Threshold GARCH modeling of the inflation inflation uncertainty relationship: historical evidence from the Turkish economy

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THRESHOLD GARCH MODELING OF THE INFLATION & INFLATION UNCERTAINTY RELATIONSHIP: HISTORICAL EVIDENCE FROM THE TURKISH ECONOMY

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ABSTRACT

In this paper, the preceding / causal relationships between inflation and inflation uncertainty have been tried to be examined for the Turkish economy. Dealing with the information content of this relationship, we estimate that positive inflationary shocks are associated with statistically significant and quantitatively larger levels of inflation uncertainty than are negative shocks. Our estimation results indicate that inflation in fact leads to inflation uncertainty in line with the Friedman-Ball hypotheses. However, our findings contradict the Cukierman-Meltzer hypotheses that inflation uncertainty leads to inflation in a positive way. We find that the larger the inflation uncertainty the lower would likely to be the level of inflation.

Key Words: Inflation; Inflation Uncertainty; Threshold GARCH Modeling; Granger Causality Analysis; Turkish Economy;

JEL Classification: C32; C51; E31;

ÖZET

Bu çalışmada, enflasyon ve enflasyon belirsizliği arasındaki önceleme / nedensellik ilişkilerinin Türkiye ekonomisi için incelenmesine çalışılmıştır. Bu ilişkinin bilgi içeriğiyle ilgili olarak, pozitif enflasyonist şokların negative şoklara göre istatistiksel olarak anlamlı ve miktar olarak daha fazla bir şekilde enflasyon belirsizliği ile özdeşleştiği tahmin edilmektedir. Tahmin sonuçlarımız enflasyonun Friedman-Ball varsayımları doğrultusunda gerçekten de

enflasyon belirsizliğine yol açtığını göstermektedir. Bununla birlikte, bulgularımız enflasyon belirsizliğinin pozitif bir şekilde enflasyona yol açtığı şeklindeki Cukierman-Meltzer varsayımları ile çelişmektedir. Daha fazla enflasyon belirsizliğinin enflasyon düzeyindeki olası bir azalma ile birlikte gerçekleştiği bulgusuna ulaşılmıştır.

Anahtar Kelimeler: Enflasyon; Enflasyon Belirsizliği; Eşik GARCH Modellemesi; Graner Nedensellik Çözümlemesi; Türkiye Ekonomisi;

JEL Sınıflaması: C32; C51; E31;

1.INTRODUCTION

The role of inflationary framework on the design and implementation of monetary policies has a considerable impact on the discussions of economic policies. The uncertainty components relating to the inflation are able to affect expectations dealing with the decision making process of economic agents and reflect to the economic behaviors of people shaped by insights as to what will happen in the future course of the aggregate economic activity. Such a phenomenon arising from relationships between inflation and its associated uncertainty has therefore been required to be elaborately examined by the researchers and policy makers, and inferences extracted from these issues of interests would likely to lead us to obtain a crucial knowledge of various other properties of inflation. As emphasized by Okun (1971), linking inflation and inflation uncertainty in a positive manner, this relationship has been of a special importance to be able to reveal the extent of the information content of the inflation uncertainty and to estimate the direction of the causal relationships between inflation and its uncertainty. Friedman (1977) in his Nobel Lecture states that high inflation rates would not likely to be steady especially during the transition decades, and the higher the inflation the more variable it is likely to be since it distorts relative prices and financial contracts which have been adjusted to a long-term "normal" price level. All these would be resulted in additional uncertainties in the economy that lead economic agents to be curious about how long it will take that policy makers try to disinflate the economy, and therefore, to bear the costs of disinflation readily. Given the presence of these types of doubts in the economy, the volume of investment and the aggregate output growth would naturally tend to be negatively influenced, that is to say, the larger the volatility of inflation and the greater the uncertainties about when and how policy authorities decide to intervene for price stability purposes, the lower would be the real income growth capacity of the economy. According to Friedman, a possible outcome of such a process would be the increasing unemployment level as well as the political unrest leading the society to be polarized.

Supporting the arguments put forward by Friedman (1977), Ball and Cecchetti (1990) find that inflation has really significant effects on its uncertainty at long horizons which lead to substantial costs due to the increased risks for individuals who have nominal contracts between themselves and these effects would be resulted in variation in policy behavior reacting to inflation, in turn, destabilizing output growth. Furthermore, Ball (1992) employing an asymmetric game perspective among monetary authorities and the decision making process of the economic agents formalizes the view of Friedman in the sense that low levels of inflation would be coincided with the policy behavior of monetary authorities to keep inflation at these levels that give rise to low inflation uncertainty, as well. However, the public would be more uncertain the higher the level and variability of inflation, as to when policy makers decide to implementing a stabilization policy to fight inflation. In this case, an information problem stemmed from activating policy would also be the length of the time lags to that the stabilization policies have been subject in achieving policy consequences consistent with *a priori* expectations of the policy makers.

On the other side, Cukierman and Meltzer (1986) and Cukierman (1992) follow the approach proposed by Barro and Gordon (1983) and try to examine the causal relationship between inflation and inflation uncertainty in an opposite way. Since governments have different objectives determined stochastically over time that lead to a trade-off between expanding output by making monetary surprises and keeping inflation at low levels, choices of policy makers in favor of creating monetary surprises to stimulate economic growth would likely to be resulted in higher money growth rates and inflation than the expectations of economic agents conditioned upon past realizations in line with some form of adaptive expectations. Following Fountas et al. (2002), of course, it is possible to assume in a different way that if increasing uncertainty has been perceived by the policy makers so much detrimental resulted in real costs, inflation uncertainty can in this case direct policy makers to applying to a tight monetary policy to lower average inflation so that they are more likely to achieve their commitment to long-run price stability.

Many papers in the contemporaneous economics literature try to examine the relationships and the direction of causality between inflation and inflation uncertainty and yield in general supportive evidence to the Friedman-Ball approach for various country cases. Among many others, Holland (1995) using post-war US data estimates that an increase in inflation precedes an increase in its uncertainty resulted in some welfare cost for the whole society. Grier and Perry (1998) employing data from the G7-countries find that inflation in general tends to raise its uncertainty, however, some mixed results are obtained for a reverse causal relationship in the sense that increased inflation uncertainty lowers inflation in the US, UK and Germany and raises inflation in Japan and France. Estimation results in Fountas (2001) and Kontonikas (2004) using the UK data demonstrate that inflationary periods are in fact associated with larger inflation uncertainty. Daal et al. (2005) using data from various country cases inclusive of both developed and emerging market economies find that positive inflationary shocks strongly affect inflation uncertainty in a positive manner, but the effects of inflation uncertainty on inflation seem to be varying and are highly sensitive to the country cases considered. For the Turkish economy, Nas and Perry (2000), Neyaptı and Kaya (2001), Akyazı and Artan (2004) and Özer and Türkyilmaz (2005) provide further evidence in support of Friedman's hypothesis that inflation leads to more uncertainty.

In this paper, we aim to re-examine the causal relationships between inflation and inflation uncertainty by applying to TGARCH (threshold generalized autoregressive conditional heteroskedasticity) estimation methodology with a long-span data for the Turkish economy. To this end, the next section describes data and briefly highlights the methodological issues used in the model estimation. The third section is devoted to employing TGARCH modeling to obtain conditional volatility estimates. The causality tests are implemented in section four. The last section summarizes results to conclude the paper.

2.PRELIMINARY DATA AND METHODOLOGICAL ISSUES

The data used consider 588 monthly frequency observations and cover the period from 1960M01 to 2008M12. For any given period t, the inflation data ($INFCPI_t$) are calculated as $[(CPI_t - CPI_{t-1}) / CPI_{t-1}]$ in its linear form using 2000: 100 based consumer price index (CPI_t) taken from the Organization for Economic Co-operation and Development (OECD) electronic statistics portal. Following the seminal paper of Engle (1982), autoregressive conditional heteroskedastic (ARCH) models and their extended version proposed by Bollerslev (1986) as generalized ARCH models have become highly popular in the economics literature to model the conditional volatility in high frequency financial and economic time series. In this sense, many other estimation techniques have also been developed by researchers as the variants of the ARCH family models. In this paper, to construct the proxy variable for inflation uncertainty, we utilize TGARCH modeling introduced independently by Glosten et al. (1993)

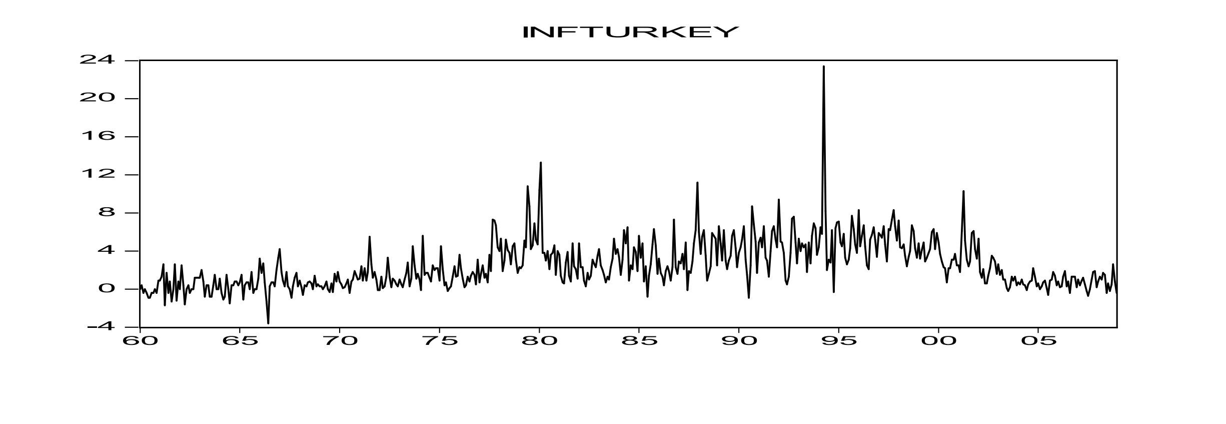
and Zakoïan (1994). For this purpose, let us first define the mean and variance equations as follows:

$$INFCPI_t = c_0 + \sum_{i=1}^{p} \alpha_i INFCPI_{t-i} + \lambda_t DUMSHIFT_t + \varepsilon_t + \phi_1 \varepsilon_{t-1} + \phi_2 \varepsilon_{t-12} \quad (1)$$

$$\sigma_t^2 = \omega_0 + \sum_{i=1}^{p} \gamma_i \varepsilon_{t-i}^2 + \sum_{j=1}^{q} \delta_j \varepsilon_{t-j}^2 d_{t-j} + \sum_{k=1}^{r} \beta_r \sigma_{t-k}^2 \quad (2)$$

where $d_t = 1$ if $\varepsilon_t < 0$ and $d_t = 0$ otherwise. In this model, we expect that negative inflation shocks, $\varepsilon_t < 0$, have a different effect on inflation uncertainty represented by conditional variance series than positive ones. More clearly to say, negative shocks would have an impact $\gamma + \delta$ whereas positive shocks tend to have an effect equal to γ. If $\delta \neq 0$, we mean that these shocks have an asymmetric effect on inflation uncertainty. Following Daal et al. (2005) and Henry et al. (2007), consider that we include a MA(1,12) process into the mean equation to provide a parsimonious estimation by reducing the order of the AR process and to account for possible seasonality in the data. As is used by Caporale and Caporale (2002), we also create a binary variable as a shify dummy ($DUMSHIFT_{80}$ or $DUMSHIFT_{94}$ or $DUMSHIFT_{01}$) and add it into the mean equation, that takes 0 before 1980M01 or 1994M01 or 2001M01, separately, and 1 otherwise to account for any structural change in the economy stemmed from either high social, political and economic uncertainty environment of 1980 military intervention period, or 1994 and 2001 economic / financial crises that the Turkish economy witnessed. To deal with potential model misspecification and to consider the possibility that the residuals of the model are not conditionally normally distributed, we have calculated robust *t*-ratios using the quasi maximum likelihood method suggested by Bollerslev and Wooldridge (1992) so that parameter estimates will be unchanged but the estimated covariance matrix will be altered. The graph and the descriptive statistics of the inflation series are reported below:

Fig. 1 Times Series Graph of the Monthly Turkish Inflation



Tab. 1 Descriptive Statistics

Series $INFCPI_t$
Sample 1960M01 2008M02
Observations 588

Mean	2.47	Skewness	1.75
Median	1.80	Kurtosis	11.6
Maximum	23.4	Jarque-Bera	211.4
Minimum	-3.60	$Q(1)$	249.6
Std. Dev.	2.48	$Q(12)$	1663.0

Fig. 1 indicates that the Turkish inflation indicates a highly volatile characteristic inside the investigation period. This is also highly evident from the fact in Tab. 1 that inflation rates have a high standard deviation nearly equal to its mean value. In Tab. 1, we observe that the mean and median of inflation is 2.48 and 1.80, respectively. Tab. 1 also presents the Ljung-Box Q statistics for the inflation rate at lag k to test for the null hypothesis that there is no autocorrelation of the deviations and the squared deviations of the inflation from its sample mean up to the order k. Skewness is a measure of asymmetry of the distribution of the series around its mean, and the skewness of a symmetric distribution, such as the normal distribution, would be zero. Descriptive statistics reveal that monthly inflation data are biased to the right and has a right tail. On the other hand, kurtosis measures the peakedness or flatness of the distribution of the series, and the kurtosis of the normal distribution is 3. If the kurtosis exceeds 3, the distribution would be peaked relative to the normal. An excess kurtosis can easily be noticed for the inflation series. Jarque-Bera is a test statistic for testing whether the series is normally distributed under the null hypothesis. The test statistic measures the difference of the skewness and kurtosis of the series with those from the normal distribution. In our case, a significant departure from normality due to the excess kurtosis is also found. Finally, $Q(k)$ is the Ljung-Box Q-statistics at lag k to test for the null hypothesis that there is no autocorrelation up to the order k. Results indicate that the large and significant autocorrelations and the significant departure from normality provide ARCH evidence.

As a next step for preliminary data analysis, we test whether it is possible to demonstrate that the data have a stationary characteristic. Spurious regression problem analysed by Granger and Newbold (1974) indicates that using non-stationary time series steadily diverging from long-run mean will produce biased standard errors and unreliable correlations within the regression analysis. This means that the variables must be differenced (d) times to obtain a covariance-stationary process. However, conventional tests for identifying unit roots in a time series such as the conventional augmented Dickey-Fuller (Dickey and Fuller, 1981) and KPSS (Kwiatkowski et al., 1992) tests are criticized strongly in the contemporaneous economics literature when they have been subject to structural breaks which yield biased estimations. For an introductory survey upon these tests, see e.g., Yavuz (2004) and Göktaş (2005). Perron (1989) in his seminal paper argues that conventional unit root tests used by researchers do not consider that a possible known structural break in the trend function may tend too often not to reject the null hypothesis of a unit root in the time series when in fact the series is stationary around a one time structural break. Selecting the date of structural break, that is, assuming that time of break is known *a priori*, however, may not be the most efficient methodology. The actual dates of structural breaks may not be coincided with dates chosen exogenously. Considering these issues, in our paper, we follow the Zivot and Andrews (1992) (henceforth ZA) methodology, allowing the data to indicate breakpoints endogenously rather than imposing a breakpoint from outside the system. The ZA methodology as a further development on Perron (1989) methodology can be explained by considering three possible types of structural breaks in a series, i.e., *Model A* assuming shift in intercept, *Model B* assuming change in slope and *Model C* assuming change in both intercept and slope. For any time series y_t, ZA (1992) test the equation of the form:

$$y = \mu + \eta y_{t-1} + \varepsilon_t \tag{3}$$

Here the null hypothesis is that the series y_t is integrated without an exogenous structural break against the alternative that the series y_t can be represented by a trend-stationary I(0) process with a breakpoint occurring at some unknown time. The ZA test chooses the breakpoint as the minimum t-value on the autoregressive y_t variable, which occurs at time 1 < TB < T leading to $\lambda = TB / T$, $\lambda \in |0.15, 0.85|$, by following the augmented regressions:

Model A:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-j} + \varepsilon_t \tag{4}$$

Model B:

$$y_t = \mu + \beta t + \gamma DT^*_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-j} + \varepsilon_t \tag{5}$$

Model C:

$$y_t = \mu + \beta t + \theta DU_t(\lambda) + \gamma DT^*_t(\lambda) + \alpha y_{t-1} + \sum_{j=1}^{k} c_i \Delta y_{t-j} + \varepsilon_t \tag{6}$$

where DU_t and DT_t are sustained dummy variables capturing a mean shift and a trend shift occuring at the break date respectively, i.e., $DU_t(\lambda) = 1$ if $t > T\lambda$, and 0 otherwise; $DT_t^*(\lambda) = t - T\lambda$ if $t > T\lambda$, and 0 otherwise. Δ is the difference operator, k is the number of lags determined for each possible breakpoint by one of the information criteria and ε_t is assumed to be an identically and independently distributed (i.i.d.) error term. The ZA method runs a regression for every possible break date sequentially and the time of structural changes is detected based on the most significant t-ratio for α. To test the unit root hypothesis, the smallest t-values are compared with a set of asymptotic critical values estimated by ZA. We must note that critical values in the ZA methodology are larger in absolute sense than the conventional ADF critical values since the ZA methodology is not conditional on the prior selection of the breakpoint. Thus, it is more difficult to reject the null hypothesis of a unit root in the ZA test. For the appropriate lag length used in estimations, we consider the Schwarz's Bayesian information criterion (SBIC)-minimizing value.

Tab. 2 ZA Unit Root Tests for the Monthly Inflation Series

Intercept			Trend			Both		
k	min *t*	TB	*k*	min *t*	TB	*k*	min *t*	TB
0	-7.799	63M04	0	-7.554	75M03	0	-7.965	75M03

Notes: Estimation with 0.15 trimmed. Lag length is determined by Schwarz Bayesian information criterion. Min t is the minimum t-statistic. 1% and 5% critical values –intercept: -5.43 and -4.80 ; trend:-4.93 and -4.42; both: -5.57 and -5.08.

The results indicate that for all three cases of the deterministic components in the ZA equation, the stationarity of inflation data cannot be rejected within the period under investigation. From now on, thus, for our empirical purposes in this paper we tend to use the level form data of the monthly inflation series.

3.CONDITIONAL VOLATILITY ESTIMATES

Following the preliminary data issues and methodological discussions, we now try to estimate the mean and variance equations using TGARCH estimation method. The autoregressive (AR) order of mean equation is determined by way of minimizing Akaike model selection information criterion, so various models including different lag structures have been estimated. Beginning from the maximum lag selection 12 and decreasing lag one at a time, we have searched for the true data generating process of our model outlined above and decided to use an AR(12) specification with the smallest estimated statistic as a chosen model. Note also that for the conditional distribution of the error structure, normal (Gaussian) distribution is assumed. The results estimated by the method of maximum likelihood and using Marquardt optimization algorithm as well as quasi-maximum likelihood covariances and standard errors described by Bollerslev and Wooldridge (1992) are presented in Tab. 3 below. Standard errors are given in parentheses.

In Tab. 3, Model 1 refers to the original TGARCH Eq. 1 without any shift dummy variable. Model 2 uses binary variable $DUMSHIFT_{80}$, and Model 3 and Model 4 consider $DUMSHIFT_{94}$ and $DUMSHIFT_{01}$ as a shift dummy, respectively. In addition to the autoregressive variables in the mean equation we see that MA(1) and M(12) process have been found highly significant. As for the dummies created, for the 10% significance level chosen, in Model 2 the dummy $DUMSHIFT_{80}$ yields a statistically insignificant result, but in Model 3 and Model 4 we observe that it turns out to be significant with a negative sign. This means that the shift dummies assigned to the post-economic / financial crises periods 1994 and 2001 catch up a diversification in the course of the level of the inflation in the sense that following the crisis periods, although having a high volatility leading to fluctuations in the mean level, a downward trend seems to dominate the Turkish inflation.

Tab. 3 Estimates of the TGARCH Equations

Dependent Variable: *INFCPI*$_t$
Method: ML – ARCH (Marquardt) – Normal Distribution
Bollerslev-Wooldridge robust standard errors & covariance

	Model 1	Model 2	Model 3	Model 4
	Mean Equation			
c	0.1952	0.2222	0.2432	0.2675
	(0.0989)	(0.0949)	(0.0930)	(0.0935)
α_{-1}	-0.0920	-0.0858	-0.0806	-0.0778
	(0.1012)	(0.1099)	(0.1097)	(0.1014)
α_{-2}	0.0832	0.0788	0.0668	0.0550
	(0.0575)	(0.0429)	(0.0510)	(0.0477)
α_{-3}	0.0027	0.0050	-0.0057	-0.0072
	(0.0401)	(0.0397)	(0.0399)	(0.0392)
α_{-4}	-0.0121	-0.0091	-0.0103	-0.0136
	(0.0300)	(0.0635)	(0.0630)	(0.0587)
α_{-5}	-0.0090	-0.0095	-0.0085	-0.0163
	(0.0312)	(0.0527)	(0.0535)	(0.0525)
α_{-6}	0.0220	0.0214	0.0200	0.0167
	(0.0338)	(0.0328)	(0.0321)	(0.0301)
α_{-7}	0.0229	0.0300	0.0309	0.0270
	(0.0289)	(0.0422)	(0.0415)	(0.0397)
α_{-8}	-0.0006	0.0017	0.0013	-0.0030
	(0.0316)	(0.0388)	(0.0382)	(0.0369)
α_{-9}	0.1317	0.1320	0.1333	0.1311
	(0.0375)	(0.0562)	(0.0575)	(0.0563)
α_{-10}	-0.0500	-0.0475	-0.0457	-0.0427
	(0.0467)	(0.0432)	(0.0446)	(0.0422)
α_{-11}	0.1012	0.1013	0.1022	0.1023
	(0.0356)	(0.0351)	(0.0348)	(0.0339)
α_{-12}	0.6384	0.6506	0.6553	0.6687
	(0.0685)	(0.0720)	(0.0726)	(0.0684)
MA(1)	0.3964	0.3829	0.3699	0.3501
	(0.1254)	(0.1328)	(0.1368)	(0.1247)
MA(12)	-0.3358	-0.3529	-0.3615	-0.3823
	(0.0616)	(0.0769)	(0.0797)	(0.0770)
DUMSHIFT	---------	-0.1236	-0.1826	-0.2353
		(0.1036)	(0.0924)	(0.0845)
	Variance Equation			
ω	0.1241	0.1269	0.1433	0.1443
	(0.0397)	(0.0357)	(0.0413)	(0.0408)
γ	0.5083	0.5035	0.5273	0.5341
	(0.2267)	(0.2284)	(0.2529)	(0.2502)
δ	-0.4020	-0.3986	-0.4158	-0.4117
	(0.2010)	(0.2216)	(0.2390)	(0.2400)
β	0.7155	0.7165	0.7014	0.6950
	(0.0326)	(0.0587)	(0.0653)	(0.0638)
Adj. R^2	0.5001	0.4842	0.4849	0.4889
F-stat.	30.959	29.413	29.492	29.130
Durbin Watson stat.	1.7649	1.7548	1.7392	1.7016

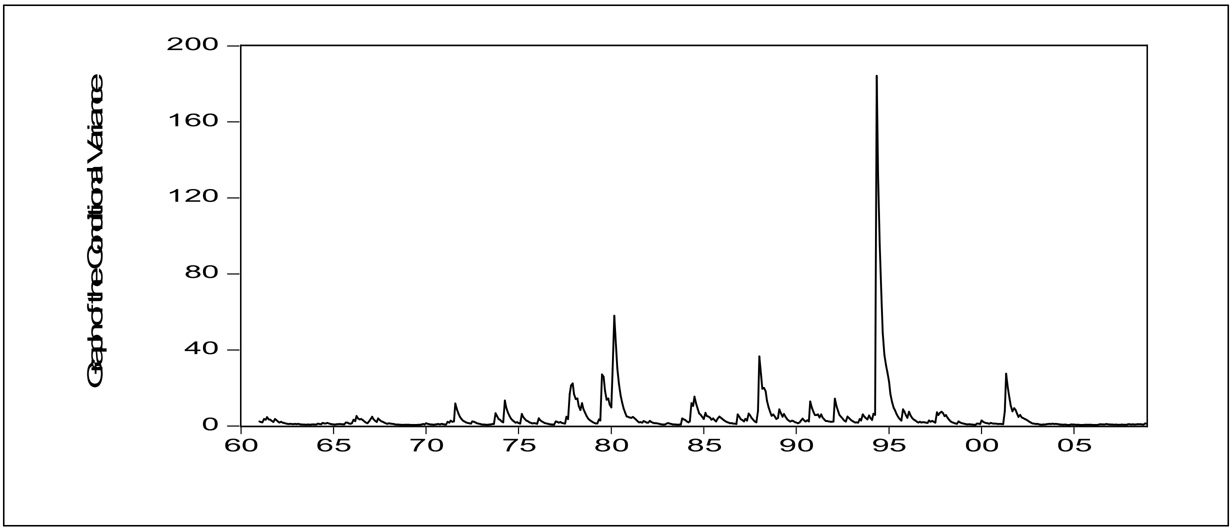
In the variance equation, the ARCH term (ε_{t-1}^2) gives the news about volatility from the previous period measured as the lag of the squared residual from the mean equation, and the lagged GARCH term (σ_{t-1}^2) refers to the last period's forecast variance. Notice that the ARCH and GARCH coefficients have highly similar characteristics to each other within own groups, no matter which model is considered. All the ARCH term coefficients have statistical significance under 5% probability levels, while the GARCH term coefficients are significant under 10% levels. As can be seen, the degree of persistence in the conditional variance is quitely high. These results mean that the information content of the forecasts of the conditional variance extracted from the TGARCH methodology used in this paper has in fact been of a special importance in modeling inflation and its uncertainty. Due to the statistical significance of the threshold term in all equations, we can infer that the news impact seems to be asymmetric. Following also Caporale and Caporale (2002), the coefficient for which the order is set to one in the variance equation indicates that positive inflationary shocks are associated with statistically significant and quantitatively larger levels of inflation uncertainty than are negative shocks. The impact of positive shocks on inflation are about five times greater than that of negative shocks. Thus we verify the asymmetric nature of the relationship between inflation and inflation uncertainty for the Turkish economy. We report in Tab. 4 below that the model satisfies the null hypothesis that there remains no autocorrelation problem of the 12th order.

Tab. 4 Some Diagnostics for the TGARCH Models

	Model 1	Model 2	Model 3	Model 4
Q(12)	6.5552	6.1332	6.1025	6.9429
$Q^2(12)$	4.4476	4.3387	4.2737	4.4996
ARCH LM(12) F-stat	0.0741	0.3514	0.3592	0.3636

Finally we give below the graph of the conditional variance series extracted from the TGARCH equation:

Fig. 2 Graph of the Conditional Variance



In Fig. 2, it is highly explicit that the late-1970s and early-1980s, the late 1980s, and the 1994 and 2001 economic / financial crisis periods witness a considerable increase in the conditional variance of the inflation. Notice that this is explicitly evident especially for the 1994 crisis period.

4.GRANGER CAUSALITY ESTIMATES

As a next step in our empirical modeling for the inflation and inflation uncertainty relationship, for which the latter is governed by the conditional variance series obtained through the TGARCH model constructed in the former section, we try to implement some Granger causality tests. We must specify that for the conditional variances considered, estimates from Model 1 in Tab. 3 are used. The Granger causality between the two variables, say X and Y, asks that how much of the current X can be explained by a regression on its past values, and then tries to test whether inclusion of the lagged values of Y into the regression to explain X have statistical significance as a whole. If so, we can infer that Y helps predict the course of X, or in other words, X is Granger-caused by Y. More formally, to test the causal relationship between the variables X_t and Y_t observed in a given period t, let us write the bivariate regressions as follows:

$$X_t = c_0 + \sum_{i=1}^{n} \rho_1 X_{t-n} + \sum_{i=1}^{n} \rho_2 Y_{t-n} + \varepsilon_t \tag{7}$$

$$Y_t = d_0 + \sum_{i=1}^{n} \rho_3 Y_{t-n} + \sum_{i=1}^{n} \rho_4 X_{t-n} + u_t \tag{8}$$

where c_o and d_0 denote the constant terms in the Granger regressions, ρ's are relevant coefficients in the Granger equations, n represents the lag length chosen for causality analysis, which is assumed in principle to correspond to the expectations for the longest time over which variables could predict the others, and ε_t and u_t are assumed as white-noise error terms in the regressions. Note that the null hypothesis in Eq. 7 is that the lags of Y_t are not significant as a whole, that is to say, Y_t does not Granger-cause X_t. Likewise, the null hypothesis in Eq. 8 is that the lags of X_t have no statistical significance in explaining Y_t, which also means that X_t does not Granger-cause Y_t. By employing F-type Wald tests, the results of pairwise Granger causality analyses which are applied on the joint significance of the sum of lags of each explanatory variable are reported below. Following Nas and Perry (2000) and Daal et al. (2005), since Granger causality tests initially indicate the temporal ordering or precedence relationship between each variable but do not reveal the sign of this relationship, we also give below the sign of the sum of the coefficients taken from each Granger equation to determine whether the Granger causality, if estimated, is in the positive or negative way. For the causality tests, various lag lengths are considered to see whether the estimation results are sensitive to the *a priori* lag selection. The asterisks ***, ** and * indicate significance at the 0.01, 0.05 and 0.10 levels, respectively. The signs (+) and (−) are used for the process by which the sum of the coefficients of Granger equation yields a positive or negative sign, respectively.

Tab. 5 Granger Causality Tests

Lag	*H_0: Inflation does not Granger-cause Inflation Uncertainty*	*H_0: Inflation Uncertainty does not Granger-cause Inflation*
3	108.841 (+)***	7.7805 (–)***
6	56.9807 (+)***	7.9084 (–)***
12	33.4626 (+)***	3.7160 (–)***
18	24.1544 (+)***	2.8049 (–)***
24	19.7678 (+)***	2.3563 (–)***

In Tab. 5, we cannot reject the null hypothesis of Granger causality between inflation and its uncertainty mutually in a strong way at the 0.01 level. When we examine the estimation findings, we can easily notice that data from the Turkish economy give support to the Friedman-Ball hypothesis that inflation Granger causes inflation uncertainty in positive way, that is to say differently, inflation precedes the course of inflation uncertainty as *a priori* hypothesized mainly by Friedman (1977). However we find a significant causal relationship running from inflation uncertainty to inflation at the 0.01 level, the sign of the sum of the coefficients in this case turns out to be negative contradicting what the Cukierman-Meltzer hypothesis adduces. This means the larger the inflation uncertainty the lower would likely to be the level of inflation. Holland (1995) explains as a possible reason of this case that an increase in inflation uncertainty can be viewed by policymakers as costly, so induces them to fight inflation to reduce it in the future. Nas and Perry (2000) also touchs upon the issue that inflation and associated uncertainty create real costs, which lead policy authorities to monetary tightening stabilization efforts to lower inflation. Therefore we can infer that inside the period we examine, had there not been chronic and high inflation rates subject to the Turkish economy, other things being equal, the uncertainty component stemmed from inflation in the economy could have been decreased by the policy makers. Furthermore we observe that monetary authorities seem to be fighting inflation due to the uncertainties associated with inflation occurred in the economy.

5.CONCLUDING REMARKS

In this paper, we try to examine the preceding / causal relationships between inflation and inflation uncertainty in the Turkish economy. For this purpose, we initially extract the knowledge of the conditional volatility from the data by using contemporaneous threshold generalized autoregressive conditional heteroskedasticity (TGARCH) estimation technique. Dealing with the information content of this relationship, we find that positive inflationary shocks are associated with statistically significant and quantitatively larger levels of inflation uncertainty than are negative shocks and that the impact of positive shocks on inflation are about five times greater than that of negative shocks. Our estimation results indicate that inflation in fact leads to inflation uncertainty in line with the Friedman-Ball hypothesis. However, our results contradict the Cukierman-Meltzer hypothesis that inflation uncertainty leads to inflation in a positive way. We find that the larger the inflation uncertainty the lower would likely to be the level of inflation. Such a finding can be attributed to the inference that inside the period examined, monetary authorities seem to be fighting inflation due to the uncertainties associated with inflation occurred in the economy.

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REFERENCES

Akyazı, H. and Artan, S. (2004), “Türkiye’de Enflasyon-Enflasyon Belirsizliği İlişkisi ve Enflasyon Hedeflemesinin Enflasyon Belirsizliğini Azaltmadaki Rolü”, **Türkiye Bankalar Birliği Bankacılar Dergisi**, 48, 3-17.

Ball, L. and Cecchetti, S.G. (1990), “Inflation and Uncertainty at Short and Long Horizons”, **Brookings Papers on Economic Activity**, 1, 215-54.

Ball, L. (1992), “Why Does Higher Inflation Raise Inflation Uncertainty?”, **Journal of Monetary Economics**, 29, 371-78.

Barro, R.J. and Gordon, D.B. (1983), “A Positive Theory of Monetary Policy in a Natural Rate Model”, **Journal of Political Economy**, 91, 589-610.

Bollerslev, T. (1986), “Generalized Autoregressive Conditional Heteroskedasticity”, **Journal of Econometrics**, 31, 307-27.

Bollerslev, T. and Wooldridge, J.M. (1992), “Quasi-maximum Likelihood Estimation and Inference in Dynamic Models with Time Varying Covariances”, **Econometric Reviews**, 11, 143-72.

Caporale, B. and Caporale, T. (2002), “Asymmetric Effects of Inflation Shocks on Inflation Uncertainty”, **Atlantic Economic Journal**, 30/4, 385-88.

Cukierman, A. (1992), **Central Bank Strategy, Credibility, and Independence**. Cambridge: MIT Press.

Cukierman, A. and Meltzer, A. (1986), "A Theory of Ambiguity, Credibility, and Inflation under Discretion and Asymmetric Information", **Econometrica**, September, 54, 1099-128.

Daal, E., Naka, A. and Sanchez, B. (2005), "Re-examining Inflation and Inflation Uncertainty in Developed and Emerging Countries", **Economics Letters**, 89, 180-86.

Dickey, D.A. and Fuller, W.A. (1981), "Likelihood Ratio Statistics for Autoregressive Time Series with Unit Roots", **Econometrica**, 49, 1057-072.

Engle, R.F. (1982), "Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of U.K. Inflation", **Econometrica**, 50, 987-1008.

Fountas, S. (2001), "The Relationship between Inflation and Inflation Uncertainty in the UK: 1885-1998", **Economics Letters**, 74, 77-83.

Fountas, S., Karanasos, M. and Kim, J. (2002), "Inflation and Output Growth Uncertainty and their Relationship with Inflation and Output Growth", **Economics Letters**, 75, 293-301.

Friedman, M. (1977), "Nobel Lecture: Inflation and Unemployment", **Journal of Political Economy**, 85/3, 451-72.

Glosten, L.R., Jaganathan, R. and Runkle, D. (1993), "On the Relation between the Expected Value and the Volatility of the Nominal Excess Returns on Stocks", **Journal of Finance**, 48, 1779-801.

Göktaş, Ö. (2005), **Teorik ve Uygulamalı Zaman Serileri Analizi**, İstanbul: Beşir Kitabevi.

Granger, C.W.J. and Newbold, P. (1974), "Spurious Regressions in Economics", **Journal of Econometrics**, 2/2, 111-20.

Grier, K.B. and Perry, M.J. (1998), “On Inflation and Inflation Uncertainty in the G7 Countries”, **Journal of International Money and Finance**, 17/4, 671-89.

Henry, Ó.T., Olekalns, N., and Suardi, S. (2007), “Testing for Rate Dependence and Asymmetry in Inflation Uncertainty: Evidence from the G7 Economies”, **Economics Letters**, 94, 383-88.

Holland, A.S. (1995), “Inflation and Uncertainty: Tests for Temporal Ordering”, **Journal of Money, Credit, and Banking**, 27/3, 827-37.

Kontonikas, A. (2004), “Inflation and Inflation Uncertainty in the United Kingdom, Evidence from GARCH Modelling”, **Economic Modelling**, 21, 525-43.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P. and Shin, Y. (1992), “Testing the Null Hypothesis of Stationary against the Alternative of a Unit Root”, **Journal of Econometrics**, 54, 159-78.

Nas, T.F. and Perry, M.J. (2000), “Inflation, Inflation Uncertainty, and Monetary Policy in Turkey: 1960-1998”, **Contemporary Economic Policy**, 18/2, 170-80.

Neyaptı, B. and Kaya, N. (2001), “Inflation and Inflation Uncertainty in Turkey: Evidence from the Past Two Decades”, **Yapı Kredi Economic Review**, 12/2, 21-5.

Okun, A. (1971), “The Mirage of Steady Inflation”, **Brookings Papers on Economic Activity**, 2, 485-98.

Özer, M. and Türkyılmaz, S. (2005), “Türkiye’de Enflasyon ile Enflasyon Belirsizliği Arasındaki İlişkinin Zaman Serisi Analizi”, **İktisat İşletme ve Finans**, 5/229, 93-104.

Perron, P. (1989), “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis,” **Econometrica**, 57, 1361-401.

Yavuz, N.Ç. (2004), “Durağanlığın Belirlenmesinde KPSS ve ADF Testleri: İMKB Ulusal-100 Endeksi ile bir Uygulama”, **İstanbul Üniversitesi İktisat Fakültesi Mecmuası**, 54/1, 239-47.

Zakoïan, J.M. (1994), “Threshold Heteroskedastic Models”, **Journal of Economic Dynamics and Control**, 18, 931-44.

Zivot, E. and Andrews, D.W.K. (1992). “Further Evidence of Great Crash, the Oil Price Shock and the Unit Root Hypothesis”, **Journal of Business and Economic Statistics,** 10, 251-70.