2017

“Is inflation mean-reverting? Evidence from Egypt during 1974-2016”

Dr. Mohamed S. Abed Mekheimar
Associate Professor of Economics
Faculty of Economic Studies & Political Science Alexandria

Introduction

Egyptian economy has witnessed unstable record of annual inflation rates over the last four decades. In the early 1970s inflation rate was at low level Shifted to a moderate level after the first oil-prices shock to the mid-1980s. Inflation took a faster track and averaged about 19% annually during second half of the 1980s to the early 1990s. After the successful implementation of the stabilization program with the International Monetary Fund (IMF) and the World Bank (WB) during 1992-1997, inflation decreased to a single digit about 4%.

The abandonment of using foreign exchange rate as a nominal anchor and several devaluations in the early 2000s inflation rate increased again to around 11%. In 2005, Central Bank of Egypt (CBE) announced his intention to implement a full-fledged inflation-targeting policy once the fundamental prerequisites are met.

In the aftermath of the revolution in 2011, inflation increased again progressively. In November 2016, Egyptian government reached an agreement with the IMF according to which it has access to billion 12$ of Extended Fund Facility over the next three years. This program entailed very restrictive monetary and fiscal policy to reduce the budget deficit as percentage of GDP and to put domestic debt on a sustainable track. Drastic measures prescribed by the IMF' staff such as floatation of the Egyptian pound against the U.S. dollar, replacing the sales tax with the value-added tax, raising the Energy prices reducing the government subsidy to food and basic utilities. All these measures and others pushed inflation rate in Egypt to unprecedented level of more than 30% in the last quarter of 2016.

The purpose of this study is to examine the mean reversion of inflation rate in Egypt over the last five decades by using stationarity tests that take into consideration the existence of multiple structural breaks.

Based on the results of the mean-reverting investigation, the study will analyze its implications for the inflation-targeting policy proposed by the CBE in the mid-2000s.

To enhance the performance of inflation-targeting policy with respect to inflation-forecasting accuracy, the study will estimate a nonlinear model to try to capture the dynamics of inflation rate in Egypt using monthly data during 1974:M1-2016:M12. Monthly data of inflation rate in Egypt exhibits clear volatility compared with annual data; this kind of instability can be better explained using time-varying conditional variance models. Nonlinear models such as Smooth Transition Autoregressive (STAR) which allows for regime switching could be suitable to explain such behavior.

The paper is organized as follows. The first part presents a brief literature review of previous studies concerning inflation mean reverting. The second part will give some stylized facts about inflation rates in Egypt over the period 1974M1-2016M12. In the third provides information about data and methodology. Testing for non-stationarity and estimating a non-linear STAR model will be given in the third part. The fourth part provides the specification of a univariate nonlinear STAR model to explain inflation behavior in Egypt. The fifth part will analyze the empirical results of estimation of the STAR model. The final part will give the conclusion and policy implications of the study.

I. Literature review

As discussed in the seminal work of Nelson and Plosser (1982), several macro variables are having unit root and this would have theoretical and policy-making implications. One of the important macroeconomic variables is inflation rate as it has wide impacts on other macroeconomic variables such as economic growth, unemployment rate, interest rates and demand for money. Therefore, the nature of the integration process of inflation time series would have important implications.

The validity of several economic models depends crucially on the integration properties of inflation rate. For example, the Fisher effect theory assumes that inflation rate is non-stationary, i.e., integrated I (1) and co-integrated with nominal interest rate in order make real interest rate stationary. Hence, Fisher effect tests are possible if inflation and interest rates series are integrated of order one Koustas and Serletis, (1999) and Mishkin, (1992). Similarly, the expectations augmented Phillips curve model requires inflation rates to be non-stationary to allow wages and prices to share a long-run relationship. The accelerationist hypothesis requires non-stationary inflation rates as authorities must accept an ever-increasing inflation level just to keep unemployment rate below its natural level.

Conversely, the rational expectation version of Cagan (1956) indicates that stable growth of money supply implies stationary inflation unless there are bubbles. The hypothesis of natural rate of inflation and the sticky price model of Taylor (1979) assume that inflation is stationarity.

The importance of identifying the integration proprieties of inflation rate is also essential for policy makers conducting monetary policy. Arize et al. (2005) and Cecchetti and Debelle (2006) argue that the cost of disinflation would be higher if inflation rate is found to be non-stationary as shocks to inflation would have a permanent effect. Whereas, a stationary (mean reverting) inflation rates would lower the cost of controlling inflation as the shocks would have a transitory impact on inflation.[1]

The empirical evidence in related literature does not provide a consensus agreement about the stationarity of inflation. There are several models that support the notion that inflation is following a nonlinear process, e.g., Arango and Gonazalez (2001), Martin and Miles (2004), and Kilian and Manganelli (2008).

Arize (2011) provides some evidence on the non-stationarity of inflation rates in thirty-four African countries, Egypt included. Using quarterly inflation rates data, he applied nonlinear KSS test developed by Kapetanios et al. (2003) and his empirical results support the non-stationarity of the inflation rates in 25 African countries. Nonlinear mean-reverting behavior of inflation was found in only nine African countries, i.e., Burundi, Central African Republic, Egypt, Ghana, Guinea-Bissau, Lesotho, Madagascar, Seychelles and Swaziland. With respect to the case of Egypt, the study used quarterly data for the period 1980Q1-2009Q3; and the null hypothesis of non-stationarity was rejected at 5% significance level.

Anoruo and Murthy (2014), use nonlinear STAR unit root testing procedures to examine the issue of inflation convergence for the Central African Economic and Monetary Community (CEMAC) member states.[2] The results from tests suggest that inflation differentials for the sample countries are nonlinear and have mean reverting processes. The finding of inflation convergence indicates the possibility of implementing a common monetary policy and using inflation targeting regime within CEMAC. Interestingly, they argue that under inflation targeting policy, inflation's Data Generating Process (DGP) becomes a nonlinear process because the responsiveness of the Central Banks may change depending on whether the current inflation rate is above or below the inflation target level, thus creating a "band of inaction."[3]

Noureldinem (2005), used quarterly data to evaluate the robustness of three approaches to forecast inflation in Egypt during the period 2001Q1-2002Q4. The alternative three models were the output gap (Philips curve) model, the money gap model and a VAR model. The empirical results have shown the superiority of the money gap model over the other models in capturing the dynamics of inflation during the period 1980Q4-2000Q4; but the three models performed poorly with respect to forecasting. The author explained such poor performance because of observed anomaly in the out-of-sample data.

Noureldinem (2005) used inflation at its first difference to avoid spurious regression and to achieve stationarity; but he did not use tests for stationarity of inflation that considers possible structure breaks or nonlinearity such as LS test introduced by Lee and Strazicich (2003) and KSS test (2003); respectively. Using these tests could have helped in understanding the observed anomaly in the data. Moreover, as we mentioned above using “expectations augmented Phillips curve” model requires inflation to be non-stationary rather than stationary to allow wages and prices to share a long-run relationship; and therefore, using first difference stationary inflation rate would invalidate the results of this model.[4]

The merits of this study are threefold. First, it covers relatively large time period 1974M1-2016M12 using monthly data which allows for better understanding of inflation dynamics in Egypt. Second, using non-stationarity tests that take into consideration the possibility of structural breaks, i.e., ZA test developed by Zivot and Andrewes (2002) for one possible break and LS test introduced by Lee and Strazicich (2003) for two possible breaks. In addition, the study is also using a more advanced nonlinear Smooth Transition Autoregressive (STAR) unit root test, i.e., the KSS test developed by Kapetanios et al. (2003). KSS test is found to be more powerful than ZA and LS as the latter tests don’t include the nonlinear STAR model in the alternative hypothesis while the former does. Third, the study will estimate a nonlinear model to explain the dynamic behavior of inflation rates in Egypt using a regime swishing STAR model.

II. Stylized facts about inflation rates in Egypt 1974-2016

Egyptian economy has witnessed unstable record of annual inflation rate over the last four decades. In the early 1970s inflation rate was at low level; but shifted to a moderate level after the first oil-prices shock to the mid-1980s. In the second half of the 1980s to the early 1990s annual inflation rate took a faster track and averaged about 19% annually. After the successful implementation of stabilization and liberalization programs with the IMF and the World Bank, inflation rate was put under control and averaged about 4% annually.

As the monetary authority abandoned the use of foreign exchange rate as a nominal anchor, several devaluations of the Egyptian Pound against the U.S. Dollar in the early 2000s caused the inflation rate to increase to 11%. In 2005, the Central Bank of Egypt CBE announced his intention to adopt a full-fledged inflation targeting regime once the fundamental prerequisites are met.[5] Unfortunately, the transition period to move towards full-fledged inflation-targeting toke a decade from 2005 to 2015. During the 2005-2010, annual inflation rate continued to increase from 5% to 11%.

In the aftermath of the revolution in 2011, the inflation rates accelerated progressively. Over the period 2011-2015, the average rate of inflation was about ten percent a year. Several factors have contributed to the increase in inflation in this turmoil period, especially the political instability and the fragile security situation triggering great economic uncertainty, increase in fiscal deficit, and high growth rate of domestic money supply.[6] As the CBE tried to maintain stability of foreign exchange rate against strong speculative attacks and abrupt capital flights, international reserves were almost depleted. Meanwhile, the main sources of foreign currencies inflows (i.e., tourism, Egyptian worker's remittances and Suez Canal fees) suffered from a great setback. This situation generated a great pressure on the Egyptian government to devaluate the pound.

To ease the pressure on the Egyptian Economy (caused by the scarcity of the U.S. dollar), the CBE decided to devaluate the pound by 13% to EGP 8.85 per US dollar in in March 2016. Even though such devaluation failed to ease the pressure in the foreign exchange market, it caused inflation to increase further through raising imports' prices. In the second half of 2016, pressure in foreign exchange market became more sever and the gap between the official price and the black-market price was getting wider.

In November 2016, the Egyptian government decided to float the pound against the U.S., so that the foreign exchange rate will be decided according to demand and supply in the market. This step was part of a three years IMF economic reform program supported by an arrangement under the Extended Fund Facility (EFF) of $12 billion.[7]

Immediately after the floatation of the Egyptian pound, half of its value was lost and eventually stabilized around LE18 per unite of dollar. Inflation has immediately jump following the depreciation of the Egyptian pound via the pass-through effect. At the end of 2016, inflation rate reached 23% on 12-months basis and continued to increase in the first quarter of 2017 to reach all times high level of 34%. Foreign exchange rate changes have played major role in high inflation rates during the period of 2011-2016.

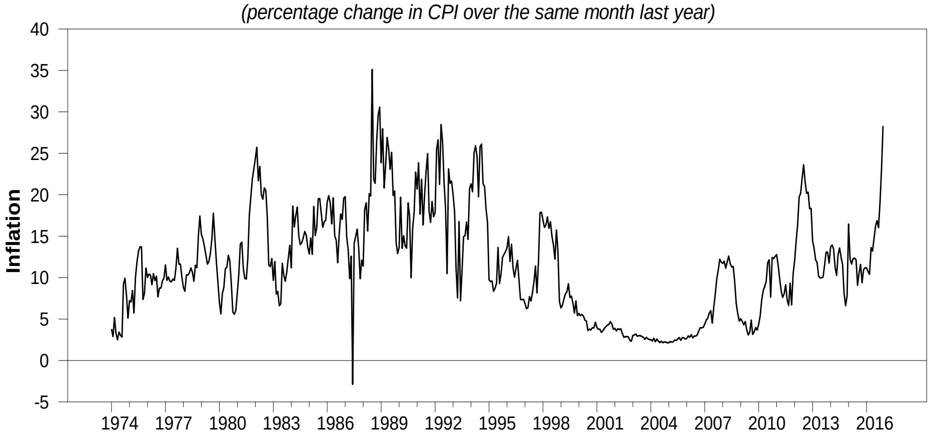
Data is drawn from the IMF's International Financial Statistics (IFS) data-base (2016) of monthly consumer price index (CPI) over the period 1974M1-2016M12, with 516 observations. The data set of 12- month inflation time series is constructed, by taking the percentage change in CPI over the same month of last year.

Apart from the period 2000-2007, Figure (1) displays visible instability of the 12-months inflation rate over the period of study. The null hypothesis of normality is rejected for the distributions of the data set based on the Jarque-Bera' statistic (JB).[8] The rejection of normality supports the use of a nonlinear model in the analysis of inflation rates in Egypt.

Compared to normal distribution with the same mean and variance, the shape of unconditional distribution of the 12-month inflation in Figure (2) suggests two separate modes; one is in the upper part including the highest value and most of the observations and the other is in the lower part embodying the smallest values of the observations. This is a clear indicator for the possible existence of two different regimes in the behavior of the 12-months inflation. Descriptive statistics in Table (1) reveals also that distribution of the dataset is found to be slightly positive skewed which means that positive shocks are more common than negative shocks to inflation rates in Egypt.

Figure (1)12-Months Inflation Rates in Egypt 1974M1-2016M12

(percentage change in CPI over the same month last year)



Source: IFS, IMF.

Figure (2) Distribution of *infmy* vs. Normal Distribution

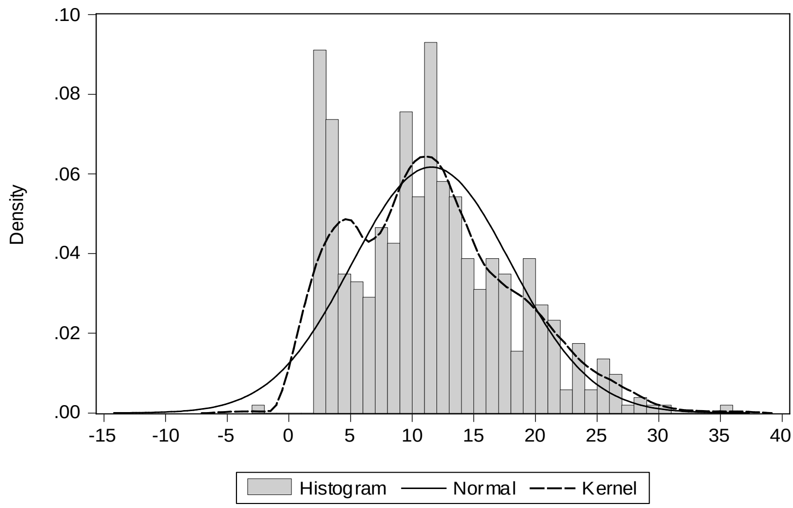


Table (1) Descriptive Statistics of 12-month Inflation in Egypt							
Mean	Median	Maximum	Minimum	Std. Dev.	Skewness	Kurtosis	**Jarque-Bera**
11.56880	11.17547	35.11111	-2.877698	6.462969	0.516442	2.837412	**23.50565 (0.00008)**
Sample 1974M01-2016M12) Observations 516							

III. Testing Stationarity and Mean Reverting of Inflation in Egypt

In this part, we are going to start with conventional linear unit-root tests that disregard the possibility of structural breaks, i.e., Augmented Dickey-Fuller (ADF) test, Elliott-Rothenberg-Stock (DF-GLS) test, and *Phillips-Perron* (PP) test. Results of the linear unit-root tests in Table (2) reveals that the null hypothesis of non-stationary *infmy* cannot be rejected at any significant levels for ADF and DF-GLS tests. For PP test, null hypothesis of non-stationary *infmy* is rejected at 1% significance level in the case of including a constant or a constant and a linear trend.

Table (2) Unite root test for 12-month Inflation (*infmy*) 1974M1-2016M12					
Linear Unit-root tests	H0: *infmy* is Non-stationary H1: *infmy* is Stationary				
Augmented Dickey Fuller	Lag Length: 14 (Automatic - based on SIC, maxlag=18)				
	t-Sta.	Prob.	Test critical values		
Exogenous:			1%	5%	10%
Constant	-2.1603	0.2214	-3.4428	-2.8669	-2.5696
Constant, Linear Trend	-2.2151	0.4796	-3.9759	-3.4186	-3.1317
None	-0.2585	0.5928	-2.5695	-1.9414	-1.6163
MacKinnon (1996) one-sided p-values.					
Elliott-Rothenberg-Stock DF-GLS	Lag Length: 14 (Automatic - based on SIC, maxlag=18)				
	t-Sta.	Prob.	Test critical values		
Exogenous:			1%	5%	10%
Constant	-0.6992	---------	-2.5695	-1.9414	-1.6163
Constant, Linear Trend	-1.5860	---------	-3.4800	-2.8900	-2.5700
Elliott-Rothenberg-Stock (1996, Table 1).					
Phillips-Perron	Bandwidth: 7 (Newey-West automatic) using Bartlett kernel				
	t-Sta.	Prob.	Test critical values		
Exogenous:			1%	5%	10%
Constant	**-4.3473***	0.0004	-3.4428	-2.8669	-2.5697
Constant, Linear Trend	**-4.4303***	0.0021	-3.9759	-3.4185	-3.1318
None	-1.3295	0.1701	-2.5695	-1.9414	-2.5695
MacKinnon (1996) one-sided p-values.					
Rejection of the Null hypothesis at *1%, **5%, and ***10% significance					
Source: Calculated by the author.					

Using unit-root tests that take into consideration the possibility of structural breaks is very important as the sample period is covering five decades that have witnessed several structural shocks. Disregarding the possibility of structural breaks in unit-root tests could lead to the false acceptance of non-stationary. Two tests are used in this paper; i.e., Zivot-Andrews (ZA) test and Lee-Strazicich (LS) test. Whereas the former is testing non-stationarity of *infmy* with one possible structure break, the latter is testing non-stationarity with two possible breaks.

Results of ZA test in Table (3) reveals that the null hypothesis of non-stationarity is rejected at 1% significance level, so *infmy* is found stationary with one structure break in the intercept in 1994M8 and with one structure break in the intercept and trend in 1998M10. Table (4) presents the results of LS test which confirms the stationarity of *infmy* with only one structure break in the trend in 1998M9, as the null hypothesis of non-stationarity is rejected at 1% significance level.

Table (3) Zivot-Andrews Unit Root Test of 12-month Inflation (*infmy)* in Egypt		
Linear Unit-root test with one Structural Break	H0: *infmy* has a unit root H1: *infmy* stationary with one structural break	
Sample: 1974M01 2016M12	**Included observations: 516**	
Intercept	**Trend**	**Intercept & Trend**
Test statistic **-5.5339***	Test statistic -4.0292	Test statistic **-5.6011***
Break point **1994:08**	Break point 2009:01	Break point **1998:10**
Critical value	**Critical value**	**Critical value**
1% * -5.3400	1%* -4.9300	1%* -5.5700
5% ** -4.8000	5%** -4.4200	5%** -4.8000
Zivot and Andrews (2002)		
Source: Calculated by the author		

Table (4) Lee-Strazicich Unit-Root Test of 12-month Inflation (*infmy*) 1974M1-2016M12		
Linear Unit-root tests with two Structural breaks	H0: *infmy* is non-stationary H1: *infmy* is stationary	
Sample period 1974:01 - 2016:12	Estimated with 7 lags chosen from 8	
Model C : Trend Break Model with 2 breaks		
Variable	Coefficient	T-stat
S{1}	-0.2166	-6.7812
Constant	1.1078	4.4596
D(1990:01)	7.4745	2.8540
DT(1990:01)	-1.2134	-3.3902
D(1998:09)	6.4396	2.4466
DT(1998:09)*	**-2.8968***	**-5.3621***
Critical Values *1% = -5.2961 **5%= -4.8053 ***10% = -4.4841		
Model A: Crash Model with 2 breaks in the level		
Variable	Coefficient	T-stat
S{1}	-0.0463	-2.9845
Constant	-0.1459	-1.0669
D(1993:05)	-7.9185	-2.9292
D(1998:11)	-6.3924	-2.4100
Critical Values *1% = -3.8341 **5% = -3.2536 ***10% = -3.0089		
Lee-Strazicich (2003)		

Results of ZA and LS tests confirm the stationarity of the 12-month inflation rate in Egypt over the period of the study with one structural break in August 1994, September or October of 1998. These results suggest that inflation is mean reverting. Unfortunately, the problem with these linear unit-root tests lies in its alternative hypotheses of linear stationarity. If the data generating process (DGP) of inflation rate in Egypt is found to be nonlinear, the results of linear unit-root test would produce a misspecification problem with the nature of the mean reversion, as the linear stationarity models assume constant speed of mean reverting to equilibrium in all cases. On the contrary, nonlinear stationary process has different speed of mean reverting depending on the extent of deviation from equilibrium.

In the case of nonlinear model of Smooth Transition Autoregressive (STAR), speed of reverting tends to be relatively slow when the deviation from equilibrium is small and inflation tends to exhibit unit-root behavior; but the speed of reverting becomes stronger the larger the deviation from equilibrium and inflation in this case behaves as a stationary process. This is exactly the reason why linear unit-root tests are found to have difficulties in differentiating between unit-root process and stationary process with long persistence in cases of nonlinear DGP.[9] Inflation persistence is defined as the speed at which inflation adjusts back or reverts (converges) to its equilibrium after a shock.

The results of nonlinear KSS test reported in Table (5) reveal that the null hypotheses of non-stationarity of *infmy* is rejected (in favor of the alternative hypotheses of nonlinear stationarity of ESTAR process) for raw, demeaned, and de-trended data at 1%, 1% and 5% significance level; respectively.

Table (5) Kapetanios-Snell-Shin (KSS) Unit-root test of 12-month Inflation in Egypt			
Non-linear Unit-Root test		Ho: *infmy* is non-stationary Ha: *infmy* is stationary nonlinear ESTAR	
Regression Run From 1974:01 to 2016:12		Estimated with lags = 18	
Model C : Trend Break Model with 2 breaks			
	Case 1 raw data	Case 2 demean data	Case 3 detrend data
KSS-Statistic p-value	**-1.632*** (0.038)	**-2.951*** (0.008)	**-2.977**** (0.010)
Critical values			
1% * 5% ** 10% ***	-2.036 -1.547 -1.335	-2.880 -2.297 -2.012	-2.979 -2.430 -2.126
Kapetanios-Snell-Shin (2004)			
Source: Calculated by the author			

The results of KSS test provides strong evidence of nonlinear mean reverting of inflation rate in Egypt during the period of the study and support the choice of ESTAR model in the regression process in the next part.

IV. Smooth Transition Autoregressive (STAR) Model[10]

Based on the work of Teräsvirta and Anderson (1992), Granger and Teräsvirta (1993) and Teräsvirta (1994), STAR model was developed as a generalized model of the Threshold Autoregressive (TAR) model initially proposed by Tong (1978) and Tong and Lim (1980). TAR model belongs to the so-called regime-switching models, which allow for the time series y_t to move from one regime to another based on the value of an observable transition variable (q_t) relative to a certain threshold value (c).[11] If the transition variable is assumed to be a lagged value of the same time series, $q_t = y_{t-d}$, where d is a positive integer, the regime will be determined by the time series itself. This special case of TAR model is called Self-Exciting TAR (SETAR), which can be presented by the following equation:

$$y_t = \begin{cases} \beta_0 + \sum_{i=1}^{p} \beta_i y_{t-i} + \varepsilon_t & \text{if } y_{t-d} \le c \\ \beta_0^* + \sum_{i=1}^{p} \beta_i^* y_{t-i} + \varepsilon_t & \text{if } y_{t-d} > c \end{cases} \quad (1),$$

where, β_0 and β_0^* are constants in regime 1 and regime 2; respectively. β_i and $\beta_i^*, i = 1, \dots, p$ are autoregressive parameters of regime 1 and regime 2, respectively; and $\varepsilon_t \sim iid(0, \delta^2)$. Eq. (3) shows a SETAR 2-regime model, in which y_t can be either in regime 1 when the transition variable $y_{t-d} \leq c$ or in regime 2 when $y_{t-d} > c$; but the movement between the two regimes is assumed to be in a discrete manner or discontinuous.

In contrast to SETAR model, STAR approach allows for a rather smooth transition process between regimes by including a continuous transition function $F(y_{t-d} - c)$.[12] STAR-model can be illustrated by equation (2):

$$y_t = \begin{cases} (\beta_0 + \sum_{i=1}^{p} \beta_i y_{t-i} + \varepsilon_t)(1 - F(y_{t-d} - c) & if\ F(y_{t-d} - c) = 0 \\ (\beta_0^* + \sum_{i=1}^{p} \beta_i^* y_{t-i} + \varepsilon_t)\ F(y_{t-d} - c) & if\ F(y_{t-d} - c) = 1 \end{cases} \qquad (2),$$

where y_t will be in regime 1 when transition function $F(y_{t-d} - c)$ equals zero; or in regime 2 when transition function equals unity. The movement of $y_t$between the two regimes will depend on the explicit form of the transition function $F(y_{t-d} - c)$. There are two prominent forms that transition function may take, the exponential form and the logistic form. The choice between the two forms depends on whether the time series exhibits symmetric or asymmetric behavior towards negative and positive shocks. If the behavior of the time series is showing symmetric reaction, then transition function should take the exponential form and the model in this case will be called Exponential STAR (ESTAR). Logistic form will be used in the case of asymmetric behavior of the time series and the model will be called Logistic STAR (LSTAR) model. In the following, ESTAR and LSTAR models will be discussed in some details. In ESTAR model, transition function will take the following form

$$F(y_{t-d} - c) = 1 - exp\,[-\gamma(y_{t-d} - c)^2] \qquad (3),$$

where d is the delay parameter, c is the threshold, y_{t-d} is the transition variable and γ measures the speed of smoothness of the transition process. The smaller is the value of γ, the smoother will be the movement of y_t between regimes. if $\gamma = 0$, transition function will be a constant and ESTAR will be linear model. But if $\gamma \to \infty$, transition process will be abrupt and ESTAR model becomes a SETAR model. As illustrated in Figure (A), exponential transition function in Eq. (3) has symmetrical inverted bell shape distribution and its value is bounded between 0 and 1. The smaller the value of γ is the slower and smoother will be the transition process from regime to other. In LSTAR model, transition function will take the following form:

$$F(y_{t-d} - c) = \{1 - exp[-\gamma(y_{t-d} - c)]\}^{-1} \qquad (4),$$

where, d, c, y_{t-d}, and γ are defined as in Eq. (4). The shape of the logistic transition function is different from that of exponential function. As can be seen from Figure (B), logistic transition function has an “s” shape of asymmetrical distribution and is showing a monotonic increasing curve from 0 to 1 as y_{t-d} increases. The smaller the value of γ is the slower and smoother will be the transition process from regime to other.

Figure (A) Exponential Transition Function

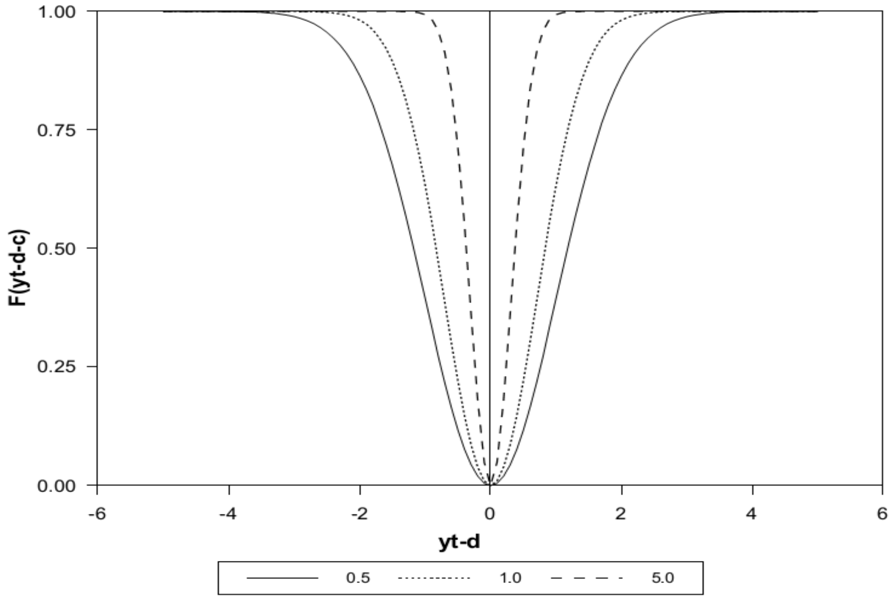
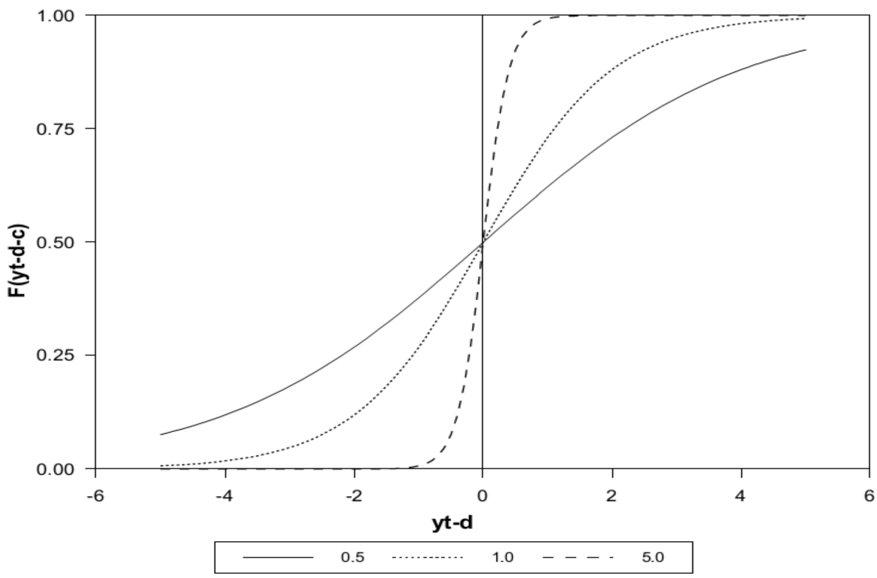


Figure (B) Logistic Transition Function



Specification of a STAR model for Inflation Rate

If y_t is defined as the inflation rate, to specify a STAR model for y_t, there is a systematic modeling cycle approach proposed by Teräsvirta (1994). This cycle consists of the following four steps: First, determine the order of autoregressive parameter *p*, based on the lower value of Akaike Information Criteria (AIC) and the Schwarz Information Criteria (BIC), or examining the partial autocorrelation function (PACF). Second, testing the autoregressive model for linearity against STAR model based on the value of *p*, as determined in the first step. Third, select the delay parameter *d*, by choosing its value that minimizes the *p-value* of the test of the null hypothesis of linearity. Fourth, choose between LSTAR and ESTAR specification by referring to a sequence of *F-test* to a third order Taylor expansion of the transition function; as illustrated in the following equation:

$$y_t = \beta_{00} + \sum_{j=1}^{p}(\beta_{0j}y_{t-j} + \beta_{1j}y_{t-j}\,y_{t-d} + \beta_{2j}y_{t-j}{y^2}_{t-d} + \beta_{3j}y_{t-j}{y^3}_{t-d}) + \varepsilon_t \qquad (5).$$

The linearity test involves testing the joint hypothesis that

$$H_0: \beta_{1j} = \beta_{2j} = \beta_{3j} = 0 \qquad (j = 1, \dots, p).$$

To determine the appropriate form for the transition function (either ESTAR or LSTAR) we need to test the following hypotheses:

$$H_{01}: \beta_{3j} = 0 \qquad (j = 1, \dots, p),$$
$$H_{02}: \beta_{2j} = 0 \mid \beta_{3j} = 0 \qquad (j = 1, \dots, p),$$
$$H_{03}: \beta_{1j} = 0 \mid \beta_{2j} = \beta_{3j} = 0 \qquad (j = 1, \dots, p).$$

The rejection of H_{01} implies selecting LSTAR form. Accepting H_{01} but rejecting H_{02} implies selecting ESTAR form. Accepting H_{01} and H_{02} but rejecting H_{03} implies selecting LSTAR form. Sarantis (1999) proposed a simpler selecting rule: if the p-value resulting from the rejection of H_{02} is the smallest ESTAR form will be selected, otherwise LSTAR form will be selected if H_{01} and H_{03} are rejected.

Escribano and Jordá (1999) proposed an alternative selection procedure using a second-order Taylor approximation of the exponential transition function:

$$y_t = \beta_{00} + \sum_{j=1}^{p}(\beta_{0j}y_{t-j} + \beta_{1j}y_{t-j}\,y_{t-d} + \beta_{2j}y_{t-j}{y^2}_{t-d} + \beta_{3j}y_{t-j}{y^3}_{t-d} + \beta_{4j}y_{t-j}{y^4}_{t-d}) + \varepsilon_t \qquad (6),$$

and testing the following hypotheses:

$$H_{oE}: \beta_2 = \beta_4 = 0,$$
$$H_{oL}: \beta_1 = \beta_3 = 0,$$

ESTAR (LSTAR) form will be selected if $H_{oE}(H_{oL})$ is rejected with the minimum p-value. This methodology will be applied, in part four, to Egyptian data over the period 1974M1-2016M12.

V. Empirical Results of the Regression of STAR Model

STAR molding cycle can be applied to the time series *infmy* as its stationarity is established by KSS, LS and ZA tests in part II. The first step in this cycle is to determine the order of autoregressive parameter *p* by examining the PACF in Figure (4) which shows that the appropriate number of lags is 27.

The second step entails testing the autoregressive model for linearity against STAR model based on the value of *p*, i.e., 27 lags. The result of linearity test is shown in Table (6) and indicates that linearity hypothesis is rejected at 5% significance level. The third step is to select the delay parameter *d* that minimizes the *p-value* of the test of the null hypothesis of linearity. Table (6) also shows that the minimum level of the F(*p-value)* for linearity test is (0.0114) achieved at lag 24. It is worth mentioning that this test is taking into consideration the problem of heteroscedasticity.[13] Table (7) shows the results of LM Nonlinear test to select the appropriate STAR-model. Based on these results ESTAR is found to be the right model to explain the nonlinear behavior of inflation rate in Egypt.

The last step in the cycle is to estimate the parameters in Eq. (2) with the specification of ESTAR transition function in Eq. (5) using nonlinear least squares. For the estimation process, we used the R-project program and RSTAR package developed by Balcilar (2008). Estimation results are shown in Table (8) indicates high explanatory power of the ESTAR model as the adjusted R-Square equals 0.90 and 28 parameters of the model are found significant. The estimated value of the threshold *c*, which represents the half way point between the two regimes, equals 17.54% and is found significant at 1% level.

Figure (4) Partial Autocorrelation Function (PACF) of *infmy*

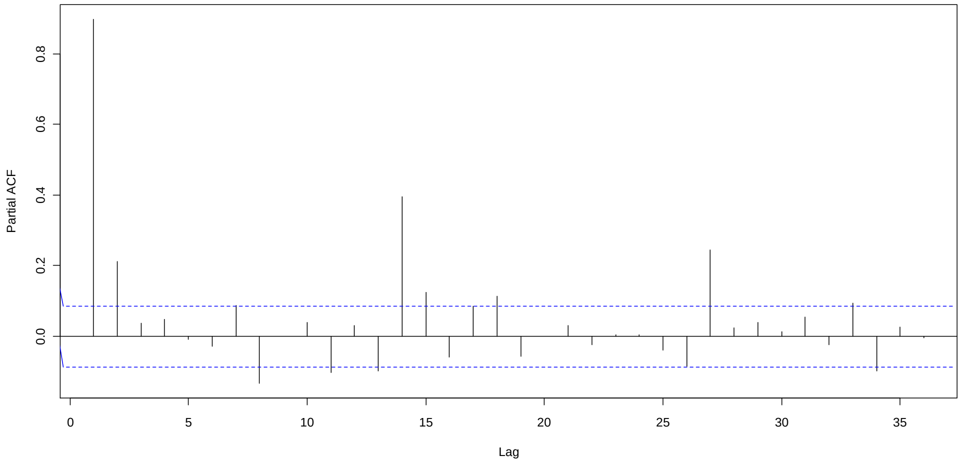


Table (6) Nonlinearity Test of STAR Model						
Null Hypothesis: no smooth threshold nonlinearity						
Allowing Heteroskedastic Errors using White Correction						
Lag	ChiSq-stat	ChiSq-dof	ChiSq.pv-val	F-stat	F-dof	F.pv-val
lag1	97.6549	81	0.1003	1.1707	-81,380	0.168
lag2	77.6822	81	0.5838	0.886	-81,380	0.7427
lag3	84.2855	81	0.3794	0.977	-81,380	0.5384
lag4	88.6733	81	0.2621	1.0391	-81,380	0.3975
lag5	91.4369	81	0.2007	1.079	-81,380	0.3157
lag6	104.3007	81	0.0417	1.2719	-81,380	0.072
lag7	87.8231	81	0.2831	1.027	-81,380	0.424
lag8	92.611	81	0.1778	1.0961	-81,380	0.2837
lag9	87.0952	81	0.3017	1.0166	-81,380	0.4471
lag10	85.032	81	0.358	0.9875	-81,380	0.514
lag11	84.6012	81	0.3703	0.9814	-81,380	0.5281
lag12	93.016	81	0.1703	1.102	-81,380	0.273
lag13	83.2432	81	0.4102	0.9625	-81,380	0.5724
lag14	83.581	81	0.4002	0.9672	-81,380	0.5614
lag15	77.3704	81	0.5936	0.8818	-81,380	0.7513
lag16	95.7691	81	0.1254	1.1426	-81,380	0.2068
lag17	77.059	81	0.6034	0.8776	-81,380	0.7597
lag18	90.045	81	0.2303	1.0588	-81,380	0.3559
lag19	100.8033	81	0.0673	1.2182	-81,380	0.115
lag20	92.9261	81	0.172	1.1007	-81,380	0.2754
lag21	102.2166	81	0.0557	1.2398	-81,380	0.0957
lag22	113.9754	81	0.0093	1.4258	-81,380	0.0153
lag23	106.3931	81	0.0308	1.3045	-81,380	0.0531
lag24	**115.5965**	**81**	**0.007**	**1.4523**	**-81,380**	**0.0114**
lag25	94.1757	81	0.1502	1.119	-81,380	0.2438
lag26	95.4566	81	0.13	1.1379	-81,380	0.2138
lag27	94.4326	81	0.146	1.1228	-81,380	0.2376
Source: Author's calculations.						

Table (7) LM Nonlinear Test to Select the STAR Model							
Test	Chi Square Variants			F Variants			
	Test Stat	p value	df	Test Stat	p value	df 1	df 2
LM.1	59.22	3.33E-04	27	2.2151	5.37E-04	27	434
LM.2	124.46	1.74E-07	54	2.5733	8.34E-08	54	407
LM.3	147.05	1.01E-05	81	2.0174	5.75E-06	81	380
LM.3e	61.7	2.46E-04	28	2.2329	3.93E-04	28	433
LM.4	172.62	7.80E-05	108	1.7833	4.41E-05	108	353
LM.S2	74.23	2.75E-06	27	2.6975	1.55E-05	27	407
LM.S3	30.3	3.01E-01	27	0.9297	5.69E-01	27	380
LM.S4	36.57	1.03E-01	27	1.0567	3.91E-01	27	353
LM.H1	59.22	3.33E-04	27	2.2151	5.37E-04	27	434
LM.H2	74.23	2.75E-06	27	2.6975	1.55E-05	27	407
LM.H3	**30.3**	**3.01E-01**	**27**	**0.9297**	**5.69E-01**	**27**	**380**
LM.H4	36.57	1.03E-01	27	1.0567	3.91E-01	27	353
LM.HE	**61.80**	**2.18E-01**	**54**	**0.9457**	**5.86E-01**	**54**	**353**
LM.HL	61.28	2.31E-01	54	0.9365	6.04E-01	54	353
Source: Author's calculations							

The value of the estimated transition parameter γ equals 3 and is found significant at 1% level. The relatively small value of γ suggests a slower and smoother transition process between the two different regimes; i.e., between high and low inflation phases. This smooth adjustment process can be observed in Figure (5) that shows the estimated transition function versus the transition variable. The shape of the estimated transition function indicates that the behavior of inflation rate in Egypt is characterized by symmetric dynamics during both the low and the high inflation phases. Remarkably, the majority of *infmy* observations are lying in the lower level of inflation which below the threshold which means that inflation in Egypt was predominately low rather than high over the last 5 decades.

Figure (6) shows how the fitted values of ESTAR-model track closely the sharp turning points of observed *infmy*. Standardized residuals of the estimated ESTAR model are shown in Figure (7). Results of the diagnostic statistics of the residuals of ESTAR model in Figure (8) reveal that the null hypothesis of Ljung-Box test (no autocorrelation in the residuals) cannot be rejected, so residuals of the estimation are found random and independent up to 10 lags. This result is also confirmed by referring to autocorrelation function (ACF) of the residuals in Figure (10) as it shows no sign of autocorrelation up to 26 lags.

Table (9) shows the results of three different tests to residuals of ESTAR model; i.e., normality, ARCH and parameter constancy tests. These results reveal the following: first, the null hypothesis of normal distributed residuals is rejected at 1% significance level. Despite the non-normal distribution of the residuals, estimators of the model are still consistent. Second, the null hypothesis of no ARCH in the residuals is also rejected at 1% significance level. Third, the null hypotheses of parameters constancy test is partially accepted at 5% significance level.

Table (8) Estimation Results of *infmy* ESTAR model					
Number of observations used: 489					
Coefficient(s)	**Estimate**	**Std.Error**	**t-value**	**Pr(>|t|)**	
Gamma	3	0.72422	4.142	3.44E-05	***
Treshold_1	17.5406	0.23748	73.861	< 2e-16	***
Intercep_1	6.6089	2.64341	2.5	0.012414	*
AR_1(1)	0.6055	0.08749	6.921	4.50E-12	***
AR_1(2)	0.3786	0.10518	3.599	0.000319	***
AR_1(3)	-0.2149	0.09554	-2.249	0.024513	*
AR_1(4)	0.1284	0.10053	1.277	0.201609	
AR_1(5)	0.1284	0.09589	1.339	0.180441	
AR_1(6)	0.1262	0.08728	1.445	0.148378	
AR_1(7)	0.0278	0.09905	0.28	0.779176	
AR_1(8)	-0.3210	0.10889	-2.948	0.003195	**
AR_1(9)	-0.0265	0.10388	-0.255	0.798555	
AR_1(10)	0.4113	0.11245	3.657	0.000255	***
AR_1(11)	-0.3727	0.11907	-3.13	0.001747	**
AR_1(12)	-0.0204	0.10659	-0.192	0.847935	
AR_1(13)	-0.6641	0.10774	-6.164	7.08E-10	***
AR_1(14)	0.5249	0.11357	4.622	3.81E-06	***
AR_1(15)	0.3703	0.1115	3.321	0.000898	***
AR_1(16)	-0.3910	0.10714	-3.649	0.000263	***
AR_1(17)	-0.1550	0.11682	-1.327	0.184557	
AR_1(18)	0.4446	0.12726	3.494	0.000476	***

AR_1(19)	0.1654	0.0984	1.681	0.092812	.
AR_1(20)	-0.2683	0.10292	-2.607	0.009145	**
AR_1(21)	-0.0646	0.11575	-0.558	0.576634	
AR_1(22)	0.1221	0.12367	0.988	0.323352	
AR_1(23)	0.0670	0.10391	0.645	0.518991	
AR_1(24)	-0.3186	0.19269	-1.654	0.098227	.
AR_1(25)	-0.0391	0.11381	-0.343	0.731382	
AR_1(26)	-0.7669	0.12395	-6.188	6.11E-10	***
AR_1(27)	0.7035	0.10024	7.018	2.25E-12	***
Intercep_2	0.08918	0.24551	0.363	0.716433	
AR_2(1)	0.92133	0.0681	13.529	< 2e-16	***
AR_2(2)	0.12739	0.08272	1.54	0.123547	
AR_2(3)	-0.12184	0.07382	-1.65	0.098854	.
AR_2(4)	0.01627	0.07745	0.21	0.833629	
AR_2(5)	0.05175	0.07465	0.693	0.488129	
AR_2(6)	-0.17353	0.08863	-1.958	0.050257	.
AR_2(7)	0.24943	0.08471	2.944	0.003235	**
AR_2(8)	-0.02663	0.08373	-0.318	0.750432	
AR_2(9)	-0.08781	0.08076	-1.087	0.276863	
AR_2(10)	-0.03693	0.07612	-0.485	0.627544	
AR_2(11)	0.07662	0.07184	1.067	0.28615	
AR_2(12)	-0.06594	0.08039	-0.82	0.412099	
AR_2(13)	-0.5121	0.0805	-6.361	2.00E-10	***
AR_2(14)	0.49889	0.08237	6.057	1.39E-09	***
AR_2(15)	0.04296	0.07981	0.538	0.590441	
AR_2(16)	0.02476	0.06821	0.363	0.716575	
AR_2(17)	0.02309	0.07023	0.329	0.742298	
AR_2(18)	0.06126	0.07093	0.864	0.387729	
AR_2(19)	-0.16964	0.08346	-2.033	0.042077	*
AR_2(20)	0.08361	0.08191	1.021	0.307367	
AR_2(21)	0.09797	0.07694	1.273	0.202902	
AR_2(22)	-0.12638	0.07791	-1.622	0.104772	
AR_2(23)	-0.0657	0.07515	-0.874	0.381951	
AR_2(24)	0.17884	0.06586	2.715	0.006618	**
AR_2(25)	-0.08923	0.07509	-1.188	0.234748	
AR_2(26)	-0.07664	0.08482	-0.904	0.366234	
AR_2(27)	0.10509	0.0621	1.692	0.090591	.
Signif. codes: 0 ‘***’ 0.001 ‘**’ 0.01 ‘*’ 0.05 ‘.’ 0.1 ‘ ’ 1					
sigma^2 estimated	4.023	**AIC Criterion:**		2184.37	
log likelihood	-1034.18	**SIC Criterion:**		2427.52	
R-Square:	0.91025	**Adjusted R-Square:**		0.89839	
Source: Author’s Calculations					

Figure (5) Transition Function *vs* Transition Variable of *infmy*

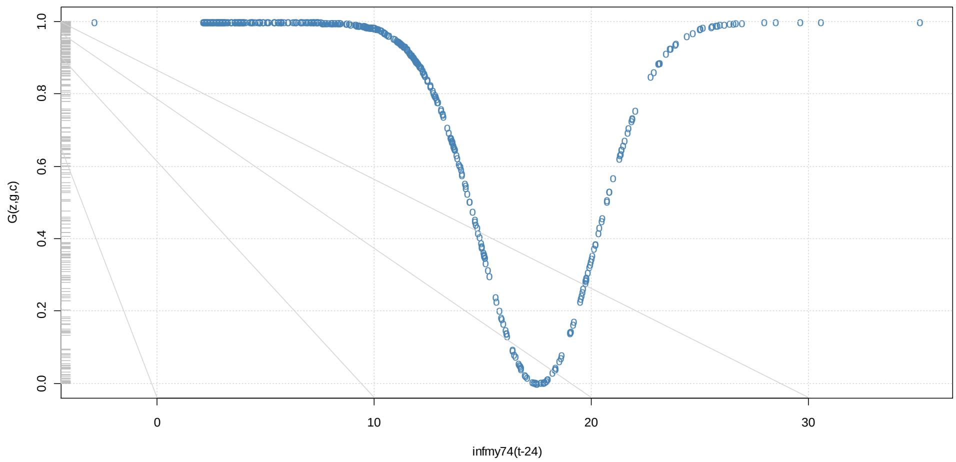


Figure (6) *infmy* and its ESTAR-model Fitted Values

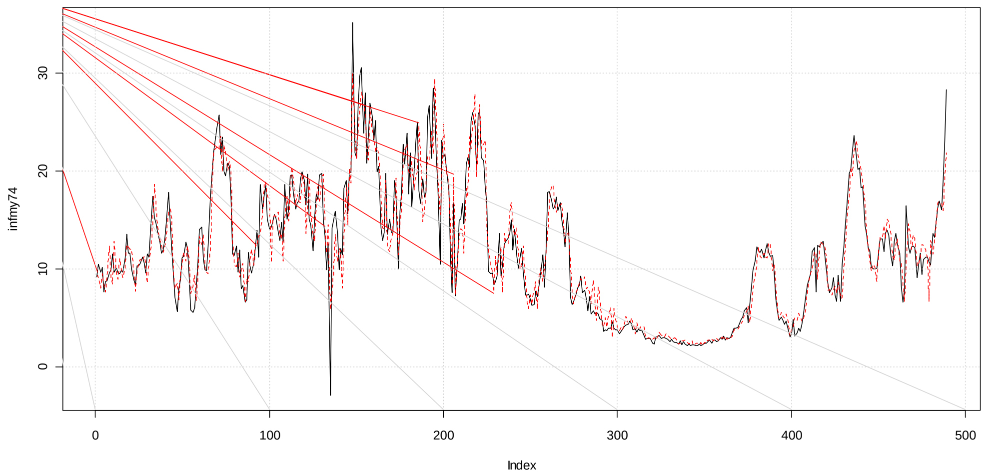


Figure (7) *infmy* Residuals of the ESTAR-model Fitted Values

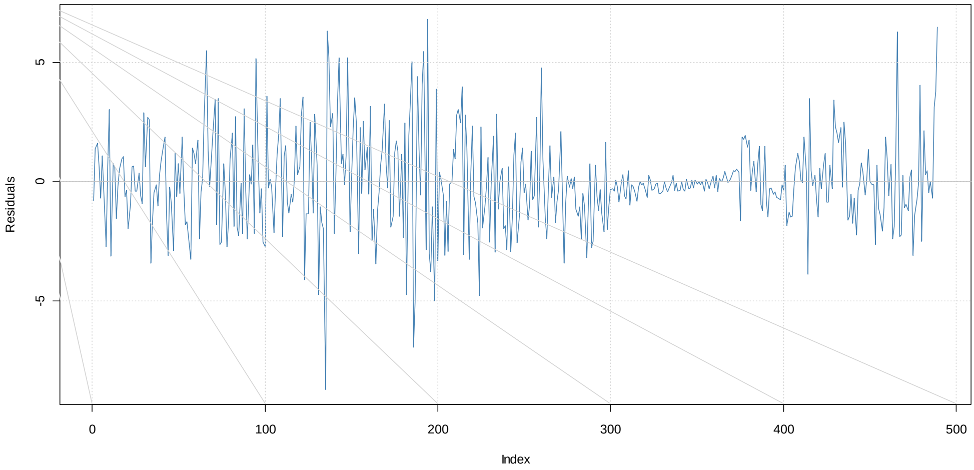


Figure (8) Ljung-Box P-value of the Residuals of the Fitted *infmy*

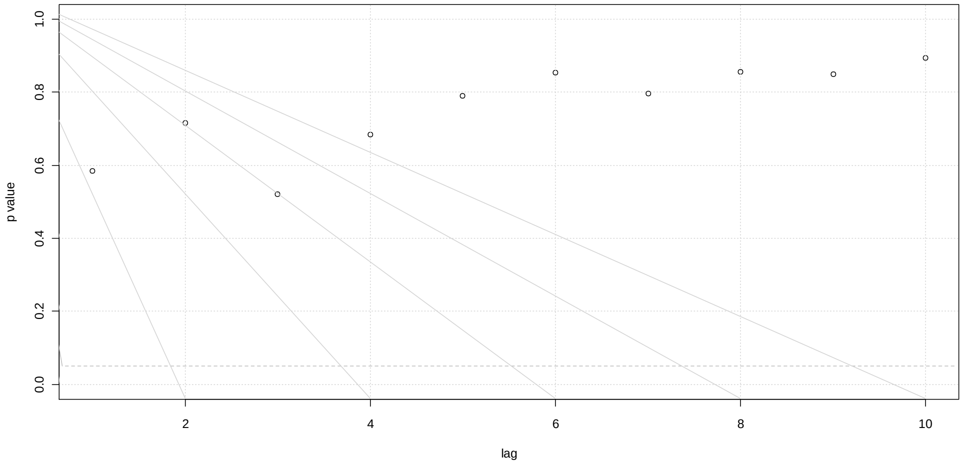


Figure (9) Autocorrelation Function (ACF) of the Residuals of the Fitted *infmy*

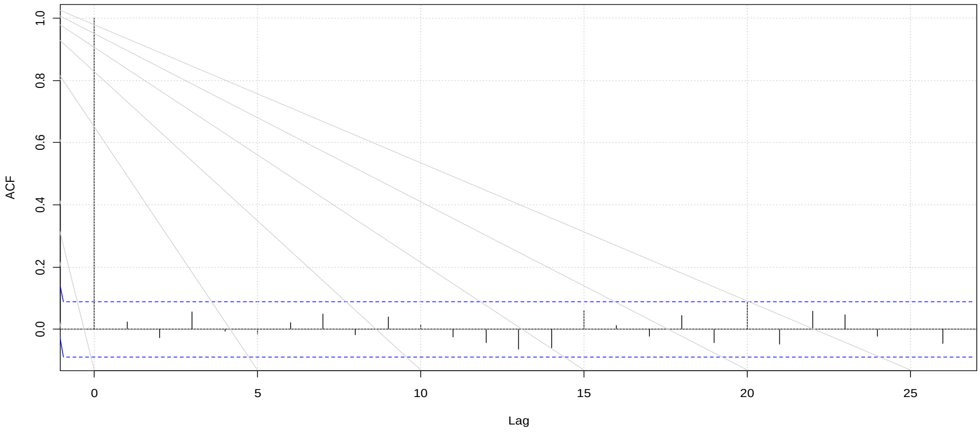


Table (9) Diagnostics of the residuals of *infmy* Fitted				
Jarque-Bera Normality Test :				
	X-Squared	p-val (X-Sq.)	F	p-val(F)
Test statistic JB	66.50184	8.216e-15	-------------	-------------
H0: Residuals are Normally Distributed				
LM Tests for ARCH:				
Estimation Method: Least Squares				
	X-Squared	p-val (X-Sq.)	F	p-val(F)
ARCH(1)	66.50184	3.330669e-16	76.67862	0
ARCH(4)	86.16491	0	25.92497	0
H0: No ARCH in the residuals				
LM Tests for Parameter Constancy (all variables):				
Estimation Method: Least Squares				
	X-Squared	p-val (X-Sq.)	F	p-val(F)
LM.C1	83.73781	1.513815e-02	1.328821	**6.424061e-02**
LM.C2	162.42478	2.904087e-03	1.350584	2.164342e-02
LM.C3	272.01849	2.855290e-06	1.851653	3.481695e-06
H0: Parameters are constant				
Source: Author's calculation.				

VI. Conclusions and policy implications

The study provides strong evidence of nonlinear mean reversion of inflation, as positive or negative shocks will have a transitory rather than permanent impact on inflation rate behavior. A nonlinear STAR-model is estimated to explain the behavior of a 12-month inflation rate in Egypt during the period 1974M01-2016M12. Based on the methodology proposed by Teräsvirta (1994) nonlinearity is detected in the data and the appropriate transition function is found to be the exponential-STAR (ESTAR) type. This transition function suggests that changes in inflation rate in Egypt exhibits symmetric behavior towards negative or positive shocks; i.e., during both the high and the low inflation phases.

The period of the study has witnessed a relatively unstable inflation rates especially after the 2011's revolution. Estimating results show high explanatory power of the ESTAR model and most of its parameters are found significant. Estimated transition parameter is relatively small and has a value of 3, which implies a slower and smoother transition process of the inflation rate adjustment between high and low regimes. The adjustment process between high and low inflation rates is found to be symmetric with a relatively slow delay of 24 months.

Based on the empirical results of the study, two main policy implications can be useful to deal with inflation in Egypt. First, monetary authority should intervene actively to prevent inflation rate from reaching the threshold of 17.5%, if they fail to avoid this level, inflation will shift to the higher phase. In this case, the cost of disinflation will be higher, in the form of severer restrictive monetary measures needed to curb inflation whether in the form of, for example, raising domestic interest rates or reducing money supply.

Second, as the study found strong evidence of nonlinear mean reversion of inflation, this will support the efforts of monetary authority to apply inflation targeting framework to reduce inflation rate in Egypt. In the case of mean reversion, positive or negative shocks will have a transitory rather than permanent impact on inflation rate behavior. Unless the monetary authority deliberately feeds the future inflation-expectations or delays its intervention to prevent inflation rate from reaching its threshold, inflation targeting could be successful at lower cost of disinflation.

Notes:

1. See Chen et al. (2016).
2. CEMAC includes Cameron, Central African Republic, Chad, Equatorial Guinea, Gabon and the Republic of Congo.
3. The link between inflation targeting and nonlinearity of inflation's DGP is investigated in Kapetanios et al. (2003) Christopoulos and Leon-Ledesma (2007).
4. For similar argument, see Arise (201), p.99.
5. Central Bank of Egypt (2005).
6. See Khan and Miller (2016).
7. International Monetary Fund (2016).
8. Jarque-Bera is testing whether the coefficients of skewness and excess kurtosis are jointly zero, the critical value of JB-statistics is 5.99. See, Bera and Jarque (1981).
9. See Froot and Rogoff (1994), pp. 7-11.
10. See Mekheimar (2013), pp. 6-9.
11. This type of regime-switching model is to be differentiated from the Markov-Switching model, proposed by Hamilton (1989).
12. See Brooks (2008, p. 474) and Franses and Van Dijk (2000, 71-73).
13. See Granger and Teräsvirta (1993).

References:

1. Anoruo, Emmanuel and Murthy, Vasudeva N.R. (2014), "Testing Nonlinear Inflation Convergence, for the Central African Economic and Monetary Community," International Journal of Economics and Financial Issues, Vol. 4, No. 1, pp.1-7.
2. Arango, L. E., and Gonazalez A. (2001). Some Evidence of Smooth Transition Non-linearity in Colombian Inflation Applied Economics, 34, 155-162
3. Balcılar, Mehmet, (2008), "RSTAR: A Package for Smooth Transition Autoregressive Modeling Using R," Prepared for The R User Conference 2008 Fakultät Statistik, Technische Universität Dortmund, Germany 12-14 August.
4. Bera, A. K. and Jarque, C. M. (1981) "An Efficient Large-Sample Test for Normality of Observations and Regression Residuals," Australian National University Working Papers in Econometrics 40, Canberra.
5. Brooks, Chris (2008), *Introductory Econometrics for Finance*, Cambridge University Press, Cambridge.
6. Cagan, P., (1956), "The Monetary Dynamics of Hyperinflation," In: Friedman, M. (ed.), Studies in the Quantity Theory of Money, Chicago: University of Chicago Press, 25-117.
7. Central bank of Egypt (CBE), (2005) Annual Report.
8. Chen, Shyh-Wei, Hsu, Chi-Sheng and Pen, Cyun-Jhen, (2016) "Are Inflation Rates Mean-reverting Processes? Evidence from Six Asian Countries," *Journal of Economics and Management*, 12, issue 1, pp. 119-155.
9. Christopoulos, D. K and Leãn-Ledesma, M. A. (2007), "A Long-Run Non-Linear Approach to the Fisher Effect," Journal of Money, Credit and Banking, 39(2-3), 543-559.
10. Elliott, G., T. J. Rothenberg, and J. Stock, (1996), "Efficient Tests for an Autoregressive Unit Root," Econometrica, 64, 813-836.
11. Escribano, A. and Jorda´,Ø. (1999), "Improved testing and specification of smooth transition regression models. In Nonlinear Time Series Analysis of Economic and Financial Data"; Rothman, P., Ed.; Kluwer: Boston, pp. 289–319.

12. Franses, P.H., van Dijk, D., (2000), Non-Linear Time Series Models in Empirical Finance. Cambridge University Press, Cambridge.
13. Froot, Kenneth A. and Rogoff, Kenneth (1994)."Perspectives on PPP and Long-Run Real Exchange Rates," NBER Working Papers 4952, National Bureau of Economic Research, Inc.
14. Granger, C.W.J., Teräsvirta, T., (1993), "Modelling Nonlinear Economic Relationships, Advanced Texts in Econometrics," Oxford University Press, Oxford.
15. Hamilton, J.D. (1989), A new approach to the economic analysis of nonstationary time series and the business cycle, Econometrica, 57, 357-384.
16. Hsiu-Yun Lee and Jyh-Lin Wu (2001), "Mean reversion of inflation rates: Evidence from 13 OECD countries," Journal of Macroeconomics, Volume 23, Issue 3, Summer 2001, Pages 477–487.
17. International Monetary Fund, (2016), International Financial Statistics, Database.
18. Khan, Mohsin and Elissa Miller (2016), "The Economic Decline of Egypt after 2011-Uprising," *Issue Brief* , Rafik Hariri Center for the Middle East, Atlantic Council.
19. Kapetanios, G., Shin, Y., Snell, A. (2003), "Testing for a Unit Root in the Nonlinear STAR Framework," *Journal of Econometrics, 112, 359–379.*
20. Koustas, Zisimos, and Apostolos Serletis (1992), "On the Fisher Effect." *Journal of Monetary Economics* 44 (August): pp. 105-130.
21. Lee, Junsoo and Mark C. Strazicich (2003), "Minimum LM Unit Root Test with Two Structural Breaks." *Review of Economics and Statistics,* 85(4), pp.1082-1089.
22. Martin, C., and Miles, C. (2004). Modelling Monetary Policy: Inflation Targets in Practice. *Economica*, 71, 209-221.
23. Mekheimar, Mohamed S. Abed (2013), "Nonlinear Model to Explain the Behavior of Daily Nominal Exchange Rate in Egypt, during the Period 2005-2010", *Journal of the Faculty of Commerce for Scientific Research*, Vol. 50. No. 1, Alexandria University, Egypt.

24. Mishkin, Frederic S. (1992), "Is the Fisher Effect for Real? A Reexamination of the Relationship Between Inflation and Interest Rates." *Journal of Monetary Economics* 30 (November): 195-215.

25. Nelson, C. R., and Plosser, C. I. (1982), Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications. *Journal of Monetary Economics*, 10, 139-162.

26. Noureldinem, Diaa (2005), "Alternative approaches to forecasting inflation in the case of Egypt, *Economic Research forum*," 12th Annual Conference, Cairo Egypt.

27. Phillips, P.C.B. and P. Perron (1988), "Testing for a Unit Root in Time Series Regression," *Biometrika*, 75, 335–346.

28. Sarantis, N. (1999), "Modeling non-linearities in real effective exchange rates," *Journal of International Money and Finance* 18, pp.27–45.

29. Taylor, J., (1979), "Staggered Wage Setting in a Macro Model," American Economic Review , 69, 108-113.

30. Teräsvirta, T. D., and Anderson, H.M. (1992), "Characterizing nonlinearities in business cycles using smooth transition autoregressive models," *Journal of Applied Econometrics* 7, S119–S136.

31. Teräsvirta, T., (1994), "Specification, estimation and evaluation of smooth transition autoregressive models," *Journal of the American Statistical Association* 89, 208–218.

32. Tong, H. (1978), On a threshold model. In Chen, C. H. (ed.): *Pattern Recognition and Signal Processing*, Applied Statistic (29), Sijthoff and Noordhoff, Netherlands Amsterdam, pp. 575-586

33. Tong, H. and Lim, K. S. (1980), Threshold Autoregression, limit cycles and cyclical data. *Journal of the Royal Statistical Society*, Vol. 42, No. 3, pp. 245-292.

34. Zivot, E. and D. W. K. Andrews, 1992, "Further Evidence on the Great Crash, the Oil Price Shock and the Unit Root Hypothesis." *Journal of Business and Economic Statistics* 10, 251-270.