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Inflation Targeting or Exchange Rate Targeting: Which Framework Supports The Goal of Price Stability in Emerging Market Economies?*

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Abstract

We investigate the relationship between inflation and inflation uncertainty under inflation targeting and a conventional fixed exchange rate system and the impact of each regime on inflation and inflation uncertainty over the span from 1980:01 to 2014:06. The results from GARCH in mean models reveal that, under the two monetary regimes, inflation increases inflation uncertainty and inflation uncertainty raises inflation. This positive bi-directional relationship between inflation and inflation uncertainty provides evidence of the importance of non-discretionary monetary policies. Both regimes appear effective in reducing inflation uncertainty in the long-run which suggests the importance of monetary regimes as signalling devices for inflation expectations. The fixed exchange rate regime has no impact on average inflation and inflation inertia, while inflation targeting has been successful at lowering average inflation and inflation persistence of its adopters. Nevertheless, the results provide evidence that inflation targeting countries have not benefited equally from this monetary regime.

Key Words: Inflation Targeting, Fixed Exchange Rate System, GARCH, Monetary Policy, Price Stability.
JEL Classification: C54, C58, E50, E58

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1 Introduction

Over the last century, monetary regimes that anchor the exchange rate have failed to ensure the goal of price stability. Indeed, the success and the prolonged use of exchange rate to fix prices have been questioned after the failure of past fixed exchange rate systems. As a consequence, many countries have moved to adopt different forms of Inflation Targeting (IT).Nonetheless, there is a consensus among economists that the social and financial costs attributed to inflation are due mainly to uncertainty about future inflation. In fact, the links between inflation and inflation uncertainty have gained attention in the literature after the Nobel lecture of Friedman (1977); he argues that inflation causes higher inflation uncertainty. The positive relationship between the two variables is theoretically justified by Ball (1992), who shows that weak policy makers are more likely to permit inflation during high-inflation episodes, generating more inflation uncertainty. Cukierman and Meltzer (1986), on the other hand, reveal that inflation uncertainty leads to higher inflation. However, it is also argued that the causality from inflation uncertainty to inflation might turn negative. In this sense, Holland (1995) states that central banks could respond to inflation uncertainty through lowering the money stock, and thereby, a negative nexus might appear as a sign of stabilising reaction of monetary policy. Consequently, over the last three decades, questioning the relationship between inflation and inflation uncertainty has encouraged a large number of empirical studies, which attempted to assess the validity of Friedman-Ball (F-B) and Cukierman and Meltzer (C-M) hypotheses. Following the development of the Auto Regressive Conditional Heteroskedasticity (ARCH) and Generalised ARCH (GARCH), by Engle (1982) and Bollerslev (1986), respectively, more studies have been conducted to examine the links between inflation and inflation uncertainty. However, while there has been a considerable number of studies on the relationship between inflation and inflation uncertainty, there have been a few studies, e.g., Chang and He (2010), Kontonikas (2004), Caporale et al. (2012), that have considered the role of monetary regimes on the nexus, but the attention has been given to IT and the euro regime. In addition, to the best of our knowledge, no study, other than Khan et al. (2013), has attempted to compare between two quantitatively-based monetary regimes. Khan et al. (2013) studied the relationship between inflation and inflation uncertainty for Eastern European countries which have a currency board or are inflation targeters. However, they fail to address distinctions between the two regimes, and thus, to specify which regime works better at reducing the nominal and real inflation uncertainty. Furthermore, far too little attention has been paid to the cases with soft fixed exchange rates. Even when some cases were considered in the studies, e.g., Daal et al. (2005) and Samimi et al. (2012), researchers have not investigated the relationship between inflation and uncertainty in much detail, i.e., the impact of monetary regimes on the nexus has been ignored.

Hence, with a similar comparison objective of Khan et al. (2013), this paper attempts to fill a gap in the empirical literature by investigating the nexus between inflation and inflation uncertainty in emerging market economies under two monetary anchors: a fixed exchange rate (FER) regime to the US dollar and inflation targeting. The aim of this paper is twofold. Firstly, we test the F-B and C-M hypotheses under the two regimes. Secondly, we evaluate the plausible effects of adopting a specific quantitative target on inflation uncertainty.

Different GARCH in Mean (GARCH-M) models are constructed to investigate the relationship in two countries with fixed exchange rate to the US dollar (FER);

Jordan and Egypt, and three inflation targeters: South Africa, Brazil and Poland.[1] We select inflation targeters, from three different continents, that had experienced an economic and/or political challenge before shifting to inflation targeting. The exchange rate targeting countries represent cases which also experienced a shift in monetary regime during the study period and suffered from political and economic pressures, criteria which do not apply to most exchange rate peggers, e.g., Gulf States. This selection allows us to highlight how the economies of the examined countries have benefited from the regime under investigation. The paper compares implicitly between the impact of monetary policy independence on inflation and inflation uncertainty, where inflation targeting enjoys more monetary policy flexibility compared to the soft pegged exchange rate system.

The plan of this paper is as follows; section two provides the literature review. Section three discusses the methodology applied. Section four presents the data. The results and concluding remarks are provided in section five and six, respectively.

2 Literature Review

Some policy makers and economists suggest that the costs of a predicted low and moderate inflation rate are acceptable and supported by economic theory. However, Okun (1971) shows that an anticipated rate of steady inflation as implied by accelerationists would be ideal to wind down the social and redistributive costs of inflation. However, such steady inflation is difficult to achieve due to inflation expectations, which hinge substantially upon the type of government in power and the trade-off between employment and inflation. Consequently, he points out that the acceptance of moderate and steady inflation would trigger higher inflation expectations, which eventually leads to a higher inflation rate. He hypothesises, by analysing the inflation behaviour for different OECD countries, that high inflation may lead to higher inflation variability, and that high inflation countries experienced higher inflation variability. Nevertheless, the link between inflation and its uncertainty has gained much interest after the Nobel lecture of Friedman (1977). Friedman states that the relationship between unemployment and nominal wage changes is not stable owning to inflation uncertainty, which increases with the level of inflation.[2] Ball (1992) supports the hypothesis suggested by Friedman that high inflation leads to higher inflation uncertainty. He bases his argument on a monetary policy-time inconsistency model of Barro and Gordon (1983), where market agents are uncertain about the type of government in power. On the other hand, on the basis of the same time inconsistency model, Cukierman and Meltzer (1986) argue that, as central bankers are motivated to create surprise inflation to stimulate economic activity, an increase in inflation variability raises the level of inflation rate.[3] In other words, for Friedman and Ball, higher inflation creates higher inflation uncertainty, while

[1] Note that Egypt has opted out from the fixed exchange rate system, so the study compares between the time during the fixed system and after abandoning it.

[2] He further argues that inflation-inflation uncertainty leads to lower output growth; however, this impact is unsettled in the literature and depends on whether money is considered neutral; see e.g., Tobin (1965), Sidrauski (1967), Stockman (1981). However, the most recent studies have shown that the link exists in practice, but the effect of inflation on economic growth comes through inflation uncertainty, see Chang and He (2010), Grier et al. (2004). Fountas (2010), on the other hand, finds that output growth is not affected by inflation uncertainty.

[3] Policy makers may increase an optimal inflation rate to benefit from low unemployment, see Tobin (1972), or to lower the public debt, see Fischer and Summers (1989) for further detail.

for Cukieman-Meltzer, the causality goes in the other direction, that is, inflation uncertainty increases inflation.

Therefore, a large number of empirical works has attempted to examine the nexus between inflation and inflation uncertainty using different measures of inflation uncertainty, finding inconclusive results. See Glezakos and Nugent (1984), Pourgerami and Maskus (1987) and Cukierman and Wachtel (Cukierman and Wachtel) amongst many others.

Although the survey-based measures were believed to be good proxies for inflation uncertainty, such measures were unable to distinguish between transitory and permanent shocks to inflation, where the latter have a much stronger effects on the intertemporal decision making of individuals and businesses. Thus, Ball et al. (1990) claim that the effect of control errors lasts temporarily and decays over short periods, whereas inflation has a severe effect on uncertainty at longer horizons, where permanent shocks dominate. Engle (1982) was the first to measure inflation uncertainty as the conditional variance of inflation to study the relationship between inflation and uncertainty in the United States. In fact, the introduction of Autoregressive Conditional Heteroscedasicity (ARCH) and General Autoregressive Conditional Heteroscedasicity (GARCH) approaches by Engle (1982) and Bollerslev (1986), respectively, encourages a large number of recent empirical work examining the link between inflation and inflation uncertainty.

Evans (1991) constructs a model which allows for the changes in the structure of inflation to affect inflation uncertainty. This is performed by incorporating the different aspects of inflation uncertainty through the Kalman filter: the conditional variance of inflation, the conditional variance of expected inflation and the conditional variance of steady-state inflation. He applies the model to the US during the period 1960:01-1988:06 and concludes that inflation raises inflation uncertainty. Grier and Perry (1998) examine the relationship utilising a GARCH model to produce a proxy for inflation uncertainty in the G7 countries. Their findings from the Granger causality tests suggest that the nexus is positive as implied by F-B, but little evidence is found in favour of the C-M hypothesis. Fountas (2001) and Kontonikas (2004), who apply GARCH and GARCH-M, respectively, find evidence that the F-B hypothesis, for different examined periods, hold true for the UK. In fact, most studies have used either GARCH or GARCH-M to investigate the relationship between inflation and uncertainty in different countries, reporting mixed results concerning the causality between inflation and uncertainty; whether it goes from inflation to uncertainty or the opposite; others even find the link to be bi-directional. Conrad and Karanasos (2005) find evidence for the Friedman hypothesis in the US and the UK, while the results for Japan show that uncertainty affected the level of inflation as suggested by C-M.

Most of the studies on the link between inflation-inflation uncertainty, e.g., Daal et al. (2005), Grier and Grier (2006), Hwang (2001), Thornton (2007), Payne (2008), Keskek and Orhan (2010), Jiranyakul and Opiela (2011), Hartmann and Herwartz (2012), Fountas and Karanasos (2007), support the Friedman hypothesis. Similarly, Grier and Grier (2006), who apply an augmented multivariate GARCH-M to study the nexus in Mexico, find that high inflation increases inflation uncertainty. Nevertheless, Hwang (2001) finds, using an ARFIMA-GARCH model, no causal nexus between inflation and its uncertainty in the United States over the period 1926-1992.

One drawback of GARCH models is that it is unable by construction to capture the asymmetric responses to positive and negative shocks to inflation. Hence a new family of asymmetric GARCH has evolved to consider the fact that bad news in financial markets has deeper effects than good news (Grier and Perry, 1998). Subsequently, many studies have constructed different models of asymmetric GARCH, such as Jiranyakul and Opiela (2010), Nazar et al. (2010) and Rizvi and Naqvi (2009) amongst others. In fact, the adoption of inflation targeting across many developed and developing central banks has increased the willingness to discover the benefit of the new framework. Fountas (2001) suggests that the announcement of an explicit inflation target has a prominent effect on lowering inflation persistence and uncertainty at long horizon. Kim (1993) also supports the Friedman hypothesis for the United States as the nexus is found to be positively correlated during high inflation periods. Likewise, Tas and Ertugrul (2013) find that the relationship between inflation and inflation uncertainty is positive and inflation variance has decreased after IT in most inflation targeters investigated in his study. Similar findings were obtained by Kontonikas (2004) for the UK. Bhar and Mallik (2013) also point out that IT is an optimal anchor for the UK, however, they find that the relationship between inflation and uncertainty has turned to be negative after inflation targeting. Furthermore, Bhar and Mallik (2012) and Bhar and Mallik (2013) claim that adopting IT in New Zealand and Australia grants monetary authorities more flexibility in setting the nominal interest rates.[4] Other studies such as Neanidis and Savva (2011), Caporale and Kontonikas (2009), Caporale et al. (2012), show that adopting the euro played a major positive role in affecting the nexus between inflation and uncertainty. The only study which compared between two different monetary regimes; currency boards and inflation targeting, for Eastern European countries, is Khan et al. (2013), who apply EGARCH model and find support for the Friedman hypothesis. However, the study fails to underline which monetary anchor worked better at reducing inflation uncertainty.

The current paper examines the relationship between inflation and inflation uncertainty in inflation targeting and exchange rate targeting countries and highlights the role of each monetary anchor in reducing the inflation uncertainty. In particular, the paper attempts to underline the impact of monetary policy flexibility on inflation and inflation uncertainty using GARCH models. Hence, it adds to the growing body of literature on the importance of inflation targeting as a framework for monetary policy.

3 Methodology

GARCH-type models have been widely employed to investigate the relationship between inflation and inflation uncertainty, as they provide a time-varying measure for volatility.Nevertheless, a standard GARCH model does not allow examining the effect of inflation and inflation uncertainty simultaneously. Hence, the previous statistical technique to study the direction of the nexus, conducted by some studies, was a two stage approach, under which the conditional variance of inflation is estimated in the first stage, then the Granger causality tests are performed to determine

[4] See also Mallik and Bhar (2011) on the relationship between inflation uncertainty and interest rates for five inflation targeters, and Wright (2011) for the effect of inflation targeting on interest rates.

the causal direction of the relationship.[5]

However, GARCH-M, developed by Engle et al. (1987), permits inflation to be specified by inflation uncertainty. Therefore, to investigate the relationship simultaneously, we allow the conditional variance of inflation to be influenced by the mean, and the inflation rate to be determined by the conditional variance. These specifications can be represented as follows:

The mean equation:[6]

$$\pi_t = \alpha_0 + \sum_{t=1}^{p} \alpha_i \pi_{t-i} + \sum_{j=1}^{q} \beta j \varepsilon_{t-j} + \delta \sqrt{h_t} + \varepsilon_t \tag{1}$$

The variance equation:[7]

$$\pi_t = \phi + \sum_{t=1}^{p} \alpha_i \varepsilon_{t-1}^2 + \sum_{j=1}^{q} \beta j h_{t-j} + \lambda Z_{t-1} \tag{2}$$

Where π_{t-i} is the lagged inflation rate; ε_{t-j} is the lagged errors and ε_t is the error term, which has conditional and unconditional mean of zero and conditional variance, h_t, given by equation (2). The conditional variance is determined by the lagged squared residuals, the lagged conditional variance and Z_{t-1}, which includes only lags of inflation. Stationarity restrictions of the model entail that α_i and βj, the non-negative parameters, must be less that unity. If the sum of the parameters is equal to one, the conditional variance must be modelled by Integrated GARCH (Harvey, 2013).

If δ in the mean equation is significantly positive, higher inflation uncertainty generates higher inflation, as argued by C-M. On the other hand, when the coefficient is significantly negative, the Holland hypothesis of monetary policy stabilising effect holds true. λ in equation (2) determines the effect of inflation on inflation uncertainty. Obtaining a positive and significant coefficient indicates that inflation uncertainty increases with inflation, as suggested by F-B.

Nevertheless, the conditional variance of inflation estimated by GARCH-M is formed to consider only the magnitude of inflation shocks, ε_{t-1}^2, and thereby the sign of innovations is ignored by the model construction. Hence, to account for possible asymmetric responses to positive and negative inflation shocks, an asymmetric GARCH model, i.e., Exponential GARCH, is used.

The conditional variance in the EGARCH model, put forward by Nelson (1991), is set in an logarithmic form, which does not require imposing artificially non-

[5]See Nas and Perry (2000) and Grier and Perry (1998). Fountas (2010) incorporates 'in mean coefficients' to capture the effect of inflation uncertainty on inflation. When the effect is found, for some countries, insignificant, he estimates the standard GARCH model, and from the estimated conditional variance he performs Granger causality tests. In the mean equation, he includes lags of inflation uncertainty to consider the inertia of money supply in responding to inflation. The influence of money supply was suggested by Holland (1995), who argues that central banks will respond to inflation uncertainty by reducing money supply and, consequently, decreasing the inflation rate.

[6]In our study, as explained in details in the results section, the mean equation is built on AR specifications rather than ARMA process, so the lagged error terms are excluded.

[7]GARCH model estimates a time-varying variance of residuals which acts as a proxy for unexpected inflation volatility (Daal et al., 2005). The GARCH regression model is built on an autoregressive moving average of a known variable, where the conditional variance is a linear function of its past values and past squared shocks. This model has an ARMA construction which makes the model soluble; however, given its quadratic specification, the model equalises the effect of positive and negative innovations (Zakoian, 1994). Inserting a one-period lagged inflation in the variance equation allows examining the hypothesis of Friedman (1977) and Ball (1992).

negativity constraints on the parameters to ensure a positive variance. The model representation can be seen as follows:

$$\log h_t^2 = \varphi + \beta_1[|\frac{\varepsilon_{t-1}}{h_{t-1}}|] + \beta_2[\frac{\varepsilon_{t-1}}{h_{t-1}}] + \beta_3 \log h_{t-1}^2 + \lambda Z_{t-1} \tag{3}$$

In this case, an asymmetric response to inflation shocks exists if $\beta_2 \neq 0$. A significantly positive β_2 implies that the inflation uncertainty increases more when the economy is hit by a positive shock, i.e., $\varepsilon_{t-1} > 0$, than a negative inflation shock, i.e., $\varepsilon_{t-1} < 0$ (Wilson, 2006).

Nevertheless, as policy makers are more concerned about the long-run impact of inflation uncertainty, and more importantly the impact of monetary anchor on reducing inflation uncertainty in the long-run, we utilise the Component GARCH (CGARCH), developed by Lee and Engle (1993). This model separates the long-run from short-run components of inflation uncertainty by allowing the mean of the conditional variance to vary around a time varying level, φ.

$$h_t = \varphi_t + \alpha_1(e_{t-1}^2 - \varphi_{t-1}) + \beta_1(h_{t-1} + \varphi_{t-1}) \tag{4}$$

$$h_t = \varphi + \rho\varphi_{t-1} + \mu(e_{t-1}^2 - h_{t-1}) + \lambda Z_{t-1} \tag{5}$$

Equation (4) represents the transitory component, which approaches zero with the power of $\alpha_1 + \beta_1$. ρ in the long run component, shown in equation (5), is usually close to one, as the time varying trend converges to the mean very slowly. If, $1 > \rho > \alpha_1 + \beta_1$, the short run component of inflation uncertainty will die out rapidly more than the trend. This indicates that the forecasts of the conditional variance will depend essentially upon the trend (Kontonikas, 2004).

4 Data

Monthly data on Consumer Price Index (CPI) for the period from 1980:01 to 2014:06 are extracted from the International Financial Statistics of the International Monetary Fund for the two FER countries: Jordan and Egypt, and the three inflation targeters: South Africa, Brazil and Poland. The monetary regime shift experienced by all the countries allows to spot the benefits of the examined regime in terms of inflation and inflation uncertainty.

Inflation is computed as π_t = log cpi$_t$ - log cpi$_{t-1}$. The inflation series is then adjusted to remove the seasonality by executing the Census Bureaus X12 in an additive default mode. Table 1 shows the summary statistics of the inflation series properties for each respective country. The statistics of the average inflation, displayed in Table 1-Panel(a), reveal that inflation targeting countries have experienced a marked lower average inflation after shifting to IT. However, this direct effect of the monetary anchor on inflation process is not obvious for the two fixed exchange rate targeters.[8]

The statistics, reported in Table 1-Panel(b), indicate that the distribution of inflation is heavy-tailed. For all the countries, the distribution is positively skewed and leptokurtic, and inflation series failed to satisfy Jarque and Bera (1980) test for normality. This asymmetry and peakness of the distribution are considered in modelling the relationship between inflation and inflation uncertainty by using

[8] This cannot be attributed to the recent economic and political disturbances in the Middle East region. The two countries have been reeling from unstable political and social conditions for the full period of the study.

a generalised error distribution with normalised density of zero mean and unity variance, in which the normal distribution is a special case (Nelson, 1991).

The seasonally adjusted inflation series for each country exhibits stationary process, as shown in Table 2. According to Dickey and Fuller (1979) (ADF) and Phillips and Perron (1988) (PP) tests, with and without a trend, the null hypothesis that inflation contains a unit root is rejected at a high level of significance.

Breusch and Godfrey test for serial correlation is first executed to ensure the whiteness of residuals before testing for ARCH effect. The results, presented in Table 1-Panel(b), imply that the inflation series' residuals are serially independent and conditionally heteroskedastic. A constant and slope regime dummy variable are imposed in the mean equation to account for the shift in monetary regime. We also account for the effect of certain political events in some countries on inflation and inflation uncertainty like Arab Spring in Jordan and Egypt and Apartheid in South Africa.

5 Results

The conditional mean of inflation is specified by constructing several ARMA models. Given that inflation is seasonally adjusted, it is found that including only the autoregressive terms yields the best model specifications, which is the common case in modelling inflation in the empirical literature (Kontonikas, 2004). For each country, we begin by incorporating up to twelve AR specifications to allow capturing the persistence of the data. The length of AR components is shortened on the basis of Akaike and Schwartz information criteria and by ensuring that all autocorrelation coefficients up to twelve lags fall inside the non-rejection region, which is also confirmed by the Q-statistics of Box and Pierce (1970). Accordingly, the selected AR process forms the following benchmark-mean specifications:

Jordan:

$$\pi_t = \gamma_0^{JO} + \gamma_1^{JO}\pi_{t-2} + \gamma_2^{JO}\pi_{t-5} + \gamma_3^{JO}\pi_{t-9} + \gamma_4^{JO}\pi_{t-12} + u_t \tag{6}$$

Egypt:

$$\pi_t = \gamma_0^{EG} + \gamma_1^{EG}\pi_{t-1} + \gamma_2^{EG}\pi_{t-9} + \gamma_3^{EG}\pi_{t-12} + u_t \tag{7}$$

South Africa:

$$\begin{aligned}\pi_t = \gamma_0^{SA} + \gamma_1^{SA}\pi_{t-1} + \gamma_2^{SA}\pi_{t-2} + \gamma_3^{SA}\pi_{t-3} + \gamma_4^{SA}\pi_{t-7} \\ +\gamma_5^{SA}\pi_{t-8} + \gamma_6^{SA}\pi_{t-11} + \gamma_7^{SA}\pi_{t-12} + u_t\end{aligned} \tag{8}$$

Brazil:

$$\pi_t = \gamma_0^{BR} + \gamma_1^{BR}\pi_{t-1} + \gamma_2^{BR}\pi_{t-2} + \gamma_3^{BR}\pi_{t-8} + u_t \tag{9}$$

Poland:

$$\pi_t = \gamma_0^{PO} + \gamma_1^{PO}\pi_{t-1} + \gamma_2^{PO}\pi_{t-2} + \gamma_3^{PO}\pi_{t-5} + \gamma_4^{PO}\pi_{t-9} + \gamma_5^{PO}\pi_{t-11} + u_t \tag{10}$$

For each country, we split the inflation series between the time before and after adopting the monetary regime of interest. The preliminary evidence from the OLS regression of the benchmark models, shown in Tables 3 to 7, suggests that IT has been successful at reducing the volatility of inflation as ARCH effect turned insignificant after adopting IT. Interestingly, for Egypt, inflation volatility has become lower not during the fixed exchange rate system but after opting out, while for Jordan, the impact of the exchange rate targeting is unclear as the volatility for the period after exchange rate targeting is found at high lags.

For our motivation to examine the simultaneous relationship between inflation and inflation uncertainty, we incorporate the standard deviation in the mean equation as a volatility measure and augment the variance equation with lagged inflation.[9] Furthermore, to model the impact of monetary regime on inflation dynamics, slope dummies are plugged in the conditional mean equations, in which the dummy takes the value of one when the examined monetary regime is in effect, and zero otherwise. We first attempt to introduce the regime slope dummies via different lags, but only two interactive dummies are selected to interact with their corresponding inflation lags based upon a significant improvement in the fit of the model. We also employ a constant regime dummy in the mean equation for the cases where doing so is found to substantially improve the overall statistical performance. In addition, we account for political circumstances in some countries. For the two exchange rate targeters, a dummy variable is added to the mean equation to capture the impact of the Arab Spring on average inflation, in which the dummy is assigned one for the period from 2011:01 onwards.[10] For South Africa, a constant dummy is included to consider the effect of apartheid on inflation, which takes the value of one for the period before January 1995, and zero for the months after. Hence, the augmented mean equations can be represented as follows:

Jordan:

$$\pi_t = (\delta_1^{JO} + \delta_3^{JO})\pi_{t-2} + \delta_4^{JO}\pi_{t-5} + \delta_5^{JO}\pi_{t-9} + (\gamma_2^{JO} + \delta_6^{JO})\pi_{t-12} + u_t \tag{11}$$

Egypt:

$$\pi_t = (\delta_1^{EG} + \delta_3^{EG})\pi_{t-1} + \delta_4^{EG}\pi_{t-9} + (\gamma_2^{EG} + \delta_5^{EG})\pi_{t-12} + u_t \tag{12}$$

South Africa:

$$\begin{aligned}\pi_t = (\delta_1^{SA} + \delta_3^{SA})\pi_{t-1} + \delta_4^{SA}\pi_{t-2} + \delta_5^{SA}\pi t - 3 + \delta_6^{SA}\pi_{t-7} \\ + \delta_7^{SA}\pi_{t-8} + \delta_8^{SA}\pi_{t-11} + (\delta_2^{SA} + \delta_9^{SA})\pi_{t-12} + u_t\end{aligned} \tag{13}$$

Brazil:

$$\pi_t = (\delta_1^{BR} + \delta_3^{BR})\pi_{t-1} + \delta_4^{BR}\pi_{t-2} + (\delta_2^{BR} + \delta_5^{BR})\pi_{t-8} + u_t \tag{14}$$

Poland:

$$\begin{aligned}\pi_t = (\delta_1^{PO} + \delta_3^{PO})\pi_{t-1} + \delta_4^{PO}\pi_{t-2} + (\delta_2^{PO} \\ + \delta_5^{PO})\pi_{t-5} + \delta_6^{PO}\pi_{t-9} + \delta_7^{PO}\pi_{t-11} + u_t\end{aligned} \tag{15}$$

A joint significance of the interactive regime dummies is confirmed by a Wald test of $\delta_1 = \delta_2 = 0$. For each country, Chi-square statistics reject the hypothesis that

[9] Inserting S.D in the mean equation is used by Baillie et al. (1996) and Kontonikas (2004).

[10] The unrest and tensions across the Middle East spread to Jordan and had negative effects on the economy. For instance, the pipelines that carried gas from Egypt to Jordan were targeted and bombed several times during the uprising, resulting in oil supply shortage. As a consequence, Jordan was forced to deal with Israel to import Gas, as Israel has become a major gas exporter in the region. However, the pace to deal with Israel triggered more domestic opposition and increased the external debt; see the Daily Mail on 11^{th} December 2014 for more details. The Arab Spring countries are still, at the time of writing this, being affected by the adverse consequences of the social unrest.

the dummies are zero at 1% level of significance, as shown in the last row of Tables 8 to 14. The effect of the monetary regime on inflation inertia is reflected by the sum of the coefficients of the regime interactive dummy and that of their corresponding lags. A negative slope dummy indicates that inflation persistence has declined after adopting the examined regime.[11] The results, reported in Tables 8 to 14, imply that IT has been successful at reducing inflation persistence at a high lag order, as the coefficients of the second interactive dummies, i.e., δ_2, appear with a negative sign.[12] This, however, does not apply to Poland, where all the slope dummies are non-negative, but its regime constant dummy, D_t, plugged in the mean equation, shows that the mean of inflation was reduced by IT; this also applies to all the ITers in the sample.[13] For South Africa, the years of apartheid were associated with higher average inflation, as the constant dummy, APART, presented in Table 12, is significantly positive under all GARCH models. For Egypt and Jordan, the Arab Spring dummy is found insignificant as a constant, but its slope, POL, reported in Tables 8 and 10, has a positive and significant influence on the trend at 1%.[14] Unlike the ITers, the inflation mean of the FER targeters is not affected by the fixed exchange rate regime. The dummy appears insignificant for Jordan and positive for Egypt. Nonetheless, the FER system appears to be influential in lowering the inflation inertia in Egypt and Jordan at the first inflation lag.

The parameter estimates of the inflation uncertainty proxy, incorporated in the mean equation, have a significantly positive sign for all the countries, indicating that inflation uncertainty increases inflation, as argued by C-M. On the other hand, we find support for the F-B hypothesis; inflation does generate inflation uncertainty, irrespective of the regime followed. Remarkably, the coefficient of the inflation regime slope dummy, λ_1, employed in the variance equation, is significant and negative in all the respective countries; however, the magnitude of the effect is almost negligible for Egypt.[15] Nevertheless, the significance proves that both regimes could to a certain extent reduce inflation uncertainty, albeit not directly via decreasing the mean of inflation for FER targeters.

The results of asymmetric responses to increasing or falling in inflation are not the expected ones for all the cases. For Jordan, the results imply that both negative and positive inflation shocks have the same influence on inflation uncertainty. The positive and significant β_2 for Egypt, presented in the second column, EGARCH, of Table 10, indicates that positive shocks trigger more conditional inflation uncertainty than negative shocks. Similarly, for Brazil, the asymmetric coefficient, presented in

[11]Note that, as stated by Kontonikas (2004), the effect of inflation regime on the inflation persistence is preferred to be analysed in the context of the Kalman filter.

[12]A large number of literature has pointed out that IT helps countries with their disinflationary efforts, stabilises inflationary expectations and enhances the monetary authority's accountability and transparency, see for example Batini et al. (2007), Capistrán and Ramos-Francia (2010), Neumann and Von Hagen (2002), Bernanke and Mishkin (1997), Hu (2003), Walsh (2009), Batini et al. (2007).

[13]We did not report the results with a constant dummy for the other ITers as the results were found to be better off, in terms of diagnostics, without adding the regime constant. However, the constant dummy appeared with a negative sign for all the IT cases. We also attempted to incorporate the regime dummy variable in the variance equations, but the dummy was insignificant for all the cases.

[14]Note that the ARCH effect exists in the estimated EGARCH-M model for Egypt, indicating that the model is not well specified. In general, for both countries, all GARCH-M models showed better results when the POL-slope dummy and the regime constant dummy were dropped from the mean equation, see Table 8 and 10.

[15]Note that this ignores the models estimated for Egypt and Jordan with slope political dummy and regime constant dummy.

Tables 13, suggests that inflation uncertainty increases following a positive inflation shock. However, no asymmetry is found for the two other ITers, implying that inflation uncertainty process is not influenced by the direction of inflation shocks. Interestingly, the slope regime dummy, λ_1, incorporated in the mean equation, remains negative for all the countries after controlling for asymmetries. Even when CGARCH-M is used in modelling the inflation variance, the slope dummy has a significantly negative sign for all the countries, except for South Africa, where the dummy turns insignificant. This finding suggests that FER system and IT alike are effective in reducing inflation uncertainty in the long run. Generally, the inflation trend of the ITers approaches the mean quicker than that of the FER targeting countries. The power of the short-run component of inflation uncertainty is also higher in the FER targeters. The diagnostic statistics, reported below each estimated GARCH model, from Table 8 to Table 14, indicate that the GARCH models are well-specified. The 1^{st} and 12^{th} lag order of Ljung-Box and the 12^{th} lag squared residuals as well as the LM test for ARCH suggest neither remaining autocorrelation nor a non-constant variance for all the countries, except for Poland, where ARCH effect remains in the error terms.[16]

The findings of a positive bi-directional relationship between inflation and inflation uncertainty suggest the need for a monetary anchor to reduce both inflation and inflation uncertainty. In general, inflation targeting countries enjoy lower average inflation and inflation persistence compared to the time before adopting IT and to the countries with FER regime. The two monetary regimes appear effective in reducing inflation uncertainty, even under the presence of asymmetries in some cases, like Brazil and Egypt. The effect also remains in the long run, except for South Africa.

For FER countries, the benefits of the regime are not reflected in lower average inflation and inflation inertia and a stable volatility. The constant regime dummy variable, incorporated in the mean equation, is found to be insignificant for Jordan and positive for Egypt, and although the FER system appears effective in reducing uncertain inflation, the magnitude of the regime dummy coefficient is close to zero for Egypt. It could be argued that anchoring the exchange rate could still influence inflation uncertainty as long as the possibility to renege on FER commitment is not perceived by the market. The several depreciations in the Egyptian Pound affected the mechanism of the FER system and its credibility in the market. The weak economic institutions and dependency on political authority due to the absence of mutual and clear vision between the central bank and government might hamper the role of FER as a device for decreasing inflation and inflation uncertainty. It could be argued that providing the market with a quantitative target for the price stability objective would not be optimal if it were not accompanied with a clear central bank's roles and objectives, the features which distinguish IT framework.

The results provide evidence, represented by lower inflation and inflation persistence for inflation targeting countries, that the framework which has a direct quantitative target of inflation could be a better signalling device than the soft peg. The institutional features which accompany adopting IT might uphold the economy to move from high inflation to low inflation levels, as the credibility of the system, and thereby the policy outcome, hinges upon the development of monetary constitutions and transparent policies. The differences in such institutional arrangements

[16] Poland adopted different monetary regimes during the 1990s. Moreover, Cuestas et al. (2016) find that the Polish inflation rates are co-moved with that of the Euro Zone. When the sample is splitted to cover the time after IT, all GARCH models exhibit no remaining ARCH effect.

might explain why the advantages of the monetary framework differ across countries.

6 Conclusions

This paper investigates the relationship between inflation and inflation uncertainty under two monetary regimes with a nominal price level target: inflation targeting and a soft fixed exchange rate regime. Given the important monetary implications of the nexus between inflation-inflation uncertainty, a large number of studies have been carried out to examine whether inflation causes inflation uncertainty, as suggested by Friedman (1977) or inflation uncertainty leads to higher inflation as stated by Cukierman and Meltzer (1986). However, few studies have shed light on the plausible influence of each regime on reducing inflation uncertainty. This study is carried out to fill the gap in the literature by empirically assessing the validity of the Friedman and Cukierman-Meltzer hypotheses under the two regimes and examining the effect of monetary regimes on the relationship between inflation and inflation uncertainty. In particular, it aims to investigate the impact of quantitative targets of monetary policy upon reducing inflation uncertainty. To do this, we utilise the monthly CPI data collected from the IMF-IFS database of two fixed exchange rate targeters: Jordan and Egypt, and three inflation targeting countries: South Africa, Brazil and Poland, over the span 01:1980-06:2014.

In order to examine the two hypotheses simultaneously, we apply GARCH-M model, which allows the inflation rate to be determined by the conditional variance. We incorporate the standard deviation as a proxy for inflation uncertainty in the mean equation and augment the variance equation with lagged inflation. Nevertheless, as inflation series exhibit positive skewness and leptokurtosis, all the GARCH models are estimated with the assumption that the errors have a generalised error distribution. The impact of the regime is assessed by employing a constant and regimes slope dummy. For each country we start by constructing different ARMA models. The general to specific approach leads to AR specifications which ensure the whiteness of the residuals. The regimes dummies are introduced to different lags of inflation, but only two interactive dummies are selected to be imposed on the mean equation, based on the improvement in the model fit. Furthermore, we account for the time of apartheid in South Africa and the political disturbances, due to the Arab Spring, in Jordan and Egypt. The results from the OLS of the conditional mean of inflation, run to the time before and after adopting the examined regime, reveal that inflation targeters, unlike fixed exchange rate countries, have experienced stable inflation after IT. The "IT regime" constant dummy variable, plugged in the mean equation, is found significant and negative, reflecting the direct effect of the regime on lowering the average inflation.[17] This is confirmed by the statistics of average inflation, in which the ITers have enjoyed lower average inflation after shifting to IT. On the contrary, such desirable effects are not found for exchange rate targeters, as the regime constant dummy appears significantly positive for Egypt, and insignificant for Jordan. Political dummies incorporated for the two exchange rate targeters and South Africa are found to positively affect the inflation trend and the mean.

As the construction of GARCH-M considers only the magnitude of inflation shocks, we apply asymmetric GARCH-M type model to account for the responses of the conditional inflation uncertainty to increasing or falling in inflation. More importantly, the long-run effect of inflation on inflation uncertainty is examined

[17] The results with a regime constant dummy reported only for the case of Poland.

by the CGARCH-M, which allows decomposing the short-run from the long-run component.

The relationship between inflation and inflation uncertainty appear to be consistent with Friedman and Cukierman-Meltzer hypotheses. The parameter estimates of inflation uncertainty proxy and the one-period lagged inflation, from all the GARCH-M models and for all the countries, are significantly positive. The asymmetric GARCH specifications show that inflation affected the process of inflation uncertainty differently in Egypt and Brazil, while for Jordan, South Africa and Poland, the conditional inflation uncertainty is not influenced by the direction of inflation shocks.

The impact of the monetary regime on inflation uncertainty is examined by employing the slope inflation dummy in the conditional variance equation for all GARCH-M models. A negative and significant parameter indicates that inflation is reduced across all the countries, irrespective of the monetary regime followed by the central bank. The negative effect holds even when we account for asymmetries; the slope dummy remains with a negative sign. Nevertheless, the influence of the regime is not clear for Egypt, as the coefficient appears with weak magnitude. This could be due to the instability of the soft exchange rate regime in Egypt during the analysed period.[18]

The negative estimates of the slope dummy under the CGARCH-M reflects the ability of both regimes to reduce inflation uncertainty in the long-run; however, this effect is found to be insignificant for South Africa. It is also noted that the inflation trend of fixed exchange rate targeters converges to the mean slower that that of the ITers.

In general, the paper has made a way towards enhancing the understanding of the effectiveness of announcing an explicit quantitative target on inflation uncertainty. Furthermore, the positive relationship found between inflation and inflation uncertainty in both directions underlines that a monetary regime with an ultimate goal of price stability is a necessity to keep inflation and inflation uncertainty constrained. Our findings add to the growing body of literature on the importance of inflation targeting as a framework for monetary policy. IT, according to our results, appears effective in lowering the inflation persistence and inflation uncertainty more than the soft fixed exchange rate regime. However, it is shown that ITers have not equally benefited from IT. One possible explanation for this might be attributed to the institutional differences among countries in terms of the level of central bank independence and transparency.[19]

[18] Exchange rate devaluations occurred between 1991 and 1992. In 2001, the government announced a new parity to the US dollar and shifted gradually to a crawling peg system then the central bank abandoned the fixed system in January 2003.

[19] Note that countries under the fixed exchange rate lose monetary freedom, as they have to keep their monetary policies in tune with the base country, but they can still ensure institutional independence from the political authority.

Table 1: : Inflation Properties

	Jordan	South Africa	Egypt	Brazil	Poland
Panel (a)					
Average	2.12	3.88	4.6	33.95	4.44
Maximum	37.65	18.81	51.51	361.23	46.78
Minimum	-34.65	-3.87	-36.16	-2.66	-7.68
Average-New Anchor	3.43	1.83	4.92	2.77	2.32
Average-Before	2.62	4.91	3.96	58.04	6.25
Panel (b)					
Skewness	0.65	0.78	0.95	2.45	2.75
Kurtosis	9.22	4.57	9.575	10.31	15.59
ARCH effect	8.20***	23.01***	76.12***	27.36***	19.56***
Breusch-Godfrey LM	1.86	3.8	0.53	1.27	4.81
Jarque-Bera	1385.88	331.09	2226.67	7703.5	942.89

Note: Average-New anchor represents the average of inflation for each country for the adoption period of the monetary regime under investigation. Average-Before shows the average before adopting the examined regime. However, since Egypt opted out of the pegged exchange rate regime in January 2003, the average of the new anchor reflects the average during the exchange rate targeting, and hence, the Average-Before is the average after abandoning the fixed exchange rate system. Brazil was subject to hyperinflation during 1980s until March 1994 owning to the default related fear of the international government debt, see Garcia (1996) for more details. The statistics reported under the Breusch-Godfrey LM test shows the Obs*R-squared statistics with a chi-squared distribution. The null hypothesis of the test is that there is no autocorrelation up to lag order 'p'. ARCH effect test is a Lagrange multiplier test, by Engle (1982), in which the null hypothesis indicates homoskedasticity.

Table 2: : Unit Root Tests

	ADF con.	ADF con./trend	PP con.	PP con./trend
Jordan	-12.05***	-12.07***	-19.02***	-19.01***
South Africa	-5.75***	-10.16***	-15.71***	-17.41***
Egypt	-24.35***	-24.94***	-23.88***	-24.55***
Brazil	-3.97***	-4.64***	-6.13***	-7.59***
Poland	-3.52***	-3.91**	-14.21***	-14.94***

Note: In the ADF test, the lag length is determined by Schwartz information criterion, while for the Phillips and Perron test, the default Bartlett Kernel and Newy-West bandwidth are used. The asterisks ***, **, * denote rejection of the null hypothesis, i.e., inflation has a unit root for the ADF and PP tests, at 1%, 5% and 10% levels of significance, respectively.

Table 3: **OLS estimates of inflation conditional mean for Jordan** 6

coefficient	**full sample** 1981:02-2014:06	**pre-target** 1981:02-1995:09	**post-target** 1995:10-2014:06
γ_0^{JO}	0.003***	0.004***	0.003***
γ_2^{JO}	0.117**	0.154**	0.002
γ_5^{JO}	0.173***	0.259***	-0.031
γ_9^{JO}	0.099**	0.125*	0.019
γ_{12}^{JO}	-0.100**	-0.082	-0.174***
ARCH(1)	8.20***	5.68**	0.218
ARCH(2)	17.36***	13.70***	0.217
ARCH(12)	26.66***	20.25*	21.13**

Note: ARCH(1), ARCH(2) and ARCH(12) are ARCH test at 1^{st}, 2^{nd} and 12^{th} lag, respectively.

Table 4: **OLS estimates of inflation conditional mean for Egypt** 7

coefficient	**full sample** 1981:02-2014:06	**during-target** 1981:02-2002:12	**opting out** 2003:01-2014:06
γ_0^{EG}	0.008***	0.009***	0.007***
γ_1^{EG}	-0.155***	-0.199***	0.380***
γ_9^{EG}	0.124***	0.142**	-0.098
γ_{12}^{EG}	-0.144***	-0.144***	-0.084
ARCH(1)	65.02***	38.68***	0.162
ARCH(2)	71.46***	42.85***	0.249
ARCH(12)	104.90***	62.33***	8.85

Note: ARCH(1), ARCH(2) and ARCH(12) are ARCH test at 1^{st}, 2^{nd} and 12^{th} lag, respectively.

Table 5: **OLS estimates of inflation conditional mean for South Africa** 8

coefficient	**full sample** 1981:02-2014:06	**pre-target** 1981:02-2000:01	**post-target** 2000:02-2014:06
γ_0^{SA}	0.007***	0.008***	0.004***
γ_1^{SA}	0.231***	0.139**	0.397***
γ_2^{SA}	0.213***	0.228***	0.048
γ_3^{SA}	0.134***	0.108*	0.176
γ_7^{SA}	0.131***	0.173***	-0.032
γ_8^{SA}	0.100***	0.106*	0.07
γ_{11}^{SA}	0.106***	0.114**	0.003
γ_{12}^{SA}	-0.124***	-0.089	-0.214***
ARCH(1)	34.47***	12.82***	0.27
ARCH(2)	34.90***	13.16***	1.77
ARCH(12)	56.67***	26.60***	18.12

Note: ARCH(1), ARCH(2) and ARCH(12) are ARCH test at 1^{st}, 2^{nd} and 12^{th} lag, respectively.

Table 6: **OLS estimates of inflation conditional mean for Brazil** 9

coefficient	**full sample** 1981:02-2014:06	**pre-target** 1981:02-1999:06	**post-target** 1999:07-2014:06
γ_0^{BR}	0.063**	0.110***	0.005***
γ_1^{BR}	0.462***	0.447***	0.741***
γ_2^{BR}	0.367***	0.358***	-0.062
γ_8^{BR}	0.085**	0.065	0.035
ARCH(1)	25.58***	13.08***	1.29
ARCH(2)	25.69***	13.13***	4.36
ARCH(12)	29.16***	14.63	8.3

Note: ARCH(1), ARCH(2) and ARCH(12) are ARCH test at 1^{st}, 2^{nd} and 12^{th} lag, respectively.

Table 7: **OLS estimates of inflation conditional mean for Poland** 10

coefficient	**full sample** 1981:02-2014:06	**pre-target** 1981:02-1998:08	**post-target** 1998:09-2014:06
γ_0^{PO}	0.008***	0.013***	0.003***
γ_1^{PO}	0.218***	0.181***	0.294***
γ_2^{PO}	0.232***	0.221***	0.136*
γ_5^{PO}	0.138***	0.118*	0.128*
γ_9^{PO}	0.114**	0.099	0.093
γ_{11}^{PO}	0.181***	0.177***	0.113
ARCH(1)	19.96***	8.34***	0.02
ARCH(2)	20.61***	9.79***	0.03
ARCH(12)	36.78***	18.46	1.4

Note: ARCH(1), ARCH(2) and ARCH(12) are ARCH test at 1^{st}, 2^{nd} and 12^{th} lag, respectively.

Table 8: : **GARCH-M models for Jordan 11 (with dummies)**

Coefficients	GARCH-M	EGARCH-M
Conditional mean		
D_t	-3.97E-04	-3.20E-04
Pol	0.176***	0.187***
δ_1^{JO}	-0.148***	-0.125***
δ_2^{JO}	0.091***	0.077***
δ_3^{JO}	0.116***	0.081***
δ_4^{JO}	0.061**	0.047**
δ_5^{JO}	0.036	0.019
δ_6^{JO}	-0.207***	-0.208***
δ	0.328***	0.337***
conditional variance		
ϕ	7.47E-06**	-1.590***
α_1	0.027	
β_1	0.810***	0.127
β_2		0.142
λ_0	0.003***	19.19
λ_1	-0.002***	-25.67***
diagnostic statistics		
	$Q(1)$=2.15	$Q(1)$=1.27
	$Q(12)$=13.22	$Q(12)$=13.70
	$Q^2(12)$=14.05	$Q^2(4)$=13.94
	$TR^2(12)$=14.6	$TR^2(12)$=14.28
Wald test	20.7***	

Note: δ tests the validity of Cukierman and Meltzer (1986) hypothesis, where a positive δ indicates that inflation uncertainty increases inflation. λ_0 is the one-period lagged inflation and tests the validity of Friedman (1977)-Ball (1992) hypothesis, where a positive λ_0 means that inflation raises inflation uncertainty. D_t is the constant-monetary regime, i.e., fixed exchange rate system, dummy variable. Pol is the slope dummy that acts for the effect of Arab Spring on average inflation. Wald test examines the significance of the interactive regime dummy, i.e., $\lambda_1 = \lambda_2 = 0$.

Table 9: : **GARCH-M models for Jordan 11 (without dummies)**

Coefficients	GARCH-M	EGARCH-M	CGARCH-M
Conditional mean			
δ_1^{JO}	-0.143***	-0.131***	-0.139***
δ_2^{JO}	0.063*	0.026	0.051
δ_3^{JO}	0.095**	0.085***	0.117***
δ_4^{JO}	0.059**	0.113***	0.063**
δ_5^{JO}	0.058**	0.045***	0.044*
δ_6^{JO}	-0.194***	-0.153***	-0.179***
δ	0.361***	0.341***	0.338***
conditional variance			
ϕ	7.95E-06**	-2.493**	4.74E-05
α_1	0.034		0.005
β_1	0.785***	0.121	0.764
β_2		0.052	
λ_0	0.003***	28.851*	0.003***
λ_1	-0.002**	-24.709*	-0.003**
τ			
μ			0.025
ρ			0.823***
diagnostic statistics			
	$Q(1)$=2.21	$Q(1)$=2.34	$Q(1)$=2.27
	$Q(12)$=12.23	$Q(12)$=17.85	$Q(12)$=13.01
	$Q^2(12)$=14.71	$Q^2(12)$=21.18	$Q^2(12)$15.26
	TR^2(12)=15.22	TR^2(12)=21.05	TR^2(12)=15.82
Wald test	10.58***		

Note: δ tests the validity of Cukierman and Meltzer (1986) hypothesis, where a positive δ indicates that inflation uncertainty increases inflation. λ_0 is the one-period lagged inflation and tests the validity of Friedman (1977)-Ball (1992) hypothesis, where a positive λ_0 means that inflation raises inflation uncertainty. Wald test examines the significance of the interactive regime dummy, i.e., $\lambda_1 = \lambda_2 = 0$.

Table 10: : GARCH-M models for Egypt 12 (with dummies)

Coefficients	GARCH-M	EGARCH-M
Conditional mean		
D_t	0.002***	0.0003*
Pol	0.999***	0.999***
δ_1^{EG}	-0.223***	-0.131***
δ_2^{EG}	-0.008	-0.137***
δ_3^{EG}	0.151***	-0.008***
δ_4^{EG}	-0.033	-0.001***
δ_5^{EG}	-0.083	0.002***
δ	0.621***	0.679***
conditional variance		
ϕ	9.29E-09	0.170***
α_1	0.399***	
β_1	0.664***	0.265***
β_2		0.956***
λ_0	3.65E-06	-59.505***
λ_1	4.43E-04*	30.83***
diagnostic statistics		
	$Q(1)$=3.29	$Q(1)$=2.78
	$Q(12)$=40.10	$Q(12)$=14.54
	$Q^2(12)$=4.43	$Q^2(12)$=24.69
	$TR^2(12)$=4.16	$TR^2(12)$=26.19**
Wald test	16.39***	

Note: δ tests the validity of Cukierman and Meltzer (1986) hypothesis, where a positive δ indicates that inflation uncertainty increases inflation. λ_0 is the one-period lagged inflation and tests the validity of Friedman (1977)-Ball (1992) hypothesis, where a positive λ_0 means that inflation raises inflation uncertainty. D_t is the constant-monetary regime, i.e., fixed exchange rate system, dummy variable. Pol is the slope dummy that acts for the effect of Arab Spring on average inflation. Wald test examines the significance of the interactive regime dummy, i.e., $\lambda_1 = \lambda_2 = 0$.

Table 11: : **GARCH-M models for Egypt 12 (without dummies)**

Coefficients	GARCH-M	EGARCH-M	CGARCH-M
Conditional mean			
δ_1^{EG}	-0.282***	-0.263***	-0.161**
δ_2^{EG}	-0.068	-0.047	-0.034
δ_3^{EG}	0.189***	0.204***	0.138***
δ_4^{EG}	-0.105***	-0.071**	-0.05
δ_5^{EG}	-0.159***	-0.166***	-0.151***
δ^{EG}	0.986***	0.987***	0.984***
conditional variance			
ϕ	3.18E-07	-0.062	9.88E-05**
α_1	0.091***		0.149***
β_1	0.892***	0.244***	0.276
β_2		0.100***	
λ_0	0.0003**	-5.647*	2.45E-05
λ_1	-0.0004***	-1.545	-0.0004***
τ			
μ			0.032***
ρ			0.991***
diagnostic statistics			
	$Q(1)$=0.0004	$Q(1)$=0.544	$Q(1)$=0.18
	$Q(12)$=5.91	$Q(12)$=5.604	$Q(12)$=5.71
	$Q^2(12)$=20.35	$Q^2(12)$=14.74	$Q^2(12)$=12.94
	$TR^2(12)$=18.44	$TR^2(12)$=13.48	$TR^2(12)$=12.53
Wald test	10.39***		

Note: δ tests the validity of Cukierman and Meltzer (1986) hypothesis, where a positive δ indicates that inflation uncertainty increases inflation. λ_0 is the one-period lagged inflation and tests the validity of Friedman (1977)-Ball (1992) hypothesis, where a positive λ_0 means that inflation raises inflation uncertainty. Wald test examines the significance of the interactive regime dummy, i.e., $\lambda_1 = \lambda_2 = 0$.

Table 12: : **GARCH-M models for South Africa** 13

Coefficients	GARCH-M	EGARCH-M	CGARCH-M
Conditional mean			
APART	0.003***	0.004***	0.003***
δ_1^{SA}	0.305***	0.348***	0.300***
δ_2^{SA}	-0.112*	-0.104	-0.107*
δ_3^{SA}	0.042	0.003	0.032
δ_4^{SA}	0.131**	0.165***	0.135**
δ_5^{SA}	0.129***	0.159***	0.115***
δ_6^{SA}	0.001	-0.009	0.005
δ_7^{SA}	0.077**	0.064*	0.079**
δ_8^{SA}	0.037	0.03	0.036
δ_9^{SA}	-0.166***	-0.147***	-0.160***
δ	1.557***	1.467***	1.546***
conditional variance			
ϕ	5.89E-6**	-7.736**	7.14E-06
α_1	0.038		0.001***
β_1	0.135	0.034	-0.001***
β_2		0.158	
λ_0	0.001***	31.769	-0.073
λ_1	-0.001**	-70.414**	-0.131
τ			
μ			0.028**
ρ			0.870***
diagnostic statistics			
	$Q(1)$=1.02	$Q(1)$=1.25	$Q(1)$=1.27
	$Q(12)$=13.00	$Q(12)$=15.21	$Q(12)$=14.06
	$Q^2(12)$=25.96	$Q^2(4)$=20.27	$Q^2(12)$=25.53
	TR^2(12)=21.50	TR^2(12)=18.87*	TR^2(12)=21.16*
Wald test	10.82***		

Note: δ tests the validity of Cukierman and Meltzer (1986) hypothesis, where a positive δ indicates that inflation uncertainty increases inflation. λ_0 is the one-period lagged inflation and tests the validity of Friedman (1977)-Ball (1992) hypothesis, where a positive λ_0 means that inflation raises inflation uncertainty. APART is a constant dummy variable capturing the effect of apartheid on average inflation. Wald test examines the significance of the interactive regime dummy, i.e., $\lambda_1 = \lambda_2 = 0$.

Table 13: : **GARCH-M models for Brazil** 14

Coefficients	GARCH-M	EGARCH-M	CGARCH-M
Conditional mean			
APART	0.003***	0.004***	0.003***
δ_1^{BR}	0.223***	0.848***	0.227***-
δ_2^{BR}	-0.294***	-0.228***	-0.031***
δ_3^{BR}	0.532***	0.344***	0.777***
δ_4^{BR}	0.321***	0.228***	-0.192***
δ_5^{BR}	0.069***	0.093***	-0.042***
δ	0.063***	0.456***	0.464***
conditional variance			
ϕ	0.006***	5.321***	0.004***
α_1	6.861***		0.174***
β_1	0.142***	-0.041	0.167***
β_2		0.225***	
λ_0	0.256***	9.919***	0.054***
λ_1	-0.427***	-100.193***	-0.080***
τ			
μ			0.169***
ρ			0.812***
diagnostic statistics			
	$Q(1)$=0.593	$Q(1)$=6.18	$Q(1)$=1.577
	$Q(12)$=13.267	$Q(12)$=46.93	$Q(12)$=202.72
	$Q^2(12)$=0.453	$Q^2(12)$=0.135	$Q^2(12)$=1.72
	TR^2(12)=0.432	TR^2(12)=1.51	TR^2(12)=1.65
Wald test	1163.90***		

Note: δ tests the validity of Cukierman and Meltzer (1986) hypothesis, where a positive δ indicates that inflation uncertainty increases inflation. λ_0 is the one-period lagged inflation and tests the validity of Friedman (1977)-Ball (1992) hypothesis, where a positive λ_0 means that inflation raises inflation uncertainty. Wald test examines the significance of the interactive regime dummy, i.e., $\lambda_1 = \lambda_2 = 0$. The sum of α_1 and β_1 in the conditional variance is larger than one. This is because Brazil was subject to hyperinflation during 1980s to March 1994.

Table 14: : **GARCH-M models for Poland** 15

Coefficients	GARCH-M	EGARCH-M	CGARCH-M
Conditional mean			
D_t	-0.001*	-0.004***	
δ_1^{PO}	0.141*	0.272***	0.142*
δ_2^{PO}	0.185***	0.264***	0.192***
δ_3^{PO}	0.182***	0.114**	0.054
δ_4^{PO}	0.001	-7.51E-05	0.026
δ_5^{PO}	-0.012	-0.061**	-0.12
δ_6^{PO}	0.072***	0.064**	0.068
δ_7^{PO}	0.071**	0.081***	0.133***
δ	1.439***	1.690***	1.722***
conditional variance			
ϕ	2.14E-06**	-3.227***	6.15E-06***
α_1	0.024		0.047***
β_1	0.693***	0.035	0.004
β_2		0.017	
λ_0	0.001***	29.109***	0.001***
λ_1	-0.001***	-19.046**	-0.001***
τ			
μ			0.052***
ρ			0.562***
diagnostic statistics			
	$Q(1)$=0.01	$Q(1)$=0.002	$Q(1)$=13.15
	$Q(12)$=9.38	$Q(12)$=2.89	$Q(12)$=60.50
	$Q^2(12)$=27.85	$Q^2(12)$=59.30	$Q^2(12)$=69.59
	$TR^2(12)$=26.21***	$TR^2(12)$=53.59***	$TR^2(12)$=56.38***
Wald test	12.46***		

Note: δ tests the validity of Cukierman and Meltzer (1986) hypothesis, where a positive δ indicates that inflation uncertainty increases inflation. λ_0 is the one-period lagged inflation and tests the validity of Friedman (1977)-Ball (1992) hypothesis, where a positive λ_0 means that inflation raises inflation uncertainty. D_t is the monetary regime-constant dummy variable. Wald test examines the significance of the interactive regime dummy, i.e., $\lambda_1 = \lambda_2 = 0$.

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