

Article

Inflation and Inflation Uncertainty Revisited: Evidence from Egypt

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Abstract: The welfare costs of inflation and inflation uncertainty are well documented in the literature and empirical evidence on the link between the two is sparse in the case of Egypt. This paper investigates the causal relationship between inflation and inflation uncertainty in Egypt using monthly time series data during the period January 1974–April 2015. To endogenously control for any potential structural breaks in the inflation time series, Zivot and Andrews (2002) and Clemente–Montanes–Reyes (1998) unit root tests are used. The inflation–inflation uncertainty relation is modeled by the standard two-step approach as well as simultaneously using various versions of the GARCH-M model to control for any potential feedback effects. The analyses explicitly control for the effect of the Economic Reform and Structural Adjustment Program (ERSAP) undertaken by the Egyptian government in the early 1990s, which affected inflation rate and its associated volatility. Results show a high degree of inflation–volatility persistence in the response to inflationary shocks. Granger-causality test along with symmetric and asymmetric GARCH-M models indicate a statistically significant bi-directional positive relationship between inflation and inflation uncertainty, supporting both the Friedman–Ball and the Cukierman–Meltzer hypotheses. The findings are robust to the various estimation methods and model specifications. The findings of this paper support the view of adopting inflation-targeting policy in Egypt, after fulfilling its preconditions, to reduce the welfare cost of inflation and its related uncertainties. Monetary authorities in Egypt should enhance the credibility of monetary policy and attempt to reduce inflation uncertainty, which will help lower inflation rates.

Keywords: Inflation; Inflation uncertainty; GARCH-M; Granger causality; Egypt

JEL Classifications: C22; E31; E52

1. Introduction

The welfare costs of inflation and inflation uncertainty are well documented in the literature. Several studies found that high non-predictable inflation distorts relative prices and the inter- and intra-temporal decisions of economic agents, leading to an inefficient allocation of resources and lower output growth (Friedman, 1977 [1]; Dotsey and Ireland; 1996 [2]; Lucas, 2003 [3]). For example, using a general equilibrium monetary model, Dotsey and Ireland (1996) [2] found that high inflation distorts a variety of marginal decisions. These include diverting resources from productive to speculative investment, reducing holdings of real cash balances and substituting leisure for work. Similarly, the real effects of inflation uncertainty have been established in the literature since Friedman's Nobel lecture on inflation and unemployment in 1977 in which he showed that high inflation uncertainty distorts the informational content of prices. This reduces the ability of the price system to allocate resources efficiently. Moreover, high inflation uncertainty distorts individual decisions on consumption, saving and investment. High uncertainty makes it expensive for investors to borrow and to detect profitable investment opportunities. Elder (2004) [4] found that U.S. output growth decreased by 22 basis points, over a three month period, due to an average shock to inflation uncertainty. In another study, Fountas *et al.* (2006) [5] found, using data on G7 countries, that inflation causes negative welfare effects, both directly and indirectly through the inflation uncertainty channel. They also found that more inflation uncertainty provides an incentive to monetary authorities to surprise the public by raising inflation rates. Inflation uncertainty could also have a negative effect on economic growth by increasing the uncertainty of investments' profits and hence reducing investments. In a recent study, Chowdhury (2014) [6] found that inflation uncertainty is detrimental to output growth in India.

Inflation rates in Egypt have increased significantly in recent years, reaching 18% in 2008, after a period of low rates at the beginning of the new millennium at 2% in 2001. This increase in inflation rates was associated with an increase in inflation volatility. Given the substantial economic costs of inflation and inflation uncertainty, understanding the nature of the link between the two variables become an interest to academics and policy makers as this will help guide in designing policies that achieve macroeconomic stability and increase economic welfare.

The objective of this paper is to examine the causal relationship between inflation and inflation uncertainty, using Egyptian data on which limited research has been conducted. A special feature of this paper is that both the two-step approach as well as simultaneous modeling approach is incorporated to check the robustness of results.

The paper is organized as follows: Section 2 presents a review of the theoretical and empirical literature. Section 3 presents a brief historical development of inflation rate in Egypt. The data is described in section 4. The empirical methods of the paper are presented in Section 5 and the obtained results are discussed in Section 6. Section 7 summarizes the findings of the paper and discusses the policy implications.

2. Literature Review

The substantial economic costs of inflation and inflation uncertainty have inspired a growing literature to examine the link between the two. Two main competing theories exist in the literature to explain the relationship between inflation and inflation uncertainty and its causal direction. In his Nobel lecture on inflation and unemployment, Friedman (1977) [1] considers a positive causation from inflation-level to inflation-uncertainty. Friedman's hypothesis was further extended by Ball (1992) [7] within the context of a game with asymmetric information between the public and the monetary authority. In Ball's (1992) [7] model, two types of policy makers exist: one is unwilling to disinflate because of the fear of the resulting recessionary effects, and the other is willing to bear the cost of disinflation. In this model, high inflation leads to greater uncertainty because the public is uncertain about the type of future policy maker.

As opposed to the Friedman–Ball hypothesis, Cukierman and Meltzer (1986) [8] hypothesize that causality runs from inflation-uncertainty to inflation-level. According to this argument, during periods of high inflation uncertainty, the monetary policy is discretionary due to lack of a commitment mechanism. This gives an incentive to the monetary authority to act opportunistically to stimulate output growth by making monetary expansions, which lead to higher inflation.

In contrast to the dominant hypothesis of a positive association between inflation level and inflation uncertainty, other theories postulate a negative relationship (see for example Pourgerami and Maskus, 1987 [9]; Holland, 1995 [10]). According to Pourgerami and Maskus (1987) [9] higher inflation reduces inflation uncertainty as people invest resources to anticipate the future inflation rate better and to shelter themselves from its adverse effects. The same view was further developed by Ungar and Zilberfarb (1993) [11]. In support of this negative relationship, Holland (1995) [10] argues that higher inflation uncertainty reduces the inflation rate due to a stabilization motive by the monetary authority, which seeks to reduce the welfare costs of inflation by disinflationary policies when inflation uncertainty is high. This is known in the literature as the “*Stabilizing Fed hypothesis*” and postulates a negative effect of inflation uncertainty on inflation level.

A growing body of empirical literature has emerged to examine the causal relationship between inflation and its uncertainty, in a wide range of countries, using different econometric techniques, and covering different time periods, with mixed findings (for a survey of the literature see Davis and Kanago (2000) [12]). The Friedman–Ball hypothesis is supported, for example, by: Grier and Perry (1998) [13], Daal *et al.* (2005) [14], Thornton (2008) [15], Keskek and Orhan (2010) [16], and Karahan (2012) [17]. On the other hand, the Cukierman and Meltzer hypothesis is supported by the findings of Conrad and Karanasos (2005) [18], Wilson (2006) [19], and Fountas (2010) [20]. Keskek and Orhan (2010) [16] found support for the stabilizing Fed hypothesis in the case of Turkey, and Thornton (2007) [21] found evidence for this hypothesis in Colombia, Mexico and Turkey.

Empirical evidence on the direction of causality between inflation and inflation uncertainty in cross-country studies is inconclusive. For example, in a cross country study of twelve emerging market economies, and using Granger causality test, Thornton (2007) [21] found strong support for the Friedman–Ball hypothesis in all of the studied countries, while the evidence on the effect of inflation uncertainty on inflation was mixed. The author found that higher inflation uncertainty leads to lower inflation rate in Colombia, Mexico and Turkey, which supports Holland hypothesis, while in Hungary,

Indonesia, and Korea, the Cukierman and Meltzer hypothesis holds. In a recent study, Zivkov *et al.* (2015) [22] examined, using GARCH type model along with quantile regression, the link between inflation and inflation uncertainty in eleven Eastern European countries, and found mixed evidence on the direction of causality. Both Friedman's and Cukierman–Meltzer's hypotheses have been confirmed mainly for the largest Eastern European countries with flexible exchange rate, while no support for any of these hypotheses was found in smaller, open economies with fixed exchange rate regime. In a panel study of 105 countries over the period 1960–2007, Kim and Lin (2012) [23] used a system of simultaneous equations and found a two-way interaction between inflation and its variability, which supports both the Friedman–Ball and the Cukierman–Meltzer Hypotheses.

While the extant literature is largely dominated by cross-country studies, several individual country studies have investigated the relationship between inflation and inflation uncertainty. Similar to the cross-country studies, evidence based on country-specific studies on the direction of causality between inflation and inflation uncertainty is equally mixed. For example, using data from U.K. over the period 1972–2002, and using different GARCH-M models of inflation, Kontonikas (2004) [24] found a positive relationship between past inflation and uncertainty about future inflation, which is consistent with the Friedman–Ball causal link. Kontonikas (2004) [24] also found that the adoption of inflation target in UK since 1992 has helped eliminate inflation persistence and reduced long-run uncertainty. Using a two-step procedure, Karahan (2012) [17] examined the relationship between inflation and inflation uncertainty in Turkey from 2002 to 2011 and found evidence in favor of the Friedman–Ball hypothesis that inflationary periods result in high inflation uncertainty in the Turkish case. In another study, Heidari and Bashiri (2010) [25] used the Full Information Maximum Likelihood method to examine the inflation–inflation uncertainty relationship in Iran and found support for the Friedman–Ball hypothesis.

Using quarterly data from Australia during the period 1949 to 2013, Hossain, (2014) [26] found a feedback relation between inflation and inflation volatility with an adverse effect of inflation volatility on the rate of unemployment. The author also found that inflation shock has an asymmetric impact on inflation volatility. Payne (2009) [27] found support for Holland's stabilization hypothesis in Thailand. Using Granger-causality test, he found that an increase in inflation causes an increase in inflation uncertainty, while an increase in inflation uncertainty causes a decrease in inflation in Thailand. Payne (2009) [27] also found that inflation targeting reduced inflation uncertainty in response to inflationary shocks.

Several country-specific studies found a two way causal relation between inflation and its uncertainty. For example, using GARCH model and Granger Causality test, Chowdhury (2014) [6] investigated the relationship between inflation and inflation uncertainty in India and found a positive feedback relationship between the two variables, supporting both the Friedman–Ball and Cukierman–Meltzer hypotheses. The author also found that inflation uncertainty is detrimental to output growth in India. Keskek and Orhan (2010) [16] examined, using various types of GARCH-M models, the relationship between inflation and inflation uncertainty in Turkey over the period 1984 to 2005. They found that higher inflation rates lead to greater inflation uncertainty, while the effect of inflation uncertainty on inflation is found to be negative due to a stabilization motives' dominating the opportunistic incentives of monetary authorities.

The mixed findings found in the previous empirical studies on the direction of causality between inflation and inflation uncertainty could be due to differences with respect to the sample of countries

that are examined, the used econometric technique, and the covered time periods. There are naturally socio-economic and country-specific factors that differ between countries. Countries may also differ in their macroeconomic policies, the nature of economic problems they face and the dynamics of these problems over time. Accordingly, it is to be expected that, in practice, the inflation–inflation uncertainty relationship is country-specific, and varies depending on the period under investigation.

3. Inflation in Egypt: A Brief Historical Background

Keeping inflation low, stable and predictable has continually remained a key objective of the macroeconomic policies in Egypt. However, data shows that this is merely an objective and has not often been realized. During the study period, the Egyptian economy has witnessed substantial economic and structural changes, as well changes in the economic policies, both fiscal and monetary. These developments have affected the inflation rates in Egypt. Figure 1 displays the evolution of the annual inflation rate in Egypt during the period from 1975 to 2014. Inflation rates were in general relatively high during the 1980s before dropping substantially in the 1990s after the adoption of the Economic Reform and Structural Adjustment Programme. Afterwards, inflation rates started to increase once again reaching 18% in 2008 compared to 2% in the year 2001.

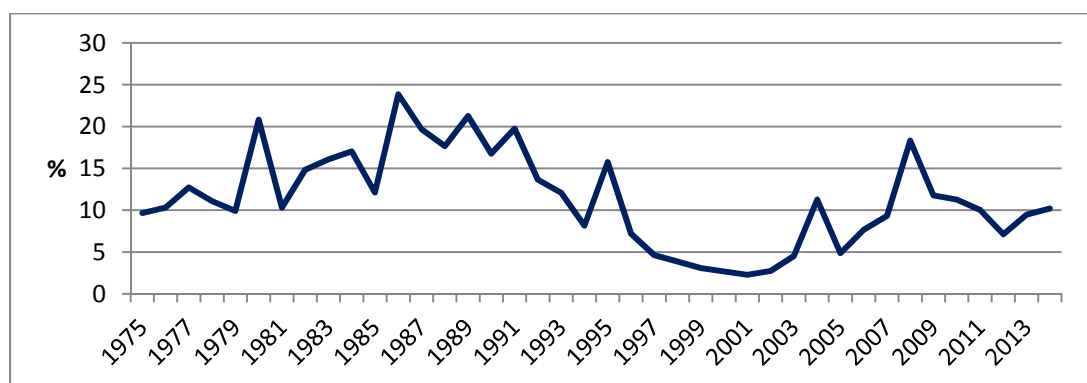


Figure 1. Annual Inflation Rate in Egypt (1975–2014). Source: Data is obtained from the International Financial Statistics database.

The adoption of an outdoor economic policy in the second half of the 1970's has lead to a significant increase in imports and a deterioration of the current account. During that period, government budget deficit increased substantially and was largely financed by money creation and debt financing, all of which exerted upward pressure on inflation rate. The budget deficit deteriorated further as a result of the expansion in government subsidies to mitigate the inflationary pressures resulting from the rising international prices. The growing money supply and domestic liquidity have increased aggregate demand and hence inflation rates. Consequently, inflation rates at the end of the 1970s doubled, within one year, jumping dramatically from 10% in 1979 to 21% in 1980. In 1981, inflation rate dropped back to 10%. During the 1980s, inflation rate averaged 17%, and reached its highest value in 1986 at 24%.

To face the deteriorating economic conditions, including high unemployment and high inflation rates, as well as the substantial deficit in the balance of payment and government budget at the end of the 1980s, in the early 1990s, Egypt adopted a battery of reform policies under the Economic Reform and Structural Adjustment Program (ERSAP) after consultation with the IMF and the World Bank, to restore

the internal as well as the external balance. Macroeconomic stabilization, including inflation reduction, was a key objective of this program. One important policy change that had an implication for inflation rate was the switch in financing budget deficits from monetary expansion to non-inflationary finance measures such as Treasury bill auctions. This was also associated with a drop in budget deficit and a significant reduction in external debts and its associated servicing. In 1991, Egypt's participation in the coalition to free Kuwait in the first Gulf war was generously rewarded by the Paris Club creditors. Egypt was granted a debt relief package of around \$20 billion, which enabled Egypt to save an average of over two percentage points of GDP a year in servicing the debts from 1992 to 1997. As a result of this debt relief, along with the ERSAP program and its associated structural adjustment policies, inflation rates decreased substantially from 20% in 1991 to only 3% in 1999 and in 2001, inflation rate reached its lowest level at 2%.

In the new millennium, Egypt continued in adopting economic reform policies in several areas such as foreign exchange rates, taxes, and banking systems. In 2003, a managed floating exchange rate system came into effect to replace the old system that pegs the Egyptian pound against the US dollar. This flexibility in the exchange rate was intended to achieve price stability and reduce the inflationary pressures by correcting for the overvaluation of the Egyptian pound and to face the growing black market for currency, all of which will enhance the competitiveness of the Egyptian exports. Following the new floating system, the Egyptian pound depreciated by around 25%. However, the depreciation of the Egyptian pound seems to have contributed rather than to remedy the increase in the inflation rate given the high dependence on imports to satisfy domestic demand for most of the essential goods. In addition, changes in Egypt's nominal exchange rate were very limited and have experienced a slow depreciation during 2003–2012, pointing to a *de facto* crawling peg regime. Despite the undertaken policy reforms, inflation rates increased significantly during this period, reaching 11% in 2004 and 18% in 2008.

4. Data

This paper uses monthly time series data on the Egyptian consumer price index (CPI) during the period January 1974–April 2015. Data on the CPI are obtained from the International Financial Statistics database, issued by the International Monetary Funds. Inflation is measured as the first difference of the seasonally adjusted log consumer price index, where $\pi_t = (\ln \text{CPI}_t - \ln \text{CPI}_{t-1}) \times 100$ (the difference from the moving average method is used to seasonally adjust the log of the CPI).

Figure 2 plots the Egyptian seasonally adjusted monthly inflation rate over the period January 1974–April 2015. It is evident from the figure that the monthly inflation rate is highly volatile and that the extent of this volatility varies from one period to another. In particular, inflation experienced wide volatility especially in the period preceding the ERSAP program.

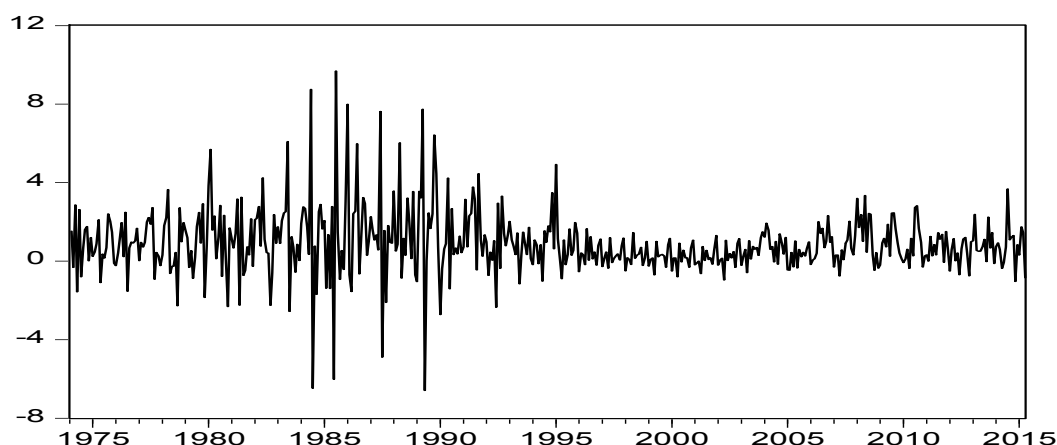


Figure 2. The Egyptian seasonally adjusted monthly inflation rate over the period January 1974–April 2015.

The summary statistics of the inflation rate time series presented in Table 1 show that the inflation distribution deviates from the normal distribution with positive skewness and bigger peakedness and this is also shown by the statistically significant Jarque–Bera statistic.

Table 1. Summary statistics of the monthly inflation rate over the period January 1974–April 2015. Source: Author’s compilation based on data from International Financial Statistics database.

Mean	Median	Std.Dev	Skewness	Kurtosis	Jarque-Bera
0.88	0.71	1.62	0.77	9.76	994.2* (0.00)

* Indicates rejection of the null hypothesis of zero skewness and zero excess kurtosis at 1% significance level.

5. Materials and Methods

5.1. Unit Root Tests of Inflation Time Series

Prior to any estimation, it is essential to check whether the inflation time series is stationary to ensure a non-spurious estimation and to have time-invariant estimates. In this regard, two unit root tests are used; the Augmented Dickey–Fuller (ADF) test and the Phillips–Perron (PP) test. Three versions of the Augmented Dickey–Fuller (ADF) and the Phillips–Perron tests are conducted. One version allows for an intercept, a second allows for an intercept and a deterministic trend, and a third version excludes the intercept and the deterministic trend. The ADF test is applied to the following model:

$$\Delta y_t = \alpha + \beta T + \gamma y_{t-1} + \sum_{i=1}^p \mu_i \Delta y_{t-i} + u_t \quad (1)$$

where, α is an intercept, β is the coefficient on the time trend T , Δ is a difference operator, μ_i is the coefficient on the lagged dependant variable, p is the lag order of the autoregressive process and $u_t \sim i.i.d$ random variable. The hypotheses of the ADF are H_0 : the time series contains a unit root *versus* H_1 : the time series is stationary. The null hypothesis is rejected if the ADF statistic, defined as the t -ratio

of the coefficient γ in Equation (1), is greater than the critical value from the Dickey–Fuller table. The PP test is similar to the ADF test but it uses a non-parametric correction of any serial correlation and heteroskedasticity in the errors (u_t) of the test regression by directly modifying the test statistics.

One limitation of the ADF and PP unit root tests is their inability to control for structural breaks in a time series with the confusion of considering structural breaks in the series as evidence of non-stationarity. Perron (1990) [28] showed that not allowing for structural breaks in the series when testing for unit root leads to inaccurate hypothesis testing. While Perron (1990) [28] suggested allowing for an exogenous structural break in the Augmented Dickey–Fuller (ADF) tests, Zivot and Andrews (2002) [29] developed a way that allows for an endogenously determined structural break. A key limitation of the Zivot and Andrews (2002) [29] unit root test is its inability to account for multiple break points in a time series. To address this, Clemente, Montanes and Reyes (1998) [30] proposed a test to endogenously account for two structural breaks in a series. This test has two versions, one that allows for any gradual shift in the mean of the series known as Innovational Outlier (IO) model, and the other version of the test allows for a sudden shift in the time series known as Additive Outlier (AO) model. The advantage of these tests is that they do not require any *a priori* knowledge of the structural break dates, which will be endogenously determined. To allow for the possibility of structural breaks in the inflation time series, the current paper uses the Clemente, Montanes and Reyes (1998) [30] test to model any potential additive outlier (AO) or innovational outlier (IO) schemes, this is in addition to the Zivot and Andrews (2002) [29] unit root test.

5.2. Modeling Inflation and Inflation Uncertainty Relationship: The Two-Step Approach

A standard approach used to examine the relationship between inflation and inflation uncertainty is a procedure called the two-step procedure. According to this approach, an ARMA-GARCH model of inflation is estimated in the first step, with the generated conditional variance used as a measure of inflation uncertainty. In the second step, a Granger causality test is performed to determine the direction of causality between the level of inflation and the generated inflation uncertainty measure from the first step. Among the studies that used this approach include Grier and Perry (1998) [13], Nas and Perry (2000) [31], Daal *et al.* (2005) [14], Payne (2008) [32], and Karahan (2012) [17]. For example, Nas and Perry (2000) [31] examined the relation between inflation and inflation uncertainty in Turkey from 1960 to 1998 using the two-step procedure. Similarly, Payne (2008) [32] carried out a corresponding study of the Caribbean region.

Studies using the two-step procedure consider that sequential modeling is appropriate since inflation may influence inflation uncertainty over several periods. Moreover, lagged inflation in the conditional variance equation can create problems with the non-negativity of the variance (Payne, 2008 [32]). In the first step of this sequential modeling, the mean equation of inflation is modeled using an Autoregressive, Moving average ARMA (p, q) model, namely,

$$\pi_t = \alpha_0 + \sum_{i=1}^p \alpha_i \pi_{t-i} + \sum_{j=1}^q \beta_j \varepsilon_{t-j} + \varepsilon_t \quad (2)$$

in which $E(\varepsilon_t | \alpha_{t-1}) = 0$; $\text{var}(\varepsilon_t | \alpha_{t-1}) = \sigma_t^2$.

In Equation (2), the conditional mean of inflation has optimal lag length (p, q) , which is determined by the Schwartz Bayesian Criterion (SBIC) and the Akaike Information Criterion (AIC). (Algebraically, the information criteria could be expressed as

$$AIC = \ln(\hat{\sigma}^2) + \frac{2k}{T}$$

$$SBIC = \ln(\hat{\sigma}^2) + \frac{k}{T} \ln T$$

where $\hat{\sigma}^2$ is the residual sum of squares divided by the number of observations (T), and k is the total number of estimated parameters).

Before modeling inflation uncertainty, it is essential to check whether the conditional variance of the error terms $\hat{\sigma}_t^2$ in Equation (2) has Autoregressive Conditional Heteroskedasticity (ARCH) effects, and the residuals are serially uncorrelated. For these purposes, a set of diagnostic tests is used. These include: Engle's (1982) Lagrange Multiplier test (LM) for the existence of ARCH effects; the Ljung–Box Q -test and the Breusch–Godfrey tests for detecting serial correlation in the residuals.

The LM test involves regressing the square of the OLS residuals, $\hat{\varepsilon}_t^2$, from the conditional mean inflation equation on a constant and q lags as follows:

$$\hat{\varepsilon}_t^2 = \beta_0 + \sum_{i=1}^q \beta_i \hat{\varepsilon}_{t-i}^2 + v_t \quad (3)$$

The LM procedure tests the joint null hypothesis that the coefficients of all the q lags of the squared residuals are not significantly different from zero. The null hypothesis is rejected if the value of the test statistic (defined as the number of observations multiplied by the coefficient of determination from Equation (3)) is greater than the critical value of the χ^2 -distribution.

The Ljung–Box Q -test is a joint test of whether the first q autocorrelation coefficients of the residuals are jointly zero or not. The Breusch–Godfrey test relies on the null hypothesis that the current residual is uncorrelated with any of its r previous values. The test statistic is $(T - r) \times R^2 \sim \chi^2$, where T is the number of observations, r is the lag length of the serial correlation terms and R^2 is the coefficient of determination obtained from regressing $\hat{\varepsilon}_t$ on all the regressors in Equation (2) plus r lags of $\hat{\varepsilon}_t$. The null hypothesis of no serial correlation is rejected if the test statistic exceeds the critical value from the chi-squared statistical table.

To construct a measure of inflation uncertainty, the mean inflation equation is augmented to incorporate the presence of the time-varying variances in the residuals using the following ARMA-GARCH model in Equations (4) and (5).

$$\pi_t = \alpha_0 + d_{91}D_{91} + \sum_{i=1}^p \alpha_i \pi_{t-i} + \sum_{j=1}^q \beta_j \varepsilon_{t-j} + \varepsilon_t \quad (4)$$

$$\sigma_t^2 = \gamma_0 + d_{91}D_{91} + \sum_{i=1}^p \gamma_i \sigma_{t-i}^2 + \sum_{j=1}^q \eta_j \varepsilon_{t-j}^2 \quad (5)$$

in which, $E(\varepsilon_t | \alpha_{t-1}) = 0$; $\text{var}(\varepsilon_t | \alpha_{t-1}) = \sigma_t^2$

From this model, the estimated conditional variance σ_t^2 is used as a measure for inflation uncertainty. The sum of the coefficients of the ARCH (η) and GARCH (γ) terms in the conditional variance equation determine the persistence of inflation volatility due to inflationary shocks.

One important policy reform that took place during the study period is the Economic Reform and Structural Adjustment Program (ERSAP) in 1991 with price stability as a key objective of the program. This period witnessed exogenous shocks to the Egyptian economy such as the Iranian–Iraqi war in 1988, and the first Gulf war between 1990 and 1991. These shocks led to higher international price for oil, which increased other commodity prices, in addition to the return of millions of Egyptians working in the Gulf region with their savings. Figures 1 and 2 seem to suggest a structural break taking place to the inflation and inflation volatility processes around the year 1991. Though not reported, a Chow break point test was conducted and results confirm the existence of a structural break at 1991.

If there is a structural change to the inflation and inflation volatility process, this has to be controlled for when investigating the causal relation between inflation and its uncertainty to avoid misleading inferences. To account for the effect of the ERSAP, a dummy variable, D_{91} , is included in both the inflation and variance equations, where $D_{91} = 1$ for $t \geq 1991:01$ and zero otherwise.

After constructing a measure of inflation uncertainty, a bivariate vector autoregression model (VAR) is then used to test whether inflation Granger causes inflation uncertainty or *vice versa*, as in Equations (6) and (7):

$$\pi_t = \alpha_0 + d_{91}D_{91} + \sum_{i=1}^k \alpha_{1i} \pi_{t-i} + \sum_{i=1}^k \eta_{1i} \hat{\sigma}_{t-i}^2 + \varepsilon_{1t} \quad (6)$$

$$\hat{\sigma}_t^2 = \alpha_1 + d_{91}D_{91} + \sum_{i=1}^k \alpha_{2i} \hat{\sigma}_{t-i}^2 + \sum_{i=1}^k \eta_{2i} \pi_{t-i} + \varepsilon_{2t} \quad (7)$$

5.3. Robustness Check: Simultaneous Modeling of Inflation and Inflation Uncertainty Relationship

An alternative approach to the two-step procedure, a growing stream of studies has modeled the relationship between inflation and inflation uncertainty simultaneously. Several studies have considered that the simultaneous modeling of inflation and inflation uncertainty relationship is superior to the two-step procedure, given the possibility of a feedback effect between the two variables (see for example: Kontonikas, 2004 [24]; Keskek and Orhan, 2010 [16]). For example, Kontonikas (2004) [24] used symmetric, asymmetric and component GARCH-M models of inflation simultaneously to examine the relationship between inflation and inflation uncertainty in UK over the period 1972–2002.

In the current study, various versions of the Generalized Autoregressive Conditional Heteroskedasticity-in Mean (GARCH-M) model are used to estimate the conditional mean and variance of inflation simultaneously. The (GARCH-M) model augments the conditional mean equation of inflation by adding the conditional variance, or its square root, as a regressor. This allows testing the Friedman–Ball and the Cukierman–Meltzer hypotheses simultaneously, as in Equations (8) and (9):

$$\pi_t = \alpha_0 + d_{91}D_{91} + \sum_{i=1}^p \alpha_i \pi_{t-i} + \sum_{j=1}^q \beta_j \varepsilon_{t-j} + \varphi \sigma_t + \varepsilon_t \quad (8)$$

$$\sigma_t^2 = \gamma_0 + d_{91}D_{91} + \sum_{i=1}^p \gamma_i \sigma_{t-i}^2 + \sum_{j=1}^q \eta_j \varepsilon_{t-j}^2 + \dot{\mu} x_t \quad (9)$$

In Equation (9), x_t and $\dot{\mu}$ denote a vector of n-exogenous variance regressors and their corresponding coefficients. Following what is standard in the literature, x_t contains only the lagged inflation rate (see: Kontonikas, 2004 [24]; and Keskek and Orhan, 2010 [16]). The Cukierman and Meltzer (1986) [8] hypothesis is supported if φ in Equation (8) is statistically significant and has a positive sign, while a positive and statistically significant μ in Equation (9) is supportive of the Friedman–Ball hypothesis. A negative and statistically significant coefficient for σ_t in Equation (8) is supportive of Holland (1995) [10] stabilization hypothesis.

One of the restrictions of the standard GARCH-M model is that it assumes a symmetric response of inflation uncertainty, regardless of the direction of the inflationary shock. To account for possible asymmetries in the response of inflation uncertainty to positive and negative shocks in inflation, several asymmetric versions of the GARCH-M model are estimated. One asymmetric model is the Exponential GARCH model (EGARCH), proposed by Nelson (1991) [33], which uses the following equation for the conditional variance:

$$\ln(\sigma_t^2) = \gamma_0 + \gamma_1 \ln(\sigma_{t-1}^2) + \gamma_2 \left[\frac{\varepsilon_{t-1}}{\sqrt{\sigma_{t-1}^2}} \right] + \gamma_3 \left[\frac{|\varepsilon_{t-1}|}{\sqrt{\sigma_{t-1}^2}} \right] + \dot{\mu} x_t \quad (10)$$

One advantage of the EGARCH model over the symmetric GARCH specification is that there is no need artificially to impose non-negativity constraints on the parameters of the model, since the conditional variance (σ_t^2) is in logarithmic form. According to this model, there is no asymmetric effect if $\gamma_3 = 0$.

Another asymmetric model is the threshold GARCH (TGARCH) model attributable to Glosten *et al.* (1993) [34]. This model extends standard GARCH by adding an additional term to account for possible asymmetries. Accordingly, the conditional variance in Equation (9) is now given by

$$\sigma_t^2 = \gamma_0 + \sum_{i=1}^p \gamma_i \sigma_{t-i}^2 + \sum_{j=1}^q \eta_j \varepsilon_{t-j}^2 + \beta \varepsilon_{t-1}^2 I_{t-1} + \dot{\mu} x_t \quad (11)$$

in Equation (11), I_{t-1} equals 1 if $\varepsilon_{t-1} < 0$ and equals 0 otherwise.

6. Results

6.1. Unit Root Tests

Results of the ADF and PP tests are presented in Table 2. Both tests reject the null hypothesis of a unit root in the inflation series at the 1% significance level. This means that the Egyptian inflation time series is integrated of order zero $I(0)$, which implies that the effect of any shock will die out over time and the inflation series will return to its long-run mean.

Table 2. Unit root tests, Egyptian CPI inflation rate, January 1974–April 2015.

	Augmented Dickey-Fuller (ADF)	Phillips-Perron
Intercept	−25.30 ***	−25.20 ***
Trend and intercept	−25.57 ***	−25.57 ***
No trend and intercept	−3.49 ***	−24.66 ***

*** Indicate rejection of the null-unit root hypothesis at 1% level of significance. The hypotheses of interest are H_0 : inflation series contains a unit root *versus* H_1 : inflation series is stationary. The ADF augments the test using p lags of the dependent variable to ensure that the error terms of the test are not auto correlated. The Schwarz Bayesian Information Criterion (SBIC) is used to determine the optimal lag length of the ADF test.

Results of the Zivot–Andrews and Clemente, Montanes and Reyes unit root tests are presented in Table 3. Zivot–Andrews structural break unit root test show that the inflation time series is stationary at level with a single structural break in October 1991, as the t -statistic is statistically significant, which implies rejection of the null hypothesis of having a unit root with a structural break. Results of Clemente, Montanes and Reyes unit root test show that the inflation time series is stationary at level with two structural breaks under both the Additive Outlier and the Innovative outlier version of the test. According to both the Ao and Io versions, the first structural break T_{B1} , took place in 1985, and the second structural break T_{B2} took place in 1989. The structural breaks identified by the unit root tests coincide with the implementation of the Economic Reform and Structural Adjustment Programme (ERSAP). In 1991, Egypt adopted a battery of reform policies under the ERSAP after consultation with the IMF and the World Bank, to restore the internal as well as the external balance. Macroeconomic stabilization, including inflation reduction, was a key objective of this program. One important policy change that had an implication for inflation rate was the switch in financing budget deficits from monetary expansion to non-inflationary finance measures such as Treasury bill auctions. This was also associated with a drop in the Egyptian budget deficit and a significant reduction in external debts and the associated servicing due to the generous debt relief granted to Egypt after participation in the coalition to free Kuwait in the first Gulf war in 1991.

Table 3. Unit root tests with structural breaks.

Zivot–Andrews Structural Break Trended Unit Root Test		Clemente–Montanes–Reyes Unit Root Test with Double Mean Shifts					
		Additive Outlier			Innovative Outliers		
T-statistic	Time break	T-statistic	T_{B1}	T_{B2}	T-statistic	T_{B1}	T_{B2}
−26.54 ***	1991 m11	−5.591 **	1985m04	1989m02	−27.17 **	1985m05	1989m03

For Zivot–Andrews structural break trended unit root test, the hypotheses of interest are H_0 : the time series has a unit root with a structural break *versus* H_1 : the time series is stationary with a structural break. **, *** indicate rejection of the null hypothesis at 5% and 1% level of significance, respectively. Zivot–Andrews structural break trended unit root test allows for break in both intercept and trend. For the Clemente, Montanes and Reyes test, the hypotheses of interest are H_0 : the time series has a unit root with structural breaks *versus* H_1 : time series is stationary with structural breaks. T_{B1} and T_{B2} are the dates of the structural breaks.

6.2. Results of the Two-Step Approach

The estimated coefficients of the ARMA model, along with the various diagnostic tests, are presented in Table 4. The SBIC and AIC propose modeling the conditional mean equation of inflation (Equation (2)) as an ARMA (4,4) process. Statistics of both the Ljung–Box Q -test and Breusch–Godfrey test show no serial correlation in the residuals at different lag lengths. In particular, the test statistics are statistically insignificant at 1% level of significance for all lags. This is also consistent with the Durbin–Watson test for serial correlation (D.W statistic = 1.994). Results of the Engle’s (LM) test, presented in Table 5, for different lag lengths all reject the null hypothesis of homoskedasticity in the conditional variance of the inflation error terms, implying the existence of ARCH effect.

Table 4. Ordinary Least Squares estimate of inflation conditional mean equation during the period January 1974–April 2015.

Coefficients	Estimated Coefficients
α_0	1.469 ***
d_{91}	−0.785 ***
α_1	−0.765 ***
α_2	0.050
α_3	1.046 ***
α_4	0.598 ***
β_1	0.634 ***
β_2	−0.066 **
β_3	−1.053 ***
β_4	−0.499 **
Diagnostic Tests	Test statistics
Durbin-Watson statistic	1.994
$Q(2)$	0.062
$Q(4)$	3.255
$Q(8)$	3.968
$Q(12)$	5.520
$(T - r) \times R^2 (2)$	0.598
$(T - r) \times R^2 (4)$	3.690
$(T - r) \times R^2 (8)$	4.798
$(T - r) \times R^2 (12)$	6.038

Q , $(T - r) \times R^2$ refer to the Ljung–Box, Breusch–Godfrey test statistics for serial correlation. **, *** indicate statistical significance at 5% and 1% level of significance, respectively.

Table 5. Testing for the existence of ARCH effect.

q	F -Statistic	LM Statistic = $T \times R^2$
1	85.82 ***	73.28 ***
4	22.54 ***	76.74 ***
8	11.97 ***	81.20 ***
12	16.06 ***	140.14 ***

q denotes the order of augmentation of the LM test, T indicates the number of observations and R^2 is the coefficient of determination. *** indicate statistical significance at 1% significance level.

The information criteria proposed estimating an ARMA(4,4)-GARCH(1,1) model. The generated conditional variance is depicted in Figure 3 and the results of the estimated model along with the several diagnostic tests are presented in Table 6. Results show a high degree of persistence in inflation volatility in response to inflationary shocks because the sum of the coefficients for the ARCH and GARCH terms is 0.97, which is close to unity.

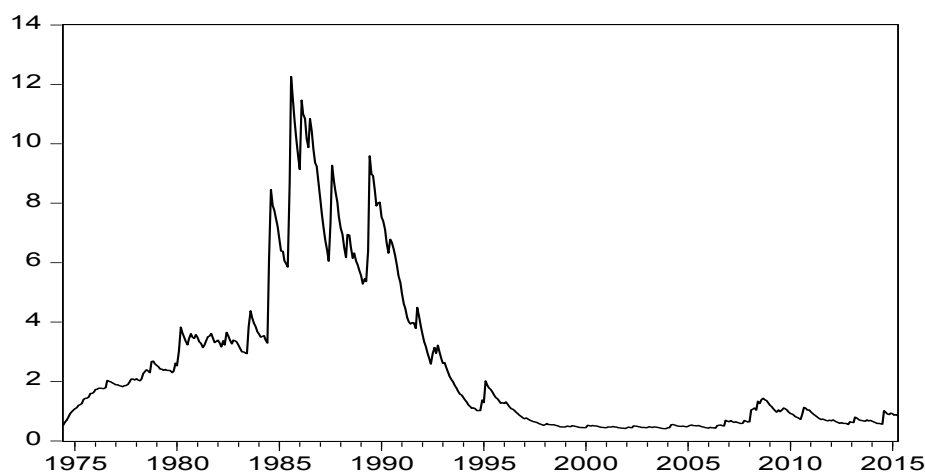


Figure 3. Conditional variance from the ARMA(4,4)-GARCH(1,1) model.

Table 6. OLS estimates of the ARMA(4,4)-GARCH(1,1) model.

Mean Equation	Estimated Coefficients
α_0	1.01 ***
d_{91}	−0.49 ***
α_1	−1.05 ***
α_2	−0.34 ***
α_3	−1.01 ***
α_4	−0.71 ***
β_1	1.11 ***
β_2	0.48 ***
β_3	1.16 ***
β_4	0.80 ***
Variance Equation	
γ_0	0.11 **
d_{91}	−0.09 **
γ_1	0.05 **
η_1	0.92 ***
Diagnostic Tests	
Durbin-Watson statistic	2.36
$Q(1)$	1.04
$Q(4)$	5.30
$Q(8)$	12.25

Q refers to the Ljung–Box test statistics for serial correlation. **, *** indicate statistical significance at 5% and 1% level of significance, respectively.

Results of the Granger causality test, presented in Table 7, show a bi-directional causality (bi-directional feedback) between inflation and inflation uncertainty. In particular, the Granger-causality test rejects the null hypothesis that inflation uncertainty does not Granger-cause inflation across all lag lengths at the one percent level of significance. Meanwhile, the null hypothesis that inflation does not Granger-cause inflation uncertainty is also rejected for all the used lag lengths at one percent level of significance.

Table 7. Granger-causality tests: inflation and inflation uncertainty.

Lag Length	H_0 : Inflation Does not Granger Cause Inflation Uncertainty	H_0 : Inflation Uncertainty Does not Granger Cause Inflation
2	16.13 ***(+)	15.03 ***(+)
4	6.25 ***(+)	9.08 ***(+)
8	5.25 ***(+)	6.04 ***(+)
12	5.10 ***(+)	4.17 ***(+)

The reported figures are the F -statistic. The asterisks *** indicate 1% significance levels. The signs (+) are based on the sum of the estimated coefficients.

6.3. Results of the Simultaneous Modeling Approach

Estimates from the symmetric GARCH-M model, depicted in the first column of Table 8, show a positive and statistically significant effect of inflation uncertainty on the level of inflation. This is reflected by a statistically significant, positive coefficient of σ_t in the mean equation at the 10% level of significance, which supports the Cukierman–Meltzer hypothesis of a positive causal relationship running from inflation uncertainty to the inflation level. For the variance equation, results show a positive effect of inflation level on inflation uncertainty at the 1% level of significance, which supports the Friedman–Ball hypothesis.

Results of the EGARCH and TGARCH models, presented in the second and third columns of Table 8, show that the past inflation level has a statistically significant positive effect on current inflation uncertainty, which supports the Friedman–Ball hypothesis. Meanwhile, inflation uncertainty has a positive and statistically significant effect on the inflation level, which is consistent with the Cukierman–Meltzer hypothesis. These results are in line with the results obtained earlier from the Granger causality test that there is a two-way positive causal relationship between inflation and its uncertainty in Egypt during the study period. Consistency of the findings of the two-step and the simultaneous modeling approaches indicates that the feedback positive relation between inflation level and inflation uncertainty in Egypt during the study period is robust and not sensitive to the way the inflation–inflation uncertainty relationship is modeled.

Table 8. Symmetric and Asymmetric GARCH (p,q)-M models augmented by lagged inflation variables.

Conditional Mean Coefficients	Symmetric GARCH-M Model	Exponential GARCH-M Model	Threshold GARCH-M Model
α_0	1.326 ***	0.466	0.111
d_{91}	−0.730 *	−0.177	−0.046
α_1	−0.592 ***	−0.853 ***	−0.327
α_2	0.135 **	0.004	−0.177
α_3	0.522 ***	0.918 ***	−0.919 ***
α_4	0.803 ***	0.881 ***	−0.087
β_1	0.516 ***	0.843 ***	0.297
β_2	−0.038	0.085 ***	0.265
β_3	−0.438	−0.830 ***	0.947 ***
β_4	−0.835 ***	−0.834 ***	0.034
φ	0.026 *	0.143 **	0.578 ***
Conditional variance coefficients	Symmetric GARCH-M Model	Exponential GARCH-M Model	Threshold GARCH-M Model
γ_0	0.706 ***	0.104 ***	0.103 *
d_{91}	−0.714 ***	−0.050 ***	−0.109 *
γ_1	0.553 ***	0.974 ***	0.900 ***
γ_2		0.091 ***	
γ_3		−0.104 ***	
η_1	0.195 ***		0.024
β			0.037
μ	0.393 ***	0.013 ***	0.087 **
Diagnostic tests	Symmetric GARCH-M Model	Exponential GARCH-M Model	Threshold GARCH-M Model
Durbin-Watson Statistic	2.124	2.23	2.25
$Q(1)$	0.197	0.637	0.580
$Q(4)$	0.473	1.056	2.701
$Q(8)$	6.529	1.957	11.95

p, q represent the order of the GARCH, ARCH term, respectively. Q denote the Ljung–Box test statistic for residual serial correlation. *, **, *** indicate statistical significance at the 10%, 5% and 1% level, respectively.

7. Conclusions and Policy Implications

This paper has investigated the linkage between inflation level and inflation uncertainty in Egypt during the period January 1974–April 2015. As a baseline specification, the inflation-uncertainty relation is modeled using the two-step procedure. The estimated conditional variance from an ARMA-GARCH model is used as a measure of inflation uncertainty, and a Granger-causality test is conducted to determine the causal relation between the two variables. To control for any potential feedback effects between inflation and inflation uncertainty, a simultaneous estimation approach is also used as a robustness check for the results. Results show a high degree of volatility persistence in response to inflationary shocks.

Results also showed that the stabilization measures adopted under the ERSAP have helped reduce inflation levels and inflation volatility as the coefficients of the dummy variable that accounts for the effect of the ERSAP were negative.

Granger-causality and asymmetric GARCH-M, EGARCH-M and TGARCH-M models indicate a statistically significant positive, two-way relationship between inflation and inflation uncertainty, supporting both the Friedman–Ball and the Cukierman–Meltzer hypotheses. This has important implications for the relationship between inflation and output given the substantial empirical evidence that greater inflation uncertainty is detrimental to economic growth.

The feedback relationship between inflation and inflation uncertainty that is found in this study has imperative implications for the monetary policy in Egypt. Since the inflationary expectations have significant effect on inflation, monetary authority in Egypt should enhance the credibility of monetary policy and attempt to control inflation uncertainty by restrictive monetary policy measures. One policy measure is establishing a nominal anchor for the monetary policy to stabilize inflation expectation and hence lower inflation rates. In addition, previous studies showed that structural reforms based on improving Egypt's productive capacity, reducing the budget deficit and national debts, are crucial for controlling inflation (El-Sakka and Ghali, 2005 [35]). Keskek and Orhan (2010) [16] found strong evidence, using data from Turkey, that inflation-oriented monetary policy effectively reduces the inflation persistence and eliminates uncertainty.

Given the substantial real costs of inflation and inflation uncertainty, the findings of this paper support the view of adopting inflation control target policy in Egypt. Several studies have found that inflation targeting, once its preconditions are satisfied, contributes to the reduction of inflation and its associated volatility. For example, in a panel study of twenty-five countries, Tas and Ertugrul (2013) [36] found evidence that the adoption of inflation targeting helped most of the countries to achieve lower inflation uncertainty. In another study, Lin and Ye (2009) [37] examined the effect of implementing inflation targeting in thirteen developing countries and found that inflation targeting has large and significant effects on lowering both inflation and inflation uncertainty. Similar evidence was found by Payne (2009) [27] that inflation targeting reduced inflation uncertainty in response to inflationary shocks in Thailand. Kontonikas (2004) [24] also found that the adoption of inflation target in U.K. since 1992 has helped eliminate inflation persistence and reduced long-run uncertainty. Monetary authorities in Egypt could benefit from the experience of those countries, which succeeded in reducing inflation and its associated uncertainty.

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Conflicts of Interest

The author declares no conflict of interest.

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