as a threshold variable. The notion of taking the level of inflation as the observed threshold variable has been influenced by a few recent studies. For instance, Baillie et al. (1996) applied an autoregressive fractionally integrated moving average (ARFIMA) model to describe the long memory process of inflation dynamics for ten countries. They found that for six low inflation countries viz., Canada, France, Germany, Italy, Japan, and the USA, there is no apparent relationship between mean and variance of inflation. However, for the high inflation economics of Argentina, Brazil, Israel, and the UK, there is strong support for the Friedman hypothesis. In a recent study, Chen et al. (2008) have observed that the effect of inflation on inflation uncertainty is asymmetric. To be more specific, they have found a U-shaped pattern for four countries of East Asia, based on a nonlinear flexible regression model of Hamilton (2001), suggesting that inflation uncertainty is more sensitive to inflation in an inflationary period than in a deflationary period. Thus, it is relevant as well as important to examine the dynamic interaction between inflation and inflation uncertainty in the different regimes considered, and also to find if these interactions are different across the different regimes.

The purpose of this study is to examine the effect of inflation on inflation uncertainty in a regime switching modelling framework where regimes are determined by the past level of inflation. This model has two other additional features as well. The first is that not only the conditional mean model but also the conditional variance model is regime specific. Since (G)ARCH (Engle, 1982; and Bollerslev, 1986) or similar other models are preferred when

[2] In this class of Markov switching regression models, regimes are not deterministic and can be determined by an underlying unobservable stochastic process depending upon the probabilities assigned to the occurrence of the different regimes.

measuring inflation uncertainty, we too have applied the GARCH model; accordingly, the proposed model is a double-threshold GARCH (DTGARCH) model, originally proposed by Li and Li (1996) in the context of stock returns[3]. The second feature for our model involves the extension of the usual GARCH specification by explicitly incorporating a lagged inflation term in the conditional variance specification, the coefficient of which depicts the impact of inflation on inflation uncertainty. One distinct advantage of this model is that it allows for different behaviours of inflation uncertainty in different regimes, including the one captured through the lagged inflation term. We have applied this model to the time series covering the period from January 1970 to December 2013 for all members of the G7 countries. In this context it may be mentioned that since we have considered a long enough time period, we have specifically considered the issue of structural break in our study. However, unlike others, we have dealt with the twin issues of structural break and stationarity together by applying the recently developed tests of Perron and Yabu (2009) and Kim and Perron (2009).

The format of the paper is as follows: Section 2 describes the proposed model. Section 3 discusses empirical findings. Section 4 ends the paper with some concluding observations.

2. THE PROPOSED MODEL

Fountas et al. (2000) first considered the model for inflation uncertainty to explicitly include a lag inflation term. Like most studies, they used the GARCH (1,1)[4] model as the basic model to measure inflation uncertainty and then augmented it by including the first lag of inflation. The conditional mean model was taken to be the usual AR model. Thus, their model, designated as the AR(k)-GARCH(1,1)L(1)[5], consists of the following specifications for the conditional mean and conditional variance.

$$\pi_t = \phi_0 + \phi_1 \pi_{t-1} + \phi_2 \pi_{t-2} + \cdots + \phi_k \pi_{t-k} + \varepsilon_t, \qquad \varepsilon_t|\Psi_{t-1} \sim N(0, h_{\pi,t}) \tag{2.1}$$
$$h_{\pi,t} = \omega + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1} + \theta\pi_{t-1} \tag{2.2}$$

where π_t stands for the value of stationary inflation at time t, $\Psi_{t-1} = \{\pi_{t-1}, \pi_{t-2}, \ldots\}$ is the information set at $t-1$, $h_{\pi,t}$ is the conditional variance of π_t at t, and k the optimal lag order in the conditional mean specification. All roots of the polynomial $(1 - \phi_1 B - \cdots - \phi_k B^k) = 0$ are assumed to lie outside the unit circle for stationarity of π_t, and the usual restrictions on the parameter in $h_{\pi,t}$ are assumed to ensure positivity. Here the coefficient θ depicts the impact of inflation on inflation uncertainty. Obviously, a significant positive value of θ provides empirical support to the Friedman-Ball hypothesis for a given time series on inflation. We use this model as a benchmark model to find if the introduction of regimes in both inflation and inflation uncertainty leads to better understanding and modelling involving these two variables.

In describing the proposed model, which is a regime-dependent model with two regimes, we allow the effect of inflation on its uncertainty to be different in the two regimes. As already

[3] See also, Chen (1998), Brooks (2001), Chen et al. (2003), Chen and So (2006), Chen et al. (2006), and Yang and Chang (2008) for applications of the DTGARCH model for other financial variables.

[4] For our study, the conditional mean model has been taken to be the AR model on the basis of behaviour of the autocorrelation function (ACF) and partial autocorrelation function (PACF) of the inflation series. The value of k has been determined by the Akaike's information criterion (AIC). As regards the values of p and q of the GARCH(p,q) specification for conditional variance, $p = q = 1$ has been found to be adequate, and hence GARCH(1,1) model has been considered.

[5] 'L(1)' stands for the fact that the specification for $h_{\pi,t}$ includes the first lag of inflation as a separate term.

stated in the preceding section, we implicitly assume that the threshold variable which defines the regime that occurs at any given point in time, is known, and takes the preceding value of inflation i.e., π_{t-1}, to be the threshold variable. We define the two regimes according to the value of stationary inflation at $t-1$, i.e., $\pi_{t-1}>0$ and $\pi_{t-1}\leq 0$, which are being labelled as high and low inflation regimes[6], respectively.

The proposed model is a generalization of the model in Fountas et al. (2000) in that here regimes are introduced in the specifications of both the conditional mean and conditional variance of inflation. Obviously, the proposed model is an extension of the DTGARCH model, where the conditional variance specification includes a lag inflation term. The resultant model, denoted as DTGARCH(1,1)L(1), is thus specified as follows:

$$\pi_t=(\phi_0^l+\phi_1^l\pi_{t-1}+\phi_2^l\pi_{t-2}+\cdots+\phi_k^l\pi_{t-k})\,I[\pi_{t-1}\leq 0]+ \tag{2.3}$$
$$+(\phi_0^h+\phi_1^h\pi_{t-1}+\phi_2^h\pi_{t-2}+\cdots+\phi_k^h\pi_{t-k})\,(1-I[\pi_{t-1}\leq 0])+\varepsilon_t, \qquad \varepsilon_t|\Psi_{t-1}\sim N(0,h_{\pi,t})$$
$$h_{\pi,t}=(\omega^l+\alpha^l\varepsilon_{t-1}^2+\beta^l h_{t-1}+\theta^l\pi_{t-1})\,I[\pi_{t-1}\leq 0]+ \tag{2.4}$$
$$+(\omega^h+\alpha^h\varepsilon_{t-1}^2+\beta^h h_{t-1}+\theta^h\pi_{t-1})\,(1-I[\pi_{t-1}\leq 0])$$

where $I[.]$ is the indicator function which takes the value 1 if $\pi_{t-1}\leq 0$ and 0 otherwise, and $\Psi_{t-1}=\{\pi_{t-1},\pi_{t-2},\ldots\}$ is the information set at $t-1$. In this model coefficient vector $\xi^l=(\phi_0^l,\phi_1^l,\ldots,\phi_k^l,\omega^l,\alpha^l,\beta^l,\theta^l)'$ comprises the coefficients in both conditional mean and conditional variance specifications for the low inflation regime and similarly for the high inflation regime $\xi^h=(\phi_0^h,\phi_1^h,\ldots,\phi_k^h,\omega^h,\alpha^h,\beta^h,\theta^h)'$. Apart from the usual GARCH coefficients, θ^l and θ^h denote the effects of inflation on inflation uncertainty in the low and high inflation regimes, respectively.

The parameters of each of the benchmark and the proposed models i.e., DTGARCH(1,1)L(1) and AR(k)-GARCH(1,1)L(1) models have been estimated by the maximum likelihood (ML) method of estimation under the assumption of normality[7] for the conditional distribution of ε_t. For the iterative optimization procedure involved in the estimation process, the well-known Berndt-Hall-Hall-Hausman (BHHH; Berndt et al., 1974) algorithm has been used. The necessary computations have been done in GAUSS.

3. EMPIRICAL ANALYSIS

Like most of the empirical studies on the relationship between inflation and inflation uncertainty (see, for instance, Grier and Perry, 1998; Fountas et al., 2002; Fountas et al., 2004; Bredin and Fountas, 2005), we have used monthly data for this study. We have taken monthly time series of consumer price index (CPI) as a measure of the monthly price level for each of the G7 countries viz., Canada, France, Germany, Italy, Japan, the UK, and the USA. These time series on CPI have been downloaded from the official website of Federal Reserve Bank of St. Louis (http://research.stlouisfed.org/). Since the available time series on CPI is not adjusted for seasonality, each series has been adjusted for seasonality by applying the X12-ARIMA filtering method. The time period considered is from January 1970 to June

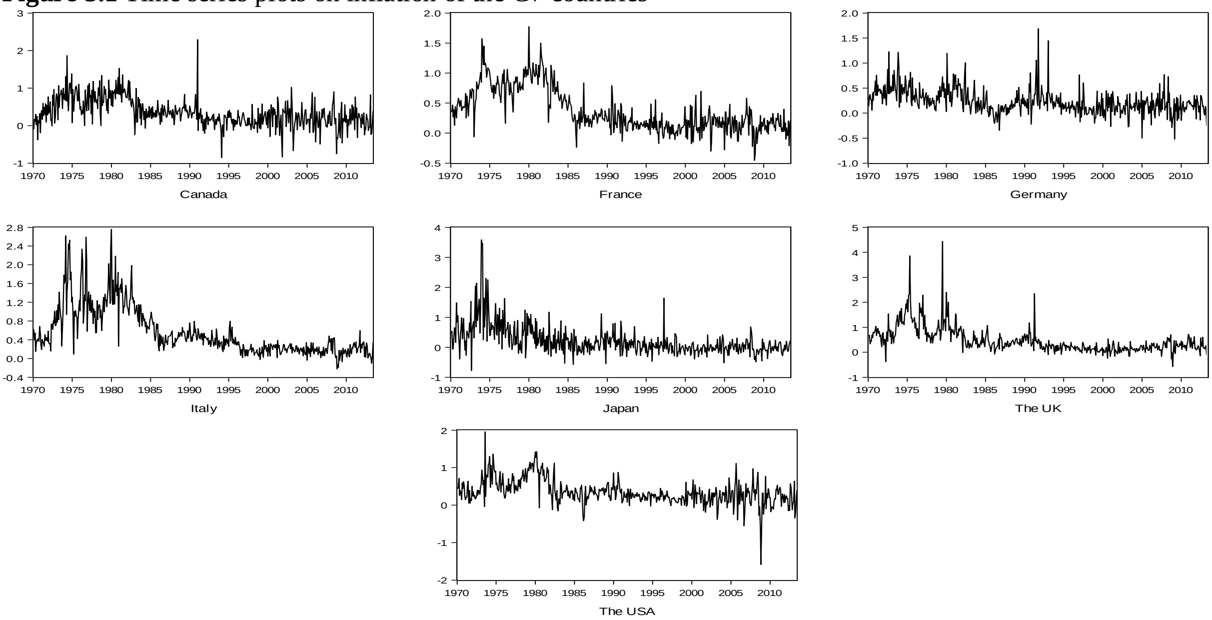
[6] The reasons for this labelling are stated in the next section (p. 41).
[7] It may be noted that like many financial time series, inflation has been found to be non-normal. However, assumption of normal distribution for estimation is very common since the resulting estimates often converge unlike some non-normal error distributions. In this particular literature on modelling inflation, the assumption of normal distribution for the error is most common, and hence we have also made this assumption of normal distribution for our analysis.

2013. The total number of observations is 522. The time series on inflation, $\tilde{\pi}_t$, has been obtained as the differences of the logarithm values of the CPI in percentage i.e.,

$$\tilde{\pi}_t = \log(CPI_t/CPI_{t-1})*100$$

The time series of inflation for all the G7 countries are plotted in Figures 3.1. It is visually evident that all the inflation series follow a trend - although segmented in nature, especially in case of Canada, France, Italy, the UK, and the USA. Further, it appears from the plots that there may be at least one structural break in each of the seven time series on inflation.

Figure 3.1 Time series plots on inflation of the G7 countries



We present the summary statistics on inflation, covering the sample period, for all the seven countries in Table 3.1 below.

Country	*Canada*	*France*	*Germany*	*Italy*	*Japan*	*The UK*	*The USA*
Mean	0.346	0.372	0.233	0.555	0.218	0.472	0.349
Std. dev.	0.379	0.359	0.251	0.508	0.498	0.513	0.331
Skewness	0.613	0.849	1.172	1.572	2.620	2.514	0.329
Kurtosis	4.765	3.245	7.066	5.594	14.544	14.298	6.650
Jarque-Bera	100.33*	63.92*	478.22*	360.62*	3489.17*	3319.76*	298.62*

Table 3.1 Descriptive statistics on inflation.
Notes: * indicates significance at 1% level of significance.

From this table it is evident that among the seven countries, Italy exhibits the highest average inflation of 0.555 per cent with the standard deviation of 0.508 per cent. The three countries viz., Italy, Japan and the UK have larger standard deviations than the remaining four countries. The skewness and kurtosis values indicate that all the seven inflation series are positively skewed and have leptokurtic distributions. Worthwhile to highlight are the high values of skewness for Japan and the UK at 2.620 and 2.514, respectively. As for the kurtosis, Japan again has the highest value of 14.544, followed by the UK with a value of 14.298. The kurtosis values for Germany and Italy are also moderately high. Therefore, as expected, results of the Jarque-Bera normality test clearly indicate that the null hypothesis of normality is rejected for all the inflation series, in particular, with very high test statistic value for countries like Japan and the UK.

We now discuss the stationarity/non-stationarity status of the inflation series. Since structural change in a time series affects inference on unit roots of the series, we examined these two issues together following the recently-developed procedures of Perron and Yabu (2009) and Kim and Perron (2009).

3.1. Stationarity and Structural Break

Since 1960s most of the industrialized countries including the G7 have experienced long-term swings in levels of inflation. Inflation had progressively risen in the 1960s and 1970s before it declined in the 1980s. Inflation further declined in the early to mid-1990s and since then remained low and stable (see, for details, Blinder, 1982; DeLong, 1997; Clarida et al., 2000; Orphanides, 2003; Meltzer, 2005; and Nelson, 2005). These observations have led many researchers to analyse the statistical properties of inflation persistence over the last two decades. This statistical issue has special relevance for inflation since in the earlier studies on developed economies, especially those on the UK and the USA, the unit root status of the time series were found to be mixed (see Kontonikas, 2004; Bhar and Hamori, 2004; Fountas et al., 2004; Daal et al., 2005; Conrad and Karanasos, 2006; and Fountas et al., 2006; for details). Thus, in this context, the span of the data sets used in this study viz., January 1970 - June 2013, is quite large. Hence, structural breaks in the trend function of these series are likely to occur, and as Perron (1989) first pointed out, disregarding this could mislead the conclusion on the presence of unit roots in the series.

To test for stationarity of a time series, most often the augmented Dicky-Fuller (ADF; Dicky and Fuller, 1979) test is used, where the null hypothesis of non-stationarity is tested against the alternative of stationarity. However, due to the influential work by Perron (1989), the commonly used ADF test has been criticized because of its bias towards non-rejection of the null hypothesis of a unit root against the alternative of trend stationarity in the presence of a structural break in the deterministic trend, as well as for its low power for near integrated process. Subsequently, Perron (1989, 1990) proposed an alternative unit root test that allows for the possibility of a structural break in the trend function under both the null and alternative hypotheses. However, one serious limitation of this test is that it is based on the assumption of a known break date. Subsequently, Zivot and Andrews (1992), Perron (1997), and Vogelsang and Perron (1998), among others, have treated the break date to be unknown, which is endogenously determined from the model.

However, recently Kim and Perron (2009) have pointed out that in all these tests with an endogenous break point i.e., those by Zivot and Andrews (1992), Perron (1997), and Vogelsang and Perron (1998), a trend break is not allowed under the null hypothesis. These tests consider a break in the time series under the alternative only. Hence, tests of this kind are likely to be affected in terms of size and power. To overcome this limitation, Kim and Perron (2009) have developed a new test procedure on the line of Perron's (1989) original formulation of trend break being allowed under both the null and alternative hypotheses, but the break date is now assumed to be unknown.

Now, prior to applying this unit root test, knowing whether a structural break is present in a given time series is crucial, especially when the break date is assumed to be unknown, as in the case with the Kim-Perron (2009) test. Further, in this context, in the absence of a trend break, the well-known ADF test has the highest power compared to any other alternative test, and it is the most appropriate test for testing the presence of unit roots in a time series. It is, therefore, all the more important that the knowledge of the presence or absence of a break in the trend function is available before deciding on the appropriate unit root test. However, as pointed out by Perron and Yabu (2009), testing for a structural break in the trend function depends on whether the noise component is stationary or non-stationary with unit roots. Thus there is some sort of a circular problem in testing for these twin issues. To deal with this, recently Perron and Yabu (2009) proposed a test for structural change in the trend function of a univariate time series, which can be performed without any prior knowledge on whether the noise component is stationary or non-stationary containing unit roots[8]. Since this test is very general in its approach insofar as the assumption on noise is concerned, we have performed this test to detect the presence of structural break, if any, in the trend function of a time series. Thus, in case the Perron-Yabu test suggests that there is no break in the deterministic trend, we apply the usual ADF test; otherwise, we use the Kim-Perron test to test the null hypothesis of unit roots against the alternative of stationarity with a break in the deterministic trend function under both hypotheses[9].

Three models - *Model I, Model II* and *Model III* - representing a single change in intercept only, a one-time change in the slope of linear trend function without a change in level, and a simultaneous change in the intercept and the slope coefficients, respectively, are considered for the Perron-Yabu test. For our study, we have considered *Model III* only. The test statistic values with trimming percentage being 0.15 are reported in Table 3.2.

Country	***Perron-Yabu structural break test***	***Kim-Perron unit root test***	***Estimated break date***
Canada	53.029*	-21.615*	June 1982
France	48.350*	-8.000*	September 1983
Germany	8.741*	-12.948*	October 1982
Italy	12.557*	-5.034*	December 1982
Japan	21.985*	-12.279*	December 1976
The UK	41.028*	-7.767*	December 1981
The USA	27.492*	-15.488*	September 1981

Table 3.2 Results of tests for structural break and unit roots in inflation series.
Notes: *, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively.

The Perron-Yabu test statistic values are significant for all seven inflation series, which clearly establish that either or both of the intercept and slope of the trend function have undergone structural change. Thus the test confirms the presence of one significant change in the deterministic trend of inflation for each of the G7 countries.

With this empirical finding on the presence of a structural break in the trend function of inflation, we have applied the third variant of additive outlier model[10], *Model* A3 in Kim and Perron (2009), to each of the seven inflation series to test the null hypothesis of unit roots with a deterministic trend break against the alternative of stationarity with a break in the deterministic trend. The values of this test statistic are reported in the 3rd column of Table 3.2. The computed values of the *t*-ratio are compared with the appropriate critical values available

[8] Perron and Yabu (2009) have considered a quasi-feasible generalized least squares technique that uses a super-efficient estimate of the sum of autoregressive parameters, which governs the stationary or integrated behaviour of a time series.
[9] A GAUSS program code for Perron and Yabu (2009) structural break test and a MATLAB code for Kim and Perron (2009) unit root test in presence of a structural break have been taken from the official website of Pierre Perron.
[10] Kim and Perron (2009) have considered three types of additive outlier models viz., A1, A2 and A3, representing the occurrence of a structural break in the intercept only, in the slope only, and both in the intercept and slope coefficients of the trend function, respectively.

in Perron (1989) and Perron and Vogelsang (1993). The findings clearly suggest that the inflation series of each of these countries are stationary when accounting for the presence of one structural break. In each case, the null hypothesis of a unit root with a trend break against the alternative of a stationary process with a trend break is rejected. We thus conclude that the underlying data generating process for each of the seven inflation series is a trend stationary process (TSP), having a structural break in the respective deterministic trend functions.

The stationary time series on inflation for each country, denoted by π_t, has then been obtained after removing the segmented deterministic trend from $\tilde{\pi}_t$. Specifically, this has been obtained by regressing $\tilde{\pi}_t$ on an intercept, an intercept dummy, a linear trend term and a trend (linear) dummy, and then collecting the detrended series, π_t, which is now stationary in the sense of having neither any stochastic trend or any deterministic trend. Thus the obtained stationary time series on inflation has been used for all subsequent analyses. In this context it may be noted that the two inflation regimes, defined as $\pi_{t-1} \leq 0$ and $\pi_{t-1} > 0$ in Section 2, essentially refer to $\tilde{\pi}_{t-1}$ being less than or equal to and greater than some positive values, as indicated by their respective deterministic trend function of each inflation series. Hence the two regimes are referred to as the low-inflation and high-inflation regimes, respectively.

The estimated break dates are reported in the 4^{th} column of Table 3.2. The estimated break dates refer to the period of 1970s and 1980s. Thus these findings provide empirical support to the general observation that the high phase of inflation during the 1960s-1970s started reducing substantially towards stabilization in the 1980s. There are quite a few studies supporting this observation as well. For instance, in their study, Cecchetti and Krause (2001) have reported that inflation in most developed and developing countries remained at a very high level during the period of 'great inflation' in 1960s, which was coupled with the oil price shock in 1973. The volatility appears to be more pronounced during that period while remaining low and stable during the latter half of 1980s due to marked improvement in macroeconomic performances in those countries. This finding has also been supported by Stock and Watson (2002), who pointed out that during the period of 1990s, not only volatility of inflation decreased sharply but also average inflation underwent a downward trend. Further, Krause (2003) reported that in a cross-section of 63 countries, mean inflation has fallen from approximately 83 per cent in the pre-1990 period to approximately 9 per cent in the latter half of 1990s. Given such empirical findings, it is expected that there would be at least one structural break in the time series on inflation, and this is, in fact, what we have found[11].

3.2. Empirical Findings on The Impact of Inflation on Inflation Uncertainty

In this section we first present the estimates of the benchmark AR(k)-GARCH(1,1)L(1) model. Next we discuss the results of estimation and testing of hypotheses of interests involving the parameters of the proposed model.

3.2.1. The Benchmark AR(k)-GARCH(1,1)L(1) *Model*

The results of estimation of the benchmark AR(k)-GARCH(1,1)L(1) model specified in equations (2.1) and (2.2) are reported in Table 3.3. As noted in the preceding section, the

[11] Clark (2006) and Levin and Piger (2003) allowed the possibility of structural breaks while examining the persistence of inflation. Rapach and Wohar (2005) also tested for breaks and found evidence of shift in the level of inflation in a wide range of European countries.

optimal lag orders of the autoregressive terms for the seven inflation series have been obtained by AIC.

	Canada	*France*	*Germany*	*Italy*	*Japan*	*The UK*	*The USA*
Conditional mean							
ϕ_0	0.007	0.006	0.003	-0.004	0.003	-0.012	-0.004
ϕ_1	-0.012	0.305*	0.078	0.266*	0.082***	0.208*	0.351*
ϕ_2	0.141*	0.026	0.184*	0.199*	0.037	0.298*	0.016
ϕ_3	0.094**	0.114**	0.110**	0.150*	0.021	0.124**	-0.008
ϕ_4	0.049	0.020	0.058	0.032	0.043	-0.096**	0.099**
ϕ_5	0.113**	-0.018	0.015	-0.016	0.114**	0.057	-0.029
ϕ_6	0.016	0.052	0.117*	0.119**	0.029	0.114*	0.019
ϕ_7	0.035	0.080***	0.038	0.091***	0.082***	-	0.067***
ϕ_8	-	0.045	0.067***	0.032	0.025	-	-0.058
ϕ_9	-	-0.003	-	0.027	0.150*	-	0.062
ϕ_{10}	-	0.057***	-	0.026	0.044	-	0.055
ϕ_{11}	-	-	-	-0.046	0.051	-	-
ϕ_{12}	-	-	-	-0.125*	-0.133*	-	-
Conditional variance							
ω	0.011*	0.009**	0.013*	0.001***	0.002***	0.008*	0.004*
α	0.237*	0.181*	0.464*	0.160*	0.080**	0.675*	0.271*
β	0.662*	0.553*	0.347*	0.830*	0.896*	0.437*	0.692*
θ	0.025**	0.007	0.023	0.002	-0.007	-0.007	0.024*
MLLV	-75.64	161.00	97.86	182.77	-121.21	-6.780	75.66

Table 3.3 Estimates of the parameters of AR(k)-GARCH(1,1)L(1) model.
Notes: *, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively. MLLV is the maximized log-likelihood value. Some entries in the table are blank since the estimates of the parameters have been reported, as expectedly, up to the lag (k) values which have been found by the AIC for the mean models of the respective countries. Evidently, the value of k has not been found to be the same for all countries.

According to the results reported in this table, the autoregressive coefficients are significant in varying numbers across the seven countries. The usual GARCH parameters viz., ω, α and β are statistically significant for all inflation series. Additionally, in agreement with the Friedman-Ball hypothesis, the estimate of the lagged inflation coefficient θ, is positive and statistically significant only for two countries viz., Canada and the USA, while for the remaining countries no significant relationship exists. It is quite possible that this finding of no such significant effect of inflation on its uncertainty for most viz., five, of the G7 countries may be due to the fact that this model does not factor for regime-specific inflation behaviour, which might have masked potentially different realizations due to probable regime shift in inflation series, especially because the span of the data set is quite large

3.2.2. The Proposed DTGARCH(1,1)L(1) *Model*

The DTGARCH(1,1)L(1) model, as specified in equations (2.3) and (2.4), incorporates differential behaviour in both inflation and inflation uncertainty from consideration of regime switching and also allows for the coefficient attached to the lagged inflation term in inflation uncertainty model specified in equation (2.4) to be different in the two regimes. This model is locally i.e., regime-wise linear but overall nonlinear in nature. The estimates[12] of the parameters of this model are presented in Table 3.4.

[12] For writing the GAUSS codes of DTGARCH (1,1)L(1) model, we have made use of codes from the GAUSS programs developed by Philip Hans Franses and Dick Van Dijk on different volatility models, which were downloaded from http://www.few.eur.nl/few/few/people/frances.

	Canada	France	Germany	Italy	Japan	The UK	The USA
Conditional mean for the low-inflation regime							
ϕ_0^l	0.007	0.005	0.010	-0.011	-0.022	-0.053*	-0.001
ϕ_1^l	0.069	0.326*	0.141	0.261*	-0.020	-0.050	0.394*
ϕ_2^l	0.087	-0.033	0.134**	0.283*	0.014	0.328*	0.025
ϕ_3^l	0.153**	0.029	0.163**	0.046	0.022	0.215*	-0.079
ϕ_4^l	0.078	0.171**	0.109***	-0.026	0.013	-0.084	0.155**
ϕ_5^l	0.160**	-0.092	-0.011	0.018	0.109***	-0.055	-0.031
ϕ_6^l	-0.073	-0.031	0.120**	0.113**	-0.009	0.178*	0.121***
ϕ_7^l	0.000	0.116	0.052	0.186*	0.100***	-	-0.021
ϕ_8^l	-	0.057	0.089***	-0.035	0.009	-	-0.106
ϕ_9^l	-	0.041	-	-0.043	0.112**	-	0.130**
ϕ_{10}^l	-	0.075	-	0.049	0.094	-	0.086
ϕ_{11}^l	-	-	-	-0.029	0.013	-	-
ϕ_{12}^l	-	-	-	-0.152*	-0.084	-	-
Conditional mean for the high-inflation regime							
ϕ_0^h	0.018	-0.015	0.014	-0.006	-0.003	-0.018	-0.006
ϕ_1^h	-0.044	0.401*	0.034	0.221**	0.121	0.425*	0.325*
ϕ_2^h	0.276*	0.141***	0.208*	0.217*	0.094	0.263*	0.020
ϕ_3^h	0.032	0.169**	0.047	0.043	0.097	0.029	0.019
ϕ_4^h	-0.004	-0.072	-0.003	0.157*	0.099	-0.031	0.040
ϕ_5^h	0.055	0.023	0.030	0.089***	0.129***	0.147*	-0.013
ϕ_6^h	0.020	0.125	0.114**	0.256*	0.001	0.138*	-0.012
ϕ_7^h	0.089	0.021	0.005	-0.020	0.076	-	0.146**
ϕ_8^h	-	0.042	0.072	-0.034	-0.030	-	-0.001
ϕ_9^h	-	-0.047	-	0.137*	0.144**	-	-0.040
ϕ_{10}^h	-	-0.007	-	0.139**	0.062	-	0.027
ϕ_{11}^h	-	-	-	-0.157*	0.060	-	-
ϕ_{12}^h	-	-	-	-0.251*	-0.195*	-	-
Conditional variance for the low-inflation regime							
ω^l	0.027**	0.004**	0.011***	0.002	0.006	0.002	0.004***
α^l	0.605*	0.363**	0.604*	0.255***	0.000	0.640*	0.304*
β^l	0.468*	0.710*	0.492*	0.323*	0.555*	0.529*	0.619*
θ^l	0.111**	0.034***	0.046	-0.055*	-0.154**	-0.017	0.010
Conditional variance for the high-inflation regime							
ω^h	0.002	0.033*	0.008	0.000	0.000	0.009**	0.000
α^h	0.000	0.120	0.000	0.948*	0.222**	0.513*	0.001
β^h	0.658*	0.001	0.280*	0.020	0.837*	0.000	0.655*
θ^h	0.119*	-0.039	0.154*	0.153*	-0.003	0.198*	0.124*
MLLV	-63.51	172.46	104.54	198.89	-116.26	16.39	88.40
Wald test for							
$H_0: \theta^l = \theta^h$	0.02	2.69	5.42**	33.89*	3.51***	13.40*	11.19*

Table 3.4 Estimates of the parameters of the DTGARCH(1,1)L(1) model.
Notes: *, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively. MLLV is the maximized log-likelihood value. Some entries in the table are blank since the estimates of the parameters have been reported, as expectedly, up to the lag (k) values which have been found by the AIC for the mean models of the respective countries. Evidently, the value of k has not been found to be the same for all countries.

Several findings from the estimation results are worth mentioning: First, the estimates of the parameters in conditional mean clearly establish regime switching behaviour in inflation for all seven series since some of the own lags in both the regimes are found to be statistically significant. Second, looking at the parameters of the model for the conditional variance in the two inflation regimes, we first note that in case of Canada, France, Germany, Japan, and the USA, only one of the parameters of α^l and α^h is significant, while for Italy and the UK, both parameters are significant. Further, only one of the parameters β^l and β^h are significant for the three countries viz., France, Italy, and the UK, and for the remaining ones i.e., Canada, Germany, Japan, and the USA, both coefficients are significant. Considering two regimes based on a past level of inflation for the conditional variance which measures inflation uncertainty, is found to be relevant and statistically meaningful. Regarding the effect of inflation on inflation uncertainty, i.e., θ^l and θ^h, we note that both these coefficients are significant in two countries only, namely, Canada and Italy, while for the remaining five countries, at least one of θ^l and θ^h is significant. This establishes the fact that the effect of inflation on inflation uncertainty is also regime specific. Hence, on the whole, any policy measure on inflation should be based, inter alia, on regime consideration of inflation.

Finally, the most important hypothesis for the proposed model is whether inflation affects inflation uncertainty differently in the two inflation regimes. In terms of the parameters, the null and alternative hypotheses are $H_0\colon \theta^l = \theta^h$ and $H_1\colon \theta^l \neq \theta^h$, respectively. This null hypothesis has been tested by using the Wald test, and the test statistic values for the seven series are reported in the last row of Table 3.4. The results of the Wald test show that the null hypothesis is rejected for all but Canada and France. Regarding the latter two countries, we can conclude that in the case of Canada, inflation has a positive impact on inflation uncertainty and does not vary with the regime change while for France, the only conclusion that can be drawn is that the coefficients do not change significantly between the two regimes. The Wald test results thus suggest that significant difference in the two regimes exists insofar the effect of inflation on inflation uncertainty is concerned in case of five members of the G7 countries. These countries are Germany, Italy, Japan, the UK, and the USA.

The positive and significant values of θ^h in four of these countries viz., Germany, Italy, the UK, and the USA indicate that inflation increases inflation uncertainty at the high-inflation regime in each of these countries, while the effect is insignificant for Germany, the UK, and the USA at the low-inflation regime and negative for Italy. This evidence for Germany, the UK, and the USA thus supports the findings of Ungar and Zilberfarb (1993) viz., that inflation affects inflation uncertainty only in the high-inflation regime but not in the low-regime. The finding that $\hat{\theta}^l$ is negative and significant for Italy and Japan is, however, somewhat unusual although not exceptional. This means that at the low-regime inflation has a negative effect on inflation uncertainty or in other words, a reduction in inflation in the low-regime increases inflation uncertainty, and thus the Friedman-Ball hypothesis fails to hold for these two countries. Such evidence has been found by a few other studies as well – although with different volatility specifications. For instance, based on a study on 12 European Monetary Union countries, Caporale and Kontonikas (2009) have found that during the post-1999 period when average inflation was low, a further reduction in inflation led to an increase rather than a decrease in inflation uncertainty for a number of countries in the Euro Zone. In a recent study based on monthly data on US inflation over the period from 1926 to 1992, Hwang (2001) has found that inflation affects its uncertainty weakly and negatively during the periods of both high and low inflation.

We report on the Ljung-Box test statistic values based on residuals of this model for all G7 countries in Table 3.5.

The test statistic has been computed for both the standardized and squared standardized residuals. It is evident from Table 3.5 that none of these are significant, and hence we conclude that the proposed models for both the conditional mean and conditional variance of inflation along with the chosen lag values are adequate for all G7 countries.

Country	$Q(1)$	$Q(5)$	$Q(10)$	$Q^2(1)$	$Q^2(5)$	$Q^2(10)$
Canada	0.007	3.746	7.027	0.670	1.386	4.238
France	0.107	1.975	4.710	0.014	4.348	5.481
Germany	0.369	2.123	4.823	0.638	3.656	7.972
Italy	0.024	2.434	11.420	2.281	5.418	12.963
Japan	0.027	0372	1.480	0.002	2.112	8.114
The UK	0.001	2.676	6.196	0.003	2.303	11.887
The USA	0.014	2.604	5.050	1.501	7.512	9.926

Table 3.5 Results of the Ljung-Box test with standardized residuals and squared standardized residuals.
Notes: *, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively. $Q(.)$ and $Q^2(.)$ denote the Ljung-Box test for autocorrelation in standardized residuals and squared standardized residuals, respectively.

Finally, we make a comparison between the benchmark AR(k)-GARCH(1,1)L(1) model and the proposed DTGARCH(1,1)L(1) model by the likelihood ratio (LR) test to find if introduction of regimes based on low and high inflation in both conditional mean and conditional variance has led to any significant gain in understanding the effects of inflation on inflation uncertainty. We report the LR test statistic values in Table 3.6.

Country	**AR(k)-GARCH(1,1)L(1)** ***versus*** **DTGARCH(1,1)L(1)**
Canada	24.26^{**}
France	22.92^{***}
Germany	13.36
Italy	32.24^{**}
Japan	9.90
The U.K.	46.34^{*}
The U.S.A.	25.48^{**}

Table 3.6 LR test statistic values.
Notes: *, ** and *** indicate significance at 1%, 5% and 10% levels of significance, respectively. $Q(.)$ and $Q^2(.)$ denote the Ljung-Box test for autocorrelation in standardized residuals and squared standardized residuals, respectively.

We observe from this table that the LR test statistic values are significant for five of the G7 countries viz., Canada, France, Italy, the UK, and the USA, and hence we can conclude that the proposed model explains the relationship ‘better’ than the benchmark model.

4. CONCLUSIONS

The effect of inflation on inflation uncertainty has been studied for the G7 countries in terms of a model called the DTGARCH(1,1)L(1) model, where the conditional mean as well as the conditional variance of inflation are based on a consideration of two regimes for inflation, and the conditional variance specification for each regime is assumed to be GARCH with an additional term of inflation with a lag of 1. The regimes are determined by the level of (stationary) inflation of the preceding lag being negative or positive.

The findings on the twin issues of structural break and stationarity of the series, based on the recently-developed tests of Perron and Yabu (2009) and Kim and Perron (2009), are that all of the series are trend stationary with a break in their respective trend functions. The estimated break date thus obtained for each country seems to broadly give empirical support to the phenomenon of 'great inflation', which refers to the high and volatile inflation that occurred in the mid-1960s and lasted for almost twenty years

Regarding the nature of the relationship, the empirical findings clearly show that the impact of inflation on inflation uncertainty is different in the two regimes for five of the G7 countries viz., Germany, Italy, Japan, the UK, and the USA. For Canada, the Friedman-Ball hypothesis for a positive effect of inflation on its uncertainty holds, but this effect is invariant to the two regimes. Further, the Friedman-Ball hypothesis holds only in the high-inflation regime for Germany, Italy, the UK, and the USA. On the other hand, the relationship is found to be insignificant for Germany, the UK, and the USA in the low-inflation regime, while it is negative and significant for Italy and Japan. This suggests that in the low-inflation regime a decline in inflation has led to an increase in inflation uncertainty for Italy and Japan, and this obviously counters the Friedman-Ball hypothesis. Thus, these findings, on one hand, provide strong empirical support to the Friedman-Ball hypothesis for high-inflation regimes of most of the G7 countries, and on the other, clearly establish the importance of considering regime-specific behaviour in modelling the impact of inflation on inflation uncertainty.

REFERENCES

Arghyrou, M., C. Martin and C. Milas (2005). Non-linear inflationary dynamics: evidence from the UK. *Oxford Economic Papers*, 57 (1), 51-69.

Baillie, R., C.F. Chung and M. Tieslau (1996). Analyzing inflation by the fractionally integrated ARFIMA-GARCH model. *Journal of Applied Econometrics*, 11 (1) , 23-40.

Ball, L. (1992). Why does higher inflation raise inflation uncertainty? *Journal of Monetary Economics*, 29 (3), 371-378.

Berndt, E.K., B.H. Hall, R.E. Hall and J.A. Hausman (1974). Estimation and inference in nonlinear structural models. *Annals of Economic and Social Measurement*, 3(4), 103-116.

Bhar, R. and S. Hamori (2004). The link between inflation and inflation uncertainty: evidence from G7 countries. *Empirical Economics,* 29 (4) , 825-853.

Blinder, A.S. (1982). The anatomy of double-digit inflation in the 1970s. In *NBER Chapters, in: Inflation: Causes and Effects,* 261-382. National Bureau of Economic Research.

Bollerslev, T. (1986). Generalised autoregressive conditional heteroskedasticity. *Journal of Econometrics*, 31 (3), 307-327.

Bredin, D. and S. Fountas (2005). Macroeconomic uncertainty and macroeconomic performance: Are they related? *The Manchester School*, 73 (s1), 58-76.

Brooks, C. (2001). A double-threshold GARCH model for the French/Deutschmark exchange rate. *Journal of Forecasting*, 20 (2), 135-143.

Caglayan, M., and A. Filiztekin (2003). Nonlinear impact of inflation on relative price variability. *Economics Letters*, 79 (2), 213-218.

Caporale, G.M. and A. Kontonikas (2009). The Euro and inflation uncertainty in the European Monetary Union. *Journal of International Money and Finance*, 28 (6), 954-971.

Cecchetti, S. G. and S. Krause (2001). Financial structure, macroeconomic stability and monetary policy. *NBER Working Paper* 8354.

Chan, K. S. and H. Tong (1986). On estimating thresholds in autoregressive models. *Journal of time series analysis*, 7 (3), 179-190.

Chang, K.L. (2012). The impacts of regime-switching structures and fat-tailed characteristics on the relationship between inflation and inflation uncertainty. *Journal of Macroeconomics*, 34 (2), 523-536.

Chang, K.L. and C.W. He (2010). Does the magnitude of the effect of inflation uncertainty on output growth depend on the level of inflation? Manchester School, 78 (2), 126-148.

Chen, C.W. (1998). A Bayesian analysis of generalized threshold autoregressive models. *Statistics & probability letters*, 40 (1), 15-22.

Chen, C.W. and M.K. So (2006). On a threshold heteroscedastic model. *International Journal of Forecasting*, 22 (1), 73-89.

Chen, C.W., T.C. Chiang and M.K. So (2003). Asymmetrical reaction to US stock-return news: evidence from major stock markets based on a double-threshold model. *Journal of Economics and Business*, 55 (5), 487-502.

Chen, C.W., M.J. Yang, R. Gerlach and H. Jim Lo (2006). The asymmetric ractions of mean and volatility of stock returns to domestic and international information based on a four-regime double-threshold GARCH model. *Physica A: Statistical Mechanics and its Applications*, 366 , 401-418.

Chen, S.W., C.H. Shen and Z. Xie (2008). Evidence of a nonlinear relationship between inflation and inflation uncertainty: The case of the four little dragons. *Journal of Policy Modeling*, 30 (2) , 363-376.

Clarida, R., J. Gali and M. Gertler (2000). Monetary policy rules and macroeconomic stability: evidence and some theory. *Quarterly Journal of Economics*, 115 (1), 147-180.

Clark, T.E. (2006). Disaggregate evidence on the persistence of consumer price inflation. *Journal of Applied Econometrics*, 21(5), 563-587.

Conrad, C. and M. Karanasos (2006). Dual long memory in inflation dynamics across countries of the Euro area and the link between inflation uncertainty and macroeconomic performance. *Studies in Nonlinear Dynamics and Econometrics*, 9 (4), 1-38.

Cosimano, T.F. and D.W. Jansen (1988). Estimates of the variance of U.S. inflation based upon the ARCH model. *Journal of Money, Credit and Banking*, 20 (3), 409-421.

Daal, E., A. Naka and B. Sanchez (2005). Re-examining inflation and inflation uncertainty in developed and emerging countries. *Economics Letters,* 89 (2), 180-186.

DeLong, J.B. (1997). America's peacement inflation. In *Reducing Inflation: Motivation and Strategy,* ed. C.D. Romer and D.H. Romer, University of Chicago Press, 247-280.

Dickey, D.A. and W.A. Fuller (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association*, 74, 427-431.

Engle, R.F. (1982). Autoregressive conditional heteroskedasticity with estimates of the variance of UK inflation. *Econometrica,* 50 (4), 987-1007.

Engle, R.F. (1983). Estimates of the Variance of US Inflation Based upon the ARCH Model. *Journal of Money, Credit and Banking*, 15 (3), 286-301.

Evans, M. and P. Wachtel (1993). Were price changes during the Great Depression anticipated? Evidence from nominal interest rates. *Journal of Monetary Economics*, 32 (1), 3-34.

Fischer, S. (1993). The role of macroeconomic factors in growth. *Journal of monetary economics*, 32 (3), 485-512.

Fountas, S. (2001). The relationship between inflation and inflation uncertainty in the UK: 1885–1998. *Economics Letters*, 74 (1), 77-83.

Fountas, S., A. Ioannidis and M. Karanasos (2004). Inflation, inflation uncertainty and a common European monetary policy. *Manchester School*, 72 (2), 221-242.

Fountas, S. and M. Karanasos (2007). Inflation, output growth, and nominal and real uncertainty: empirical evidence for the G7. *Journal of International Money and Finance*, 26 (2), 229-250.

Fountas, S., M. Karanasos and M. Karanassou (2000). A GARCH model of inflation and inflation uncertainty with simultaneous feedback. *Department of Economics, University of York , Discussion Papers* 00/24.

Fountas, S., M. Karanasos and J. Kim (2002). Inflation and output growth uncertainty and their relationship with inflation and output growth. *Economics Letters*, 75 (3), 293-301.

Fountas, S., M. Karanasos and J. Kim (2006). Inflation uncertainty, output growth uncertainty and macroeconomic performance. *Oxford Bulletin of Economics and Statistics*, 68 (3), 319-343.

Friedman, M. (1977). Nobel lecture: inflation and unemployment. *Journal of Political Economy*, 85 (3), 451-472.

Granger, C.W. and T. Teräsvirta (1993). *Modelling non-linear economic relationships*. OUP Catalogue.

Grier, K.B. and M.J. Perry (1998). On inflation and inflation uncertainty in the G7 countries. *Journal of International Money and Finance*, 17 (4) , 671-689.

Hamilton, J. (2001). A parametric approach to flexible nonlinear inference. *Econometrica*, 69 (3), 537-573.

Hwang, Y. (2001). Relationship between inflation rate and inflation uncertainty. *Economics Letters*, 73 (2), 179-186.

Kim, C.J. (1993). Unobserved-component models with Markov switching heteroskedasticity: changes in regime and link between inflation rates and inflation uncertainty. *Journal of Business and Economic Statistics*, 11 (3), 341-349.

Kim, C.J. and C.R. Nelson (1999). Has the US economy become more stable? A Bayesian approach based on a Markov-switching model of the business cycle. *Review of Economics and Statistics*, 81 (4), 608-616.

Kim, D. and P. Perron (2009). Unit root tests allowing for a break in the trend function at an unknown time under both the null and alternative hypothesis. *Journal of Econometrics,* 148 (1), 1-13.

Kontonikas, A. (2004). Inflation and inflation uncertainty in the United Kingdom, evidence from GARCH modelling. *Economic Modelling,* 21 (3), 525-543.

Krause, S. (2003). Measuring monetary policy efficiency in European Union countries. *Department of Economics, Emory University, Working Paper No.* 03-11.

Lanne, M. (2006). Nonlinear dynamics of interest rate and inflation. *Journal of Applied Econometrics*, 21 (8), 1157-1168.

Levin, A. T. and J. Piger (2003). Is inflation persistence intrinsic in industrial economies? *Federal Reserve Bank of St. Louis Working Paper* 2002-023E.

Li, C.W. and W.K. Li (1996). On a double-threshold autoregressive heteroscedastic time series model. *Journal of Applied Econometrics*, 11 (3), 253-274.

Meltzer, A.H. (2005). Origins of the great inflation. *Federal Reserve Bank of St. Louis Review*, 87 (2), 145-175.

Nelson, E. (2005). The great inflation of the seventies: what really happened? *The B.E. Journal of Macroeconomics*, 5 (1), 1-50.

Orphanides, A. (2003). The quest for prosperity without inflation. *Journal of Monetary Economics*, 50 (3), 633-663.

Perron, P. (1997). Further evidence on breaking trend functions in macroeconomic variables. Journal of Econometrics, 80 (2) , 355-385.

Perron, P. (1990). Testing for a unit root in a time series with changing mean. *Journal of Business & Economic Statistics*, 8 (2), 153-162.

Perron, P. (1989). The great crash, the oil price shock, and the unit root hypothesis. *Econometrica*, 57 (6), 1361-1401.

Perron, P. and T.J. Vogelsang (1993). The great crash, the oil price shock and the unit root hypothesis: Erratum. *Econometrica*, 61 (1), 248-249.

Perron, P. and T. Yabu (2009). Testing for shifts in trend with an integrated or stationary noise component. *Journal of Business and Economic Statistics*, 27 (3), 369-396.

Rapach, D. and M. Wohar (2005). Regime changes in international real interest rates: Are they a monetary phenomenon? *Journal of Money, Credit and Banking*, 37(5), 887-906.

Stock, J. H., and Watson, M. W. (1996). Evidence on structural instability in macroeconomic time series relations. Journal of Business & Economic Statistics, 14(1) , 11-30.

Stock, J.H. and M.W. Watson (2002). Has the business cycle changed and why? *NBER Macroeconomics Annual 2002*, 17, 159-230.

Teräsvirtra, T. (1998). Modelling economic relationships with smooth transition regressions. In *Handbook of Applied Economic Statistics*, ed. A. Ullah and D.E. Giles, 507-552. New York: Marcel Dekker.

Tong, H. and K.S. Lim (1980). Threshold autoregression, limit cycles and cyclical data. *Journal of the Royal Statistical Society, Series B (Methodological)*, 42 (3), 245-292.

Ungar, M. and B. Zilberfarb (1993). Inflation and its unpredictability - theory and empirical evidence. *Journal of Money, Credit and Banking*, 25 (4), 709-720.

Vogelsang, T.J. and P. Perron (1998). Additional tests for a unit root allowing for a break in the trend function at an unknown time. International Economic Review, 39 (4), 1073-1100.

Wilson, B.K. (2006). The links between inflation, inflation uncertainty and output growth: New time series evidence from Japan. *Journal of Macroeconomics*, 28 (3), 609-620.

Yang, Y.L. and C.L. Chang (2008). A double-threshold GARCH model of stock market and currency shocks on stock returns. *Mathematics and Computers in Simulation*, 79(3), 458-474.

Zivot, E. and D.K. Andrews (1992). Further evidence on the great crash, the oil price shock and the unit root hypothesis. *Journal of Business and Economic Statistics*, 10 (3) , 251-270.