A GARCH Model of Inflation and Inflation Uncertainty in Iran

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Abstract

The paper investigates the relationship between inflation and inflation uncertainty using the Iranian data over the period 1959:03 – 2008:02. GARCH models are used to examine this relationship. Granger methods are employed to provide statistical evidence for the relationship between average inflation and inflation uncertainty. Threshold GARCH (TGARCH) models are considered to investigate asymmetry in the conditional variance of inflation. The Component GARCH (CGARCH) models are employed to decompose inflation uncertainty into a short-run and a long-run component by permitting transitory deviations of the conditional volatility around a time-varying trend. This model examines the presence of long memory in the conditional variance of inflation.

The findings show that increased inflation raises inflation uncertainty confirming the theoretical predictions made by Friedman. Furthermore, the findings of bi-directional causality support the Cukierman and Meltzer model. Using the standard TGARCH models, the presence of asymmetry is found in the conditional variance of annualized inflation, and finally the evidence of long memory exists in the conditional variance of annualized inflation.

Keywords: Inflation; Inflation Uncertainty; AR-GARCH Models; Asymmetry; Long Memory; Islamic Republic of Iran

JEL Classification: E31; C22; C51; O53

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1 Introduction

The inflation–uncertainty[1] relationship is frequently argued by economists since 19970s [see, for example, Okun (1971), Friedman (1977), Hafer and Hafer (1981), Cukierman and Meltzer (1986), Davis and Kanago (1998), Fountas, el al. (2000), Stuber (2001), Kontonikas (2004), and Thornton (2007)]. Because of the welfare cost of inflation, inflation is extremely unpopular with the public. Inflation has indirect real cost through its effects on uncertainty that lowers welfare. Particularly, in the developing countries such as Iran where there is no perfect indexation, the deadweight loss from the inflation tax is too large. Moreover, the linkages between inflation and inflation uncertainty have real effects on the aggregate economic activity through investment, employment, financial markets and then output levels which again lower welfare of the society.

The central focus of theoretical and empirical studies is whether a rise in the level of inflation raises uncertainty about future inflation. The idea behind this relationship is that high inflation creates uncertainty about future monetary policy and makes monetary policy less stable [Ball and Cecchetti (1999)]. When the economy experiences high inflation, the central bank would like to disinflate. But, it fears the recession that would probably result. It is likely that disinflation will occur eventually, but the time is uncertain. Since the public does not know the tastes of future policymakers. Fischer and Modigliani (1978) point out those governments typically announce unrealistic stabilization programs as the inflation rate rises, thus increasing uncertainty about what actual path of prices will be. Friedman (1977) argues that burst of inflation produces strong pressure to counter it. Policy goes from one direction to the other, encouraging wide variation in inflation, which increases uncertainty and lowers output growth and welfare.

Friedman, in his Nobel address, points out that the causation runs from inflation to uncertainty while Cukierman and Meltzer (1986) argue that increases in inflation raise the optimal average inflation rate by increasing the incentive for policymaker to create inflation surprises. In contrast to the Friedman view that high inflation creates uncertainty, the causation in Cukierman and Meltzer is from increased uncertainty to higher average inflation.

This paper investigates the relationship between inflation and inflation uncertainty in Iran by using Monthly data over the period 1959:03 – 2008:02. The use of autoregressive conditional heteroscedasticity (ARCH) and generalized ARCH (GARCH) models allow us to generate a time-varying conditional variance of surprise inflation as a standard measure of inflation uncertainty. Granger methods are employed to examine bi-directional causality between inflation and inflation uncertainty. Threshold GARCH (TGARCH) models are considered to investigate asymmetry in the conditional variance of inflation. The Component GARCH (CGARCH) models are employed to decompose inflation uncertainty into a short-run and a long-run component by permitting transitory deviations of the conditional

[1] Uncertainty refers to situations in which the probability of future events cannot be determined, while in the case of a risky event an explicit probability can be assigned. Future volatility in an economic variable is the sum of both predictable and unpredictable components. In fact the uncertainty of an economic variable can be more precisely defined as its unpredictable volatility [Grier and Perry, (1998)].

volatility around a time-varying trend. The model examines the presence of long memory in the conditional variance of inflation.

The paper proceeds as follows: In section two, methodology is briefly introduced. Section three empirically investigates the relationship between inflation and its variability, asymmetry, and long memory. Section four presents policy implications and the results.

2 Methodology: ARCH Models and Their Extensions

There are opposite views and contradictory evidence about the relationship between the inflation rate and the variance of inflation. Friedman (1977) argues that higher inflation variances is associated with higher inflation rates. Engle (1982) applies an autoregressive conditional heteroscedastic (ARCH) model, and fails to confirm a positive relationship between the conditional mean and variance of inflation for the United States. Bollerslev (1986) extends the ARCH (q) model, and suggests that the conditional variance can follow an ARMA process. The Generalized ARCH model, called GARCH (p, q), contains both autoregressive and moving average components. Caporale and McKiernan (1997) employ a GARCH model and find a positive and significant relationship between the level and variability of inflation for the annualized US inflation rate. The Grier and Perry (1998) findings overwhelmingly confirm the theoretical predictions in G7 countries.

In this section, ARCH models and their extensions such as GARCH, Threshold GARCH (TGARCH), and Component GARCH (CGARCH) models are discussed to analyse the relationship between inflation and its variability. The functional form of an ARCH (q) process is formally given by

$$\varepsilon_t | \Psi_{t-1} \sim N(0, h_t^2) \tag{1}$$

$$h_t^2 = \alpha_0 + \sum_{j=1}^{q} \alpha_j \varepsilon_{t-j}^2 \tag{2}$$

where ε_t is the innovation in the ARMA model for a stationary series y_t; and h_t^2 is the conditional variance of ε_t with respect to the information set, Ψ_{t-1}. Since h_t^2 is strictly positive for all realisations of ε_t, $\alpha_0 > 0$ and $\alpha_j \geq 0$ for j = 1, 2, ..., q. Bollerslev (1986) extends ARCH (q) process to GARCH (p, q) process, defined as:

$$h_t^2 = \alpha_0 + \sum_{j=1}^{q} \alpha_j \varepsilon_{t-j}^2 + \sum_{j=1}^{p} \delta_j h_{t-j}^2 \tag{3}$$

where $p \geq 0$; $q > 0$; and $\delta_j \geq 0$, for j = 1, 2, ..., p

To test for the presence of ARCH effects, the best AR (p) model for y_t is first identified and estimated, and then the squares of the residuals, e_t^2, are obtained. In the next stage, the following equation is estimated:

$$e_t^2 = \alpha_0 + \sum_{j=1}^{q} \alpha_j e_{t-j}^2 + u_t \tag{4}$$

Under the assumption of normality, the test statistic is

$$\lambda_{LM} = T.R^2 \tag{5}$$

where T is the sample size; and R^2 is the coefficient of determination obtained from (4), the test is asymptotically distributed as χ_q^2. In the empirical analysis, the GARCH model is extended by including the lagged level of the inflation rate, y_{t-1}, as follows:

$$h_t^2 = \alpha_0 + \sum_{j=1}^{q} \alpha_j \varepsilon_{t-j}^2 + \sum_{j=1}^{p} \delta_j h_{t-j}^2 + \gamma y_{t-1} \tag{6}$$

In this model, y_{t-1} is assumed to influence the conditional error variance in addition to the past squared errors.[2] This enables us to test whether the inflation rate affects its variability.

GARCH model implies that the positive and negative residuals have a symmetric impact on the conditional variance. Glosten et al. (1993) introduce a TGARCH model to allow for negative residuals to have different impact on the conditional variance than do positive residuals. The specification of the model for conditional variance is:

$$h_t^2 = \alpha_0 + \sum_{j=1}^{q} \alpha_j \varepsilon_{t-j}^2 + \xi \varepsilon_{t-1}^2 d_{t-1} + \sum_{j=1}^{p} \delta_j h_{t-j}^2 \tag{7}$$

where d_{t-1} = 1 if $\varepsilon_{t-1} \prec 0$ and = 0 otherwise. The coefficient of $\xi \neq 0$ implies asymmetry in the conditional variance. In this model good news has an impact of α_1, while bad news has an impact of $\alpha_1 + \xi$.

GARCH model is also extended to allow the mean reversion level of the conditional variance to itself be time varying.[3] In this model, the conditional variance is divided into permanent and transitory components. The model is formally written as follows:

$$h_t^2 = q_t + \alpha_1(\varepsilon_{t-1}^2 - q_{t-1}) + \delta_1(h_{t-1}^2 - q_{t-1}) \tag{8}$$

$$q_t = \alpha_0 + \rho q_{t-1} + \mu(\varepsilon_{t-1}^2 - h_{t-1}^2) \tag{9}$$

Equation (8) describes the transitory component, while equation (9) describes the long run component. If ρ in equation (9) is equal to one, then the conditional

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[2] The standard approach is to restrict y_t to contain only past level of inflation. In this case, estimated positive and significant coefficient of γ is consistent with the Freidman-Ball link.

[3] The conditional variance displays long-run mean reversion to a constant level for all time given by α_0 in equation (3).

variance contains a unit root. If $\rho \prec 1$ and $\rho \succ \alpha_1 + \delta_1$, then q_t is the long memory component of the conditional variance.

Testing any of the theories, discussed in the paper, requires the construction of a measure for uncertainty. The empirical studies used the differences in the standard deviation of inflation across countries as a measure of the differences in inflation uncertainty across countries. In the time series studies, two other uncertainty measures are considered. They comprise the cross-sectional dispersion of individual forecasts from surveys and moving standard deviation of the variable under consideration [see, for example, Holland (1993), Golob (1993), and Hess (1993)]. Kevin et al. (1998) point out that there is large difference between variability and uncertainty, depending on whether the variability is predictable in the model under consideration. Predictable fluctuations in a variable will show up in standard deviation measures although they create no true economic uncertainty. Cukierman and Meltzer (1986) introduce the variance of the stochastic or unpredictable component of a variable as a measure of uncertainty.

GARCH techniques estimate a model of the variance of unpredictable innovations in a variable rather than simply calculating a variability measure from past outcomes like moving standard deviation. That is, a GARCH model estimates a time-varying residual variance that corresponds well to the notation of uncertainty in Cukierman and Meltzer. In this paper, A GARCH model is used to generate a time varying conditional variance of surprise inflation as a measure of inflation uncertainty.

3 Estimates

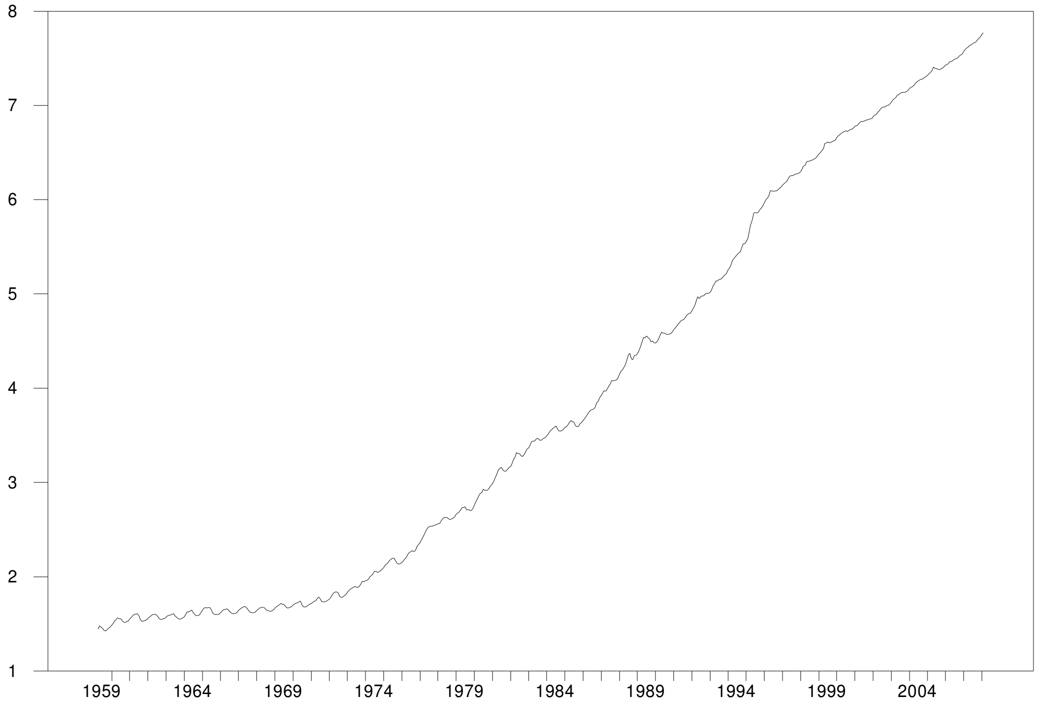
The monthly consumer price index (CPI) is employed over the period 1959:03 – 2008:02 published by the central bank of Iran. Figure 1 plots the log of CPI using monthly data. From the plots of the figure, it appears that the sample period may be split into two inflation regimes as follows:

- 1959:03 – 1972:07: relatively low and stable inflation
- 1972:08-2008:02: higher and more variable inflation

It should be pointed out that the annual average inflation rates were in single figures from 1960 to 1972. After 1972, with the oil price and the quantity of oil exports increasing, the rates of inflation rose sharply. Consequently, over the period, inflation rates exhibited large fluctuations.

Over the whole period, inflation trend exhibits phases of relatively tranquillity followed by the periods of high volatility. During the first inflationary regime, both the price level and its unconditional volatility have been lower while over the second inflationary regime, it appears that periods of higher average inflation correspond to periods of more volatile. This type of behaviour is called conditional heteroskedastic, since there are periods in which the variance is relatively high. Therefore, this stylized fact provides an interesting application of GARCH modeling in this study.

Figure 1 The Plot of the log of CPI, P,: Monthly Data (1959:03 – 2008:02)



3.1 Unit Root Tests

The Iranian economy has experienced some big shocks and major government interventions over recent decades. Possible structural breaks include first oil price shock in 1972, the revolution in 1978, the second oil price shock in 1979, the eight–year war from 1980 to 1988, the third oil shock in 1986, the economic reform programme during the period 1989 – 1993, and the fourth oil price shock started since 2001. The structural breaks in the CPI are examined. The evidence shows that there is a break in the slope of the log of CPI, P, after 1972:08 (see Figure 2).

The sample autocorrelation function (ACF) and partial autocorrelation function (PACF) are used to identify the behaviour of the variable.[4] The ACF for the level of the series does not tail off quickly, suggesting nonstationary behaviour. Similar behaviour is observed if the ACF for the two sub–sample periods is examined (results are not shown).[5] The series of $(1 - L)\, P_t$ has high and persistent autocorrelations at lags around multiples of 12 (see Figure 2) and the seasonal peaks decay slowly. They seem to be dying out very slowly, suggesting the need for seasonal differencing. Moreover, the ACF for $(1 - L^{12})\, P_t$ dies out very slowly, suggesting the need for further differencing. When the series $(1 - L)(1 - L^{12})\, P_t$ is considered, the pattern of the ACF suggests that this series is stationary.

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[4] The ACF and PACF will be used to build the best linear models later.

[5] The estimation process has been done by using Regression Analysis of Time Series (RATS), Eviews, and Microfit statistical packages.

As mentioned before, due to the breakpoint in the time series, the Perron procedure is applied to test for a unit root in the inflation measures.[6] The number of lags in the equations of unit root tests is determined by starting with some upper bound on k suggested by Campbell and Perron (1991).

Figure 2 Correlogram for ΔP_t: Monthly Data (1959:4 – 2008:02)

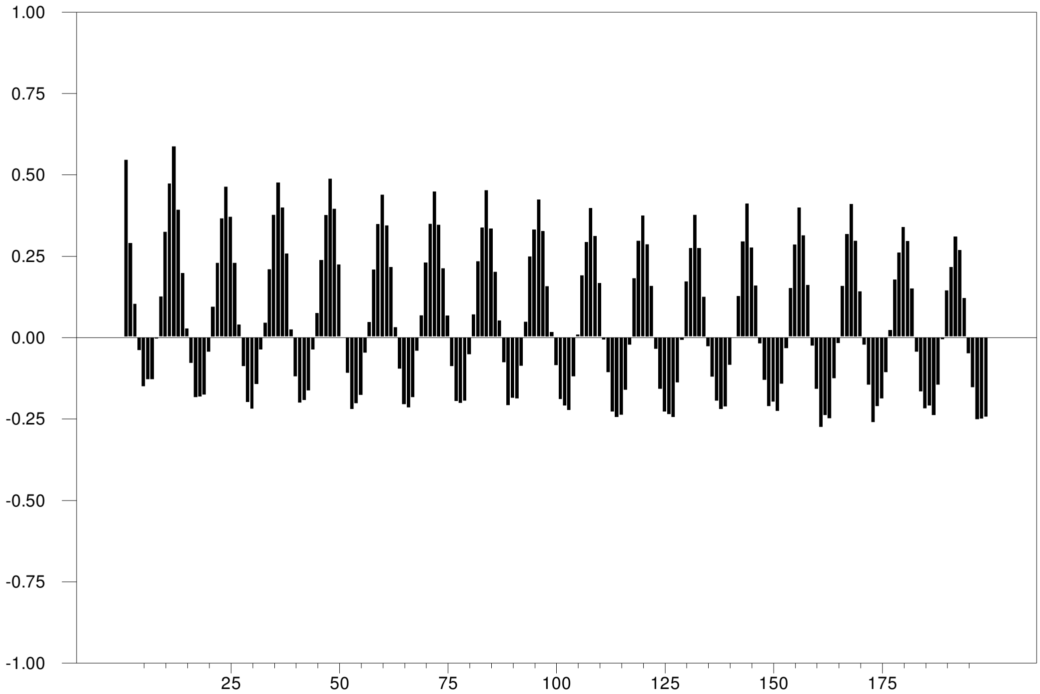

Table 1 presents the results of the univariate Perron test of unit roots for P_t, $(1 - L)\ P_t$, $(1 - L^{12})\ P_t$, and $(1 - L)(1 - L^{12})\ P_t$. Model (C) is used for the level while model (A) is used for the differenced series. To evaluate the significance of the t–statistic for the null hypothesis $\gamma = 1$, the critical values in Perron's Tables IV.A and VI.C are used. The null hypothesis of a unit root for P_t cannot be rejected. Although the test suggests that the first difference of P_t is stationary at one percent level, the ACF shows a strong seasonal pattern in this variable (see Figure 2). To eliminate seasonality, $(1 - L^{12})$ Pt is considered, which represents the annualized inflation rate. The result indicates a unit root in this variable. Finally, for the series $(1 - L)(1 - L^{12})\ P_t$, the null hypothesis of a unit root is rejected so that this series is stationary.

[6] Some other formal unit root tests such as Zivot and Andrews (1992) were used in this study. Here only the results of the Perron tests are reported.

Table 1 Univariate Perron Test for Unit Roots Using Monthly Data

Regressions: Model (A) $Y_t = \hat{\mu} + \hat{\alpha}T + \hat{\beta}D_t + \hat{d}TB_t + \hat{\gamma}Y_{t-1} + \sum_{i=1}^{k}\hat{c}_i\Delta Y_{t-i} + \hat{\varepsilon}_t$ Model (C) $Y_t = \hat{\mu} + \hat{\alpha}T + \hat{\beta}D_t + \hat{d}TB_t + \hat{\delta}DT_t + \hat{\gamma}Y_{t-1} + \sum_{i=1}^{k}\hat{c}_i\Delta Y_{t-i} + \hat{\varepsilon}_t$				
	P_t	ΔP_t	$\Delta_{12}P_t$	$\Delta\Delta_{12}P_t$
n	539	537	515	414
k	48	49	60	60
$t_{\hat{\gamma}-1}$	-3.20	-13.67* **	2.49	-12,86***

Notes:

- P_t is the log of the CPI; and Δ is the difference operator.
- D_t, TB_t and DT_t are dummy variables taking values as follows: $D_t = 1$ if $t \geq t^*$ and 0 otherwise; $TB_t = 1$ if $t = t^*$ and 0 otherwise; $DT_t = t$ if $t \geq t^*$ and 0 otherwise, where $t^* = 1972{:}08$. The ratio of pre–break sample size to total sample size is $\lambda = 0.27$.
- n is the number of observations used for estimation; and k is the number of lags.
- *** indicate statistical significance at the 1% level according to the critical values of the Perron test (1989, Table IV.A and Table VI.C).

3.2 AR and GARCH Models

An integrated I (d) series may be represented by an autoregressive integrated moving average or ARIMA (p, d, q) process as follows:

$$\phi_p(L)(1-L)^d Y_t = \theta_0 + \theta_q(L)\varepsilon_t \tag{10}$$

where $\phi_p(L) = 1 - \phi_1 L - \cdots - \phi_p L^p$; $\theta_q(L) = 1 - \theta_1 L - \cdots - \theta_q L^q$; and ε_t is zero–mean white noise. For the differenced series $\Delta^d Y_t$ to be stationary, the roots of $\phi(L)$ must be outside the unit circle. The roots of $\theta(L)$ also should be outside the unit circle for the invertibility condition to hold.

To identify the presence of ARCH effects in the inflation measures, the model building approach of Box and Jenkins (1976) is employed. The best fitting AR models of inflation are identified and estimated for monthly data over the 1959:03 – 2008:02 periods. [7]

The unit root tests for the monthly CPI show that the series $\Delta\Delta_{12}P$ is stationary so that d = 1. The differenced annualized inflation rate ($\Delta\Delta_{12}P_t = y_t$) is considered.

Table 2 shows the estimated AR (60) model for the whole period and the second sub–period, post-breakpoint, and AR (24) for the first sub–period, pre-breakpoint.

[7] In the presence of seasonality, the Bell and Hillmer's (1984) survey shows that it is wise to avoid using seasonal adjusted data, since the seasonal and autoregressive moving average (ARMA) coefficients are best identified and estimated jointly.

Table 2 AR (p) Models for the Differenced Annualized Inflation Rate ($\Delta\Delta_{12}P_t = y_t$)

	Model1 (1959:03 –2008:02)	Model2 (1959:03 –1972:07)	Model3 (1972:08 –2008:02)
$\hat{\beta}_0$	0.5E-3	0.3E-3	0.6E-3
	(0.95)	(0.41)	(0.94)
$\hat{\beta}_1$	0.31	-	0.32
	(7.40)		(7.05)
$\hat{\beta}_{12}$	-0.64	-0.60	-0.63
	(-14.68)	(-6.91)	(-13.10)
$\hat{\beta}_{13}$	0.18	-	0.17
	(3.49)		(3.15)
$\hat{\beta}_{24}$	-0.56	-0.21	-0.57
	(-11.03)	(-2.38)	(-10.12)
$\hat{\beta}_{25}$	0.15	-	0.15
	(2.77)		(2.56)
$\hat{\beta}_{36}$	-0.42	-	-0.43
	(-7.92)		(-7.28)
$\hat{\beta}_{37}$	0.13	-	0.14
	(2.62)		(2.57)
$\hat{\beta}_{48}$	-0.24	-	-0.25
	(-4.83)		(-4.59)
$\hat{\beta}_{49}$	0.10	-	0.11
	(2.41)		(2.33)
$\hat{\beta}_{60}$	-0.16	-	-0.16
	(-3.85)		(-3.57)
n	515	124	427
$\bar{R}^2$	0.366	0.277	0.368
s	0.0124	0.0080	0.0132
Q(24)	32.96 [0.105]	26.99 [0.305]	40.71 [0.284]
χ^2_{ARCH} (4)	31.95 [0.000]	0.42 [0.788]	22.74 [0.000]
χ^2_{ARCH} (12)	49.57 [0.000]	15.48 [0.218]	36.53 [0.000]
χ^2_{ARCH} (24)	66.91 [0.000]	18.66 [0.770]	53.34 [0.001]

Notes:

- y_t is the differenced annualized CPI inflation rate.
- n is the number of observations used for estimation; s is the standard error of estimate; Q(k) is the Ljung–Box statistic for residual autocorrelation up to order k; and χ^2_{ARCH} (k) is the Engle (1982) test for ARCH up to order k.
- the numbers in brackets under the coefficients are t–values; and the numbers in square brackets are P–values.

The Engle test confirms that there are ARCH effects in the whole sample, 1959:03 – 2008:02, as well as the second sub–period, 1972:08 – 2008:02. However, for the first sub–period, the test confirms that there is no ARCH.

Table 3 AR(60)-GARCH (1, 1) Models for the Differenced Annualized Inflation Rate

	Model1 (1959:03 –2008:02)	Model2 (1972:08 –2008:02)
$\hat{\beta}_0$	-0.9E-4	0.2E-3
	(-0.19)	(0.39)
$\hat{\beta}_1$	0.29	0.32
	(7.33)	(7.45)
$\hat{\beta}_{12}$	-0.68	-0.64
	(-18.61)	(-15.37)
$\hat{\beta}_{13}$	0.19	0.17
	(4.16)	(3.42)
$\hat{\beta}_{24}$	-0.53	-0.52
	(-12.15)	(-10.82)
$\hat{\beta}_{25}$	0.13	0.14
	(3.17)	(3.00)
$\hat{\beta}_{36}$	-0.37	-0.38
	(-7.69)	(-7.05)
$\hat{\beta}_{37}$	0.09	0.11
	(2.12)	(2.12)
$\hat{\beta}_{48}$	-0.23	-0.26
	(-5.79)	(-5.56)
$\hat{\beta}_{49}$	0.10	0.12
	(3.05)	(3.24)
$\hat{\beta}_{60}$	-0.15	-0.17
	(-4.59)	(-4.72)
$\hat{\alpha}_0$	0.4E-5	0.7E-5
	(3.43)	(2.58)
$\hat{\alpha}_1$	0.06	0.07
	(3.07)	(2.62)
$\hat{\delta}_1$	0.91	0.89
	(35.90)	(23.11)
$\hat{\gamma}$	0.001	0.001
	(5.73)	(5.43)
n	515	427
$\overline{R}^2$	0.353	0.357
s	0.0126	0.0133

Notes:

- n is the number of observations used for estimation; and s is the standard error of estimate.
- the numbers in brackets under the coefficients are the t–values.

In the next step, the aim is to determine appropriate values of p and q in GARCH (p, q) process and estimate the resulting GARCH models. For both, the whole period and the second sub–period, the GARCH (1, 1) model is preferred among several alternatives considered. The estimates are shown in Table 3. Comparing the results of Table 3 with the estimates in Table 2, it can be seen that the presence of ARCH does not affect the OLS estimates of the linear AR model. Furthermore, the estimates of the GARCH parameters satisfy the stationarity condition $\alpha_1 + \delta_1 \prec 1$. As can be seen from the table, the sum of the ARCH and GARCH coefficient $\alpha_1 + \delta_1$ is very close to one, indicating that inflation volatility shocks are quite persistent.

The estimate of γ (the coefficient of the lagged level of differenced annualized CPI inflation rate, y_{t-1}) is positive and significant, suggesting that there is a positive relationship between the differenced annualized inflation rate and its variance in both models supporting the Friedman hypothesis. So, accommodating leads not only to high level of inflation, but also to costly uncertainty.

3.3 Direction of Causality between Inflation and Inflation Uncertainty

Since the cost of inflation and inflation uncertainty on economic growth and welfare are significant, it is beneficial to determine the direction of the causality between inflation and uncertainty. Friedman claims that the causation runs from inflation to uncertainty. Cukierman and Meltzer (1986) argue that increases in inflation raise the optimal average inflation rate by increasing the incentive for policymaker to create inflation surprises. In contrast to the Friedman view that high inflation creates uncertainty, the causation in Cukierman and Meltzer is from increased uncertainty to higher average inflation.

In this sub-section, bi-directional relationship between inflation and inflation uncertainty is investigated to test the predictions of economic theory. The GARCH (1, 1) model, as identified and estimated in the previous sub-section, is used to provide some statistical evidence nature of the relationship between inflation and inflation uncertainty.

Table **4** and Table 5 employ the GARCH (1, 1) measure of inflation uncertainty. The results show the null hypothesis that inflation does not Granger-cause inflation uncertainty is rejected in both periods using 12, 24 and 36 lags. Furthermore, the null hypothesis that uncertainty does not Grange-cause inflation is rejected in both periods using again 12, 24 and 36 lags. Consequently, bi-directional causality is obtained for both periods, whole and second sub sample periods.

To evaluate the positive or negative bi-directional causality of relationships, the sign of coefficients are considered. The evidence shows that the sum of the coefficients on lagged uncertainty in the inflation equation is positive. Moreover, the sum of the coefficients on lagged inflation in the uncertainty equation is also positive. So, inflation uncertainty raises inflation and vice versa supporting Friedman as well as Cukierman and Meltzer hypotheses. The results confirm that increased inflation raises uncertainty and increased uncertainty raises inflation and then both create real welfare losses.

Table 4 Granger Causality Tests for Inflation and Inflation Uncertainty for the Differenced Annualized Inflation Rate over the Period 1959:03 - 2008:02

	H_0: Inflation does not Granger - Cause Inflation Uncertainty	H_0: Inflation uncertainty does not Granger - Cause Inflation
Twelve Lags	3.25 (0.000)	1.82 (0.042)
Twenty Four Lags	1.84 (0.009)	2.43 (0.000)
Thirty Six Lags	2.05 (0.000)	1.99 (0.000)

Note:

- GARCH (1, 1) is used to generate the measure of uncertainty.

Table 5 Granger Causality Tests for Inflation and Inflation Uncertainty for the Differenced Annualized Inflation Rate over the Period 1972:08-2008:02

	H_0: Inflation does not Granger-Cause Inflation Uncertainty	H_0: Inflation uncertainty does not Granger -Cause Inflation
Twelve Lags	3.25 (0.000)	1.95 (0.028)
Twenty Four Lags	1.53 (0.055)	1.99 (0.004)
Thirty Six Lags	1.77 (0.006)	1.76 (0.006)

Note:

- GARCH (1, 1) is used to generate the measure of uncertainty.

In sum, the findings of the study suggest the existence of bi-directional positive causality between inflation and inflation uncertainty in Iran. It means an increase in inflation raises uncertainty and also an increase in inflation uncertainty is positively associated with future inflation. This generates real welfare costs through higher inflation and higher uncertainty.

3.4 Asymmetry and Long Memory

First the threshold GARCH (TGARCH) model is considered to investigate asymmetry in the conditional variance of inflation for the whole period as well as the second sub-period. The results are presented in equation (11) and (12) below:[8]

[8] The coefficients for the AR models of inflation are not reported in order to save space.

whole period:

$$h_t^2 = (0.2E-5) + 0.07\varepsilon_{t-1}^2 - 0.09\varepsilon_{t-1}^2 d_{t-1} + 0.96h_{t-1}^2 \tag{11}$$

(3.83) (5.75) (-5.38) (95.52)

second period:

$$h_t^2 = (0.1E-4) + 0.18\varepsilon_{t-1}^2 - 0.17\varepsilon_{t-1}^2 d_{t-1} + 0.82h_{t-1}^2 \tag{12}$$

(2.50) (2.94) (-3.02) (12.91)

Since the estimated coefficient of α_2 is significant in both models, there is evidence of asymmetry in the conditional variance in both periods. Therefore, the positive and the negative residuals have asymmetric impact on the conditional variance. The negative sign of asymmetry parameter suggests that "good news" on inflation result in a smaller increase in inflation uncertainty than "bad news". Furthermore, the exponential GARCH (EGARCH) model proposed by Nelson (1991) is used to test asymmetry impacts. The evidence of asymmetry is again found in the conditional variance in both periods.

In the next step, the transitory and permanent components of the conditional variance are investigated. The Component GARCH (CGARCH) model is employed to decompose inflation uncertainty into a short-run and a long-run component by permitting transitory deviations of the conditional volatility around a time-varying trend, q_t. Equations (13) through (14) show the transitory and permanent components of the conditional variance in both periods. Since $\hat{\rho} \prec 1$ in equations (14) and (16), then the conditional variance does not contain a unit root and also $\hat{\rho} \succ \hat{\alpha}_1 + \hat{\delta}_1$, then q_t is the long memory component of the conditional variance of inflation. Therefore, the results confirm the presence of long memory in the conditional variance in both periods [see also equations (15) and (16)]. These findings imply that long-run mean reversion of inflation's conditional variance does not occur very slowly.

whole period:

$$h_t^2 = q_t + 0.22(\varepsilon_{t-1}^2 - q_{t-1}) + 0.28(h_{t-1}^2 - q_{t-1}) \tag{13}$$

(3.34) (1.40)

$$q_t = (0.1E-2) + 0.99q_{t-1} + .05(\varepsilon_{t-1}^2 - h_{t-1}^2) \tag{14}$$

(2.15) (135.66) (2.34)

second period:

$$h_t^2 = q_t - 0.01(\varepsilon_{t-1}^2 - q_{t-1}) - 0.97(h_{t-1}^2 - q_{t-1}) \tag{15}$$

(-0.99) (-33.18)

$$q_t = (0.1E-2) + 0.90q_{t-1} + 0.18(\varepsilon_{t-1}^2 - h_{t-1}^2) \tag{16}$$

(4.77) (21.22) (3.16)

However, it is worth to point out that the GARCH (1, 1) and Component GARCH models are considered to test between a GARCH (1, 1) and GARCH (2, 2). The results of the Likelihood Ratio (LR) tests reject the null hypothesis in favour of GRACH (1, 1) in both periods.

4 Policy Implication and Conclusion

This paper analyses the relationship between inflation and uncertainty in the context of the Iranian economy using monthly data over the period 1959:03 – 2008:02. The properties of inflation rate were investigated. The evidence showed that the monthly changes (ΔP) are stationary while the annualized changes ($\Delta_{12}P$) are not. However, considering the strong seasonal pattern in ΔP, the first difference of annualized inflation rate is stationary. Due to the existence of some major internal and external shocks in the Iranian economy, the structural breaks were examined. The findings showed that there is a break in the slope of the log of CPI, P, after 1972:08

The best fitting AR models were used to investigate the presence of ARCH effects in the inflation measures. The Engle tests provide the evidence of time-varying variances of inflation measure over the whole period as well as the second sub-sample period. The GARCH (1, 1) model was preferred among several alternatives considered for inflation measure. The results suggested that there is a positive relationship between inflation and its variability. The evidence indicated that increased inflation raises inflation uncertainty confirming the theoretical predictions made by Friedman. Furthermore, the findings of bi-directional causality between inflation and inflation uncertainty support the Friedman, and Cukierman and Meltzer hypotheses.

To evaluate the positive or negative bi-directional causality of relationships, the sign of coefficients are considered. The evidence shows that the sum of the coefficients on lagged uncertainty in the inflation equation is positive. Moreover, the sum of the coefficients on lagged inflation in the uncertainty equation is also positive. Positive bi-directional causality between inflation and inflation uncertainty affects the performance of the economy in some fundamental aspects. This decreases the effectiveness of policies, lowers the level of the aggregate output through investment and employment, and reduces welfare of the society [see, for example, Barro (1997)].

Using the standard TGARCH model, the presence of asymmetry effects and long memory is investigated in the conditional variance of annualized inflation. The evidence of asymmetry is found in the conditional variance of annualized inflation. The negative sign of asymmetry parameter suggest that "good news" on inflation result in a smaller increase in inflation uncertainty than "bad news". Moreover, the results confirm the presence of long memory in the conditional variance in both periods.

Since the rate of inflation increases uncertainty, this implies substantial costs of inflation in the economy through various channels. Firstly, when inflation is relatively high, this generates real welfare costs through higher inflation and higher uncertainty. Moreover, this affects the redistribution of wealth in the society. Consequently, greater inflation variability increases uncertainty and lowers welfare. Secondly, inflation uncertainty and variability have large effects on incentives for investment and saving in the economy. As a result, the growth rate of the economy declines. Therefore, one of the major policy implications for the economy is to aim at low average inflation rates in order to reduce the negative consequences of uncertainty. Since the findings indicate that inflation uncertainty increases with the level of inflation, the costs of inflation uncertainty might be minimized by pursuing a policy of price stability.

The presence of uncertainty in the economy is important for the effectiveness of economic policies and the decisions of agents. Uncertainty implies that the policymaker cannot guarantee that his target value is attainable, since the target is affected by other factors in addition to policy actions. There is a negative relationship between uncertainty and aggregate economic activity. Consequently, a reduction in inflation uncertainty could raise the level of real GDP. Therefore, the authorities should reduce welfare costs by reducing inflation and inflation uncertainty through monetary and exchange policy which are the main determinants of inflation in Iran [see, for example, Moradi (2002)].

Due to the close link between fiscal and monetary policy through monetization of the budget deficit in Iran, the authorities have adopted a decidedly expansionary monetary policy over recent decades. Although the government was able to earn revenue through seigniorage by accepting a higher rate of inflation [see, Moradi (2001)], higher inflation resulted in higher uncertainty in the economy and affected the effectiveness of government policies (more specifically monetary and exchange rate policies). To boost the economy, the policymakers have to take into account the uncertainty surrounding the transmission mechanisms of monetary and exchange rate policy in the economy. The policymakers should avoid an expansionary monetary policy and a depreciation of the domestic currency in order to reduce inflation. Consequently, this will reduce uncertainty.

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