**Performance of inflation targeting policy in Tunisia: Validation by Markov chains**

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**Abstract**

In this paper, we will study the inflation targeting in Tunisia as a final goal of monetary policy. This objective requires the choice of an adequate monetary aggregate to target this inflation rate within a well-defined price range. To this, we will summarize the main empirical studies that have dealt with the inflation targeting policy and validate this targeting for the Tunisian case during a study period from January 2001 to January 2018. We will use the Co-integration theory to obtain a long-term relationship that relates the inflation rate to the other explanatory variables and the Markov chains in order to study the performance of inflation targeting policy in Tunisia.

**JEL classification numbers:** **E47, E52, E58.**

**Keywords:** Markov chains, Performance, Inflation, Monetary policy.

**1. Introduction**

Inflation targeting provides a simple and predictable framework for monetary policy to anchor inflation expectations and to direct them downward through strengthening the credibility and the transparency of the central bank. Furthermore, it allows certain flexibility, in particular in case of exogenous or endogenous shocks, that is not possible with exchange rate anchoring. Finally, in a context of instability of the money demand function, inflation targeting offers a satisfactory alternative to the regulation of monetary aggregates as an intermediate objective.

This inflation targeting has brought about a clear improvement in inflation performance and was quickly chosen by some emerging countries (Chile and New Zealand). The International Monetary Fund does not anymore hesitate to encourage emerging economies to adopt inflation targeting, even if the conditions previously considered necessary for its success are not verified such as independence of the central bank from political power, institutional capacity to correctly analyze and predict inflation and the existence of sufficiently developed financial markets.

During the 1980s, Latin American countries have experienced high levels of inflation, in particular because of the monetary financing of the budget deficit. This is why it has become necessary to put in place monetary frameworks to reduce, in particular, incentives for budget slippage. Several approaches have been adopted to achieve this goal, ranging from dollarization to inflation targeting. Stabilization plans based on exchange rate policy had a rapid effect on inflation (through their impact on short-term expectations), while countries adopting inflation targeting achieved gradual performance (the credibility of the monetary policy requires more time to be built) but better in the long-run. The effectiveness of inflation targeting would therefore be less rapid but of longer duration. However, it is unclear how well inflation performance is directly attributable to inflation targeting. Indeed, inflation targeting has often been accompanied or preceded by important structural reforms (e.g. sound fiscal policies), making it more difficult to estimate the impact of these different factors on economic performance. Thus, the adoption of inflation targeting has generally been gradual (Chile 1990, Colombia 1991, Peru 1993, Mexico 1999) or, paradoxically, after a currency crisis (Brazil).

This paper aims to discuss the potential effectiveness of the inflation targeting framework that the Central Bank of Tunisia (hereafter CBT) tends to implement. Indeed, according to the declarations of its Governor Chedly Ayari on May, 23rd 2013 in the closing seminar of the twinning project between Tunisia and France, the CBT attempts to fulfill the prerequisites of inflation targeting. Applying several econometric techniques to monthly data over the period from January 2001 to January 2018, we conclude that inflation targeting fail in achieving a long-run equilibrium. In addition, we find that inflation in Tunisia exhibits cyclical behavior and that the average duration of a period of recession is longer than that of a period of expansion. The present article makes several contributions to the literature on inflation targeting. Unlike Lajnaf (2014) who used quarterly data and Amiri and Talbi (2013) who employed annual data to study the performance of inflation targeting in Tunisia, we use monthly data which allows the monetary authorities to rapidly adjust their level of inflation. In addition, unlike the majority of previous studies on this topic that used linear models, we employ a Markov switching model to detect the existence of a cyclical behavior of inflation.

The remainder of this article is structured as follows. Section 2 overviews the main previous literature that has addressed the performance of inflation targeting policies. Section 3 presents empirical analysis. We begin by investigating the existence of a long-run relationship between the consumer price index and the industrial production index, the nominal exchange rate, the monetary aggregate M3 and the money market average. This relationship is estimated using the two-step method of Engle and Granger (1987) and the Johansen (1991) approach. Then, we analyze the cyclical phenomenon of the inflation targeting policy in Tunisia through the Markov-regime switching model. Section 4 concludes.

**2. Literature review**

The empirical literature on inflation targeting is abundant. Ammer and Freeman (1995) survey the experience of inflation targeting in New Zealand, Canada and the United Kingdom and conclude that this policy is successful especially in the short-term. The authors also find that the cost of inflation reduction is important in New Zealand. Freeman and Willis (1995) assess the effectiveness of inflation targeting in four industrial countries - New Zealand, Canada, the United Kingdom and Sweden – and they find that the credibility of inflation targeting regimes have rapidly disappeared. Mishkin and Posen (1997) examine inflation targeting frameworks in New Zealand, Canada and the United Kingdom. They conclude that inflation targeting had some favorable effects on macroeconomic variables but that disinflation was already completed before inflation targeting was adopted: inflation targeting was used to keep disinflation gains rather than to facilitate disinflation in those countries. Laubach and Posen (1997) find similar results by analyzing the interest rate and consumer expectations for the same countries.

Debelle (1999) analyzes the case of several countries including New Zealand, Canada, the United Kingdom, Sweden, Finland, Spain and Australia. He shows that the adoption inflation targeting caused a decline in inflation and in the spread of long bond yields but a rise in unemployment. Kahn and Parrish (1998) analyze the evolution of inflation in New Zealand and Canada in the 1990s and conclude that Central Banks did not have better control of inflation with the adoption of inflation targeting regimes. They also estimate monetary policy responses functions for New Zealand, Canada, Sweden and the United Kingdom and find structural breaks in the functions of New Zealand and the United Kingdom.

Kuttner and Posen (1999) estimate a VAR model to test the impact of inflation targeting on the persistence of inflation. Their findings reveal a lack of change in persistence and response functions for Canada, the United Kingdom and New Zealand following the adoption of inflation targeting. Bernanke et al. (1999) conduct a comparative analysis of 9 industrialized countries and focused on three issues: does inflation targeting make disinflation less costly? Does it reduce inflation expectations?, and does it change the behavior of inflation? Their empirical analysis aims find answers to these questions and to test the stability of the parameters of the Phillips curve. They conclude that inflation targeting does not reduce the sacrifice ratio[[4]](#footnote-4).

Cecchetti and Ehrmann (2002) analyze the measure of policymakers’ inflation variability aversion before and after the adoption of inflation targeting with Ehrmann's methodology (1998) and based on structural VAR model for a group of 23 industrialized and developing countries over the period 1984-2000. They point out that both countries adopting inflation targeting and non-targeting countries increased their revealed aversion to inflation variability, and consequently experienced most increases in output volatility. Corbo, Landerretche, and Schmidt-Hebbel (2001) analyze the effectiveness of inflation targeting for three groups of countries: the first group is composed of countries that started inflation targeting before 1995, the second group comprising countries that started targeting in the 1990s and the last group is composed of industrialized countries that did not have explicit targets during the period 1980- 1999. They assess the performance of countries with and without inflation targets and look at the transmission mechanisms of inflation targeting using VAR models and the models proposed by Cechetti and Ehrmann (2002) and Corbo (1998). Their findings suggest that inflation targeting improves forward-looking expectations on inflation.

Sabban, Rozada and Powell (2003) assess the performance of inflation targeting regimes and investigate the impact of the adoption of inflation targeting on the evolution of the real and nominal exchange rates. They find that that inflation targeting policies are successful and that their performance differs across countries.

Kim (2014) examines the impact of inflation targeting on the purchasing power parity (PPP) in seven countries namely Canada, France, Japan, Italy, Sweden, the United Kingdom and the United States. He finds that inflation targeting reduces variability of real exchange rates and improves the long-run PPP.

Fazio, Tabak and Cajueiro (2015) investigate the causality between the implementation of inflation targeting policy and banking system instability. Their findings argue that bank system of inflation targeters are more stable, sounder and less distressed than those of countries without inflation targets.

Boughrara (2007) discusses the relevance of an inflation targeting policy for Tunisia. He finds that several impediments such as the fiscal dominance, the weakness of the financial system and the ignorance of the monetary transmission mechanism could lead to a failure of this policy.

Lajnaf (2014) assesses the performance of the inflation targeting policy in Tunisia using quarterly data over the period from 2000Q1 to 2011Q3. She argues that the choice of the monetary aggregate M3 as intermediate objective is not appropriate to implement this policy since it explains a low portion of the variance of the forecast error of the inflation target.

**3. Empirical analysis**

**3.1. Data**

The data used to estimate the optimal monetary policy rule in the inflation target are obtained from the Central Bank of Tunisia (BCT), the International Monetary Fund (IMF) and the World Bank (WB) for the period from January 2001 to January 2018.

**3.2. Model specifications**

Building on previous studies (Small and Porter, 1989; Mehra, 1991 and 1993; Feinman and Porter, 1992) and more recently Duca (2005), we assume that the performance of the inflation targeting policy (IT) depends on the monetary aggregate M3, the volume of transactions, the interest rate and the exchange. Inflation is measured on the basis of the harmonized index of consumer prices. Gross Domestic Product (GDP) is provided by volume and annual data. The unavailability of monthly GDP leads us replace it by the industrial production index. The money market average (MMA) is used as a proxy of interest rate[[5]](#footnote-5). We also use the intermediate objective M3[[6]](#footnote-6) chosen by the CBT as an instrument to achieve the ultimate objective of targeting the inflation rate. In addition, we use the nominal exchange rate (NER) as a measure of the external value of the national currency. In this paper, we use the following specification to assess the performance of the inflation targeting policy:

 (1)

To test the existence of a long-run relationship that describes the performance of inflation targeting policy and the cyclical phenomenon of this policy adopted by the BCT, we linearize equation (1) by introducing the natural logarithm:

 (2)

In this model, the constant Log (A) represents the average effect of the omitted variables while the non-explanatory variables are nested in the error term ε. The first step to verify the performance of inflation targeting policy is to test the stationary of each variable in the model. Table 1 reports the results of the Dickey-Fuller-Augmented test (1981) for the variables used.

**Table1. Test of Dickey-Fuller-Augmented (1981)**

|  |  |  |
| --- | --- | --- |
|  | **In level** | **In first difference** |
|  | None | Intercept | Trend and intercept | None | Intercept | Trend and intercept |
| **LCPI** |  |  | -2.1633 |  |  | -8.2404 |
| **LM3** | 9.0504 |  |  | -5.0778 |  |  |
| **LIPI** |  | -0.7664 |  |  | -12.1271 |  |
| **MMA** |  |  | -0.5478 |  |  | -4.5348 |
| **NER** | -2.7928 |  |  | -8.3644 |  |  |
| **Critical values** | -1.9425 | -2.8763 | -3.4332 | -1.9425 | -2.8764 | -3.4333 |

Table 1 shows that all the variables of this model are stationary after a first differencing since the t-statistics of Student are lower than the critical value of Mackinnon (1990) at 5%. Hence, these variables are integrated of order one or I(1). Also, we test the stationary of these variables using a unit root test with the endogenous structural change of Perron (1997). The results of the Perron test (1997) are reported in table 2.

**Table2. Model with change of the mean and the slope**[[7]](#footnote-7)

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
| Variables | t-statistics | Breakdown dates | Critical values | Number of lags |
| LCPI | -3.3491 | 2007:08 | -5.08 | 12 |
| LM3 | -2.2617 | 2009:09 | -5.08 | 12 |
| LIPI | -2.8497 | 2010:01 | -5.08 | 12 |
| NER | -6.4656 | 2007:04 | -5.08 | 0 |
| MMA | -4.2373 | 2012:01 | -5. 08 | 12 |

The presence of unit roots, with and without structural change for all variables of this model, requires the use of the co-integration approach to estimate the long-run relationship and to study the linear adjustment of the inflation rate relative to its fundamental value. We begin by estimating our model by the OLS technique. The results are reported in table 3.

**Table3. Estimating the performance of inflation targeting policy in Tunisia by OLS**

|  |  |  |
| --- | --- | --- |
| Variables | Coefficients | *p*-value |
| Constant | 1.0525 | 0.000 |
| LM3 | 0.3795 | 0.0001 |
| LIPI | -0.0241 | 0.0142 |
| NER | 0.0002 | 0.0421 |
| MMA | -0.0055 | 0.0321 |
| R-squared  | 0.9951 |
| Durbin-Watson | 0.2561 |

As can be seen, the nominal exchange rate, the industrial production index, and the money market rate have a weak effect on the performance of monetary policy. In contrast, the positive and strong coefficient of the monetary aggregate M3 suggests that this latter represents the primordial determinant of the effectiveness of the inflation targeting policy. However, it is necessary to study the stability of the residuals of this long-run relationship.

**Table4. Stationary test of the residual**

|  |  |
| --- | --- |
| Optimal number of lags | 1 |
| Type of the test : random walk without constant and without linear trend | DF |
| t-statisticsCritical values at 5% | -13.8901-1.9425 |

Since the residual is stationary in level, the co-integration relation of Engle and Granger (1987) between the variables can therefore be accepted. However, the results are less satisfactory and in line with theoretical expectations. In order to test the inflation targeting policy performance, we estimate an error correction model (ECM) model that includes short-term adjustments where used variables are stationary by differencing and the long-run equilibrium where these variables are stationary by linear combination, provided that the error correction term is negative and significant. The results of the use of the ECM are reported in table 5.

**Table5. Results of ECM**

|  |  |  |
| --- | --- | --- |
| Variables | Coefficients | *p-*value |
| Constant | 0.0016\*\*\* | 0.0000 |
| LCPIt-1 | 0.3934\*\*\* | 0.0000 |
| LM3t | 0.0401\*\* | 0.0293 |
| LIPIt | -0.0044 | 0.1199 |
| NERt | 0.0002 | 0.2719 |
| MMAt | 0.0009 | 0.3758 |
| Residualt-1 | -0.0561\*\*\* | 0.0037 |
| R-squared | 0.2096 |
| Durbin-Watson | 1.9826 |

The error correction term has a negative and significant sign. Hence, the inflation targeting policy in Tunisia is effective since the target of this long-run relationship converges partially towards a stable situation in the long-run and the disequilibrium will be corrected to around 5% per CPI. We apply the multivariate approach of Johansen (1991) to detect the number of co-integration relationships. Therefore, we use the maximum likelihood technique to estimate the VECM model. To do this, we refer to trace tests and maximum eigenvalue tests to determine the number of co-integrating vectors. But before performing these tests, it is necessary to determine the optimal number of lags of the autoregressive vector (VAR) based on the two information criteria of AIC and Schwartz and the likelihood ratio statistic (LR)[[8]](#footnote-8). Table 2 shows the optimal number of lags for VAR.

**Table6. Optimal number of lags for VAR**

|  |
| --- |
| X1t = (LCPIt, LM3t, LIPIt, NERt,MMAt) |
| Lags | 1 | 2 | 3 | 4 |
| AIC | -29.6395 | -29.9161 | -29.9262\* | -29.8941 |
| Schwartz | -29.1378\* | -28.9929 | -28.5786 | -28.1190 |
| LR | 88.4774 (0.0000) | 36.2399 (0.0680) | 28.0604 (0.3051) |

**Note: The number in parentheses indicates the marginal asymptotic level, i.e., the probability that the value of the calculated statistic exceeds the tabulated value. Thus, an asymptotic marginal 99.7% or 89.78% means that for a threshold of fewer than 99.7% and 89.78 the hypothesis H0 of a single lag is accepted.**

The two information criteria and the likelihood ratio test indicate the existence of three lags for this multidimensional autoregressive vector X1t. Table 7 below shows the number of co-integrating vectors.

**Table7. Johansen test (1991)**

|  |  |  |
| --- | --- | --- |
|  | Trace test | maximum eigenvalue test |
| H0 | r=0 | r ≤ 1 | r ≤ 2 | r ≤ 3 | r ≤ 4 | r=0 | r=1 | r=2 | r=3  | r=4 |
| H1 | r ≥ 1 | r ≥ 2 | r ≥ 3 | r ≥ 4 | r=5 | r=1 | r=2 | r=3 | r=4  | r=5 |
| Statistic value | 75.40 | 39.23 | 25.42 | 13.28  | 2.89 | 36.17 | 13.81 | 12.13 | 10.4 | 2.89 |
| Crititical value at 5% | 68.52 | 47.21 | 29.68 | 15.41  | 3.76 | 33.46 | 27.07 | 20.97 | 14.07  | 3.76 |
| Co-integrating vector normalized by LCPI | LCPI | LM3 | LIPI | NER | MMA |
| 1 | -1.296 | 1.583 | -0.007 | -0.014 |

The trace and maximum eigenvalue tests indicate the existence of a single co-integrating vector for the consumer price index in relation to the various explanatory variables. The estimate of the Co-integration relation is made from the maximum likelihood method. This method gives the Co-integrating vector normalized by the consumer price index. We found that aggregate M3, MMA, and the nominal exchange rate has positive effects in increasing the rate of inflation. On the other hand, the index of the industrial product generates a reduction of the rate of inflation. We will use the Hendrey (1995) test to detect low exogeneity for the matrix of long-run adjustment coefficients.

We perform the Hendry (1995) test to verify if certain determinants of the inflation rate are weakly exogenous for the parameters of these cointegrating relationships. If this is the case, these parameters can be estimated without loss of information using the conditional model, which is easier to manage since it is extracted from the complete VECM models. This hypothesis of low exogeneity is expressed by the nullity of a certain number of coefficients of the matrix of the long-term adjustment speed. The null hypothesis of weak exogeneity and its alternative can be formulated as follows:



Under the null hypothesis, the weak exogeneity test follows a Chi-square distribution with r degree of freedom. If the statistic of the likelihood ratio is greater than the critical value of Chi-square, the variable is not weakly exogenous, i.e., this variable undergoes a phenomenon of error correction. The exogeneity test rsults for the variables used as determinants of the performance of the inflation targeting policy.

Table8. Test of weak exogeneity

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
| Variables | LCPI | LM3 | LIPI | NER | MMA |
| Chi-square (1) | 2.1337 | 1.2258 | 19.4325\*\*\* | 1.2984  | 0.4211 |
| *p-*values | 0.1441 | 0.2682 | 0.0000 | 0.2545  | 0.5164 |

**Note: \*\*\* indicates significance at 1% level.**

The results suggest that we can reject the null hypothesis of weak exogeneity of all the variables except the industrial production at index. So, we estimate the co-integration relations using a VEC model including five variables (LCPI, LM3, LIPI, NER, MMA), four of them are weakly exogenous (LCPI, LM3, NER, MMA). It is not necessary to model explicitly these variables over the long term, although they may influence the performance of the inflation targeting policy. From the multivariate approach of Johansen (1991), we can conclude that the policy of targeting the inflation rate is not efficient, since the target of the inflation rate does not converge towards a stable situation in the long term since the speed of adjustment has a contradictory sign. Hence, the strength of the recall or the mechanism adopted by the BCT does not move this target back to long-run equilibrium. Hence, we reject the existence of an adjustment mechanism for the target of the inflation rate in relation to the equilibrium and we abandon the estimation of the VECM model and the matrix of long-term adjustments.

**3.3. Performance of inflation targeting policy in Tunisia: Validation by Markov chains**

Models with Markovian regime changes were taken from the statistical literature of Hamilton (1989), in order to take into account the non-stationary of some economic and financial series. Having pointed out that this type of series often presents breaks in their means, the original idea of Hamilton (1989) was to model this nonstationarity using a linear process per states. In particular, we assume that the observed series can be treated using an autoregressive process whose parameters vary over time. In addition, Hamilton (1989) hypothesized that the evolution of these parameters is governed by an unobserved state variable that can be modeled using a Markov chain at k regimes.

In economics, the unobserved variable denoted (st)t is often assumed to represent the current state of the economy. This variable is commonly modeled by a two-state Markov chain which means that for all t, the variable st takes the value 1 to upturns periods and the value 2 for recession periods[[9]](#footnote-9). The process Xt is defined as a process MS (2), if it satisfies the following conditions, in the case of an AR (p) process:



Where εt is a white noise process of unknown finite variance. We then specify processes MS (2) -AR (p). The above equations can be rewritten in the following form:



The complete representation of the MS (2) -AR (p) process requires the specification of the variable (st) as a two-state Markov chain, i.e., for any t, st depends only on st -1, for i, j = 1,2:



Probabilities Pij denote the probabilities of transition between states. They measure the likelihood of staying in the same regime and moving from one regime to another. We then obtain easily the following equality:



In practice, it is often interesting to determine the unconditional probabilities of being in a specific regime. We can show that (see Hamilton, 1994):



From the measures of the persistence of the regimes of the seriesP11 and P22, we can also estimate the average duration of a regime. Indeed, if let Ui denotes the random variable representing the duration of the process MS in the regime 1. Knowing that the initial regime is 1, then we easily assume that this random variable follows a geometric distribution with parameters (1-P11 ). For all n, we have: 

Consequently, the mean and the variance of the duration of the regime 1 are given by the following equations:



Obviously, an analogous reasoning provides the mean and the variance of the duration of regime 2.

Most macroeconomic series are non-stationary since they have either a linear trend or they are integrated in order one. Therefore, it is necessary to stabilize transform them before fitting a MS model. Thus, we consider the series Yt = log(Xt) – log(Xt-L) where Xt is the original series and L is the degree of smoothing of the series. The transition to logarithms stabilizes the variance of the process. The choice of the degree of smoothing is not without cost. Indeed, a high degree of smoothing makes possible to eliminate the high frequency variations, but at the same time it increases the time required to detect changes in regimes. The optimal degree of smoothing chosen is an annual smoothing (L = 1), because it is the lowest empirical degree that avoids the largest number of false signals at the 5% significance level. The application of MS model to the consumer price index, the industrial production index, the money market rate and the nominal exchange rate requires the following assumptions:

H1: We set the number of regimes: k = 2.

H2: We set the autoregressive order: p = 1.

H3: The probabilities of transition are constant over time.

H4: The conditional distribution density is the normal distribution with an identical variance for each of the two regimes.

H5: The optimal degree of smoothing chosen is L = 1

The estimation of the MS-AR model (1) for each of the five variables aforementioned is based on a new four-state Markov chain, st \*. The results obtained are reported in table 9.

**Table 9. Estimation of the parameters of the MS model to the variables of the performance of the inflation targeting policy**

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
| Parameters | LCPI | LM3 | LIPI | NER | MMA |
|  | 1.65664\* | 0.9358\* | 0.0670\* | -0.2110 | 0.0097 |
|  | -0.0050\* | 0.0528\* | 0.0670\* | -0.1802 | -0.1285\*\*\* |
| P11 | 0.8625\* | 0.4487\* | 0.3103\* | 0.3849\* | 0.3015 |
| P22 | 0.2849\* | 0.9977\* | 0.8834\* | 0.1225\* | 0.2330 |
|  | 0.8074\* | 0.1322\* | 0.7935\* | 0.2125\*  | 0.0325 |
|  | -0.1341\* |  0.1822\* | -0.5917\* | -0.2273\* | 0.0750 |
|  | 0.1476\* | 0.0387\* |  -0.2435\* | 0.0143 | -0.1069 |
|  | 0.2065\* |  0.3522\* | -0.1938\* | -0.1754 | -0.0645 |
|  | 0.0287\* |  0.0163 | 1.2099\* | 0.9644 | 2.5401 |

Notes: \*, \*\*, \*\*\* indicate significance levels at 10, 5 and 1% respectively.

We notice that the probability P11 associated with the low regime is higher than the probability P22 associated with the high regime, which is consistent with the fact that the average duration of a period of recession is longer than that of a period of expansion. We also point out that the two means of regimes are positive both for the differential of the monetary aggregate M3 and the differential of the industrial production index. In contrast, these two means have negative signs for the differential of the nominal exchange rate and contradictory signs for the differentials of the money market rate and the consumer price index. The standard deviations of the estimated parameters P11, P22, μ1, μ2 for these variables are very low, which indicates a stability of the parameters of these variables. Thus, we retain these variables in our recession-detection indicator.

We can use this econometric technique to detect the change of regimes due to the implementation of an inflation targeting policy by Tunisian central bank. In adopting this policy, the monetary authorities can easily identify the different cyclical phases. Thus, we begin by estimating the long-run relationship between the consumer price index and the M3 monetary aggregate, the industrial production index, the nominal exchange rate and the money market rate using the nonlinear least-squares technique. Then, we use the residuals of this estimation to study the change of regimes through the Markov chain. Table 10 reports the results of the Markovian estimate.

**Table 10. Estimation of the long-run relationship by the Markov chain**

|  |  |
| --- | --- |
| Parameters | LCPI |
|  |  -0.7319\* |
|  | -0.1569\*  |
| P11 | 0.2314\* |
| P22 | 0.2099\* |
|  | -0.2239\* |
|  | -0.1665\*\* |
|  | 0.1550\* |
|  | -0.6085\*\*\* |
|  | 0.9692 |

Notes: \*, \*\*, \*\*\* indicate significance levels at 10, 5 and 1% respectively.

As can be seen, table 2 shows that P11 is higher than P22 which confirms the longer duration of the contraction phase compared to the period of expansion. This evidence is consistent with the monetarist theory suggesting that inflation has a beneficial effect on economic growth in the short-run, but it represents a threat to national wealth in the long-run. We also notice that the two means have negative signs and the standard deviations of the estimated parameters are very small, which generate a total significance for the two probabilities of transition, the means and the sigma.

**Conclusion**

The objective of this article was to assess the performance of the inflation targeting policy in Tunisia. We began identifying the main determinants of the performance of this policy namely the monetary aggregate M3, the industrial production index, the nominal exchange rate and the money market rate.

The empirical analyzes using the Dickey-Fuller-Augmented (1981) test revealed that all these variables are not stationary at levels. Using the Perron (1997) test, we found that these variables are not stationary despite the elimination of trends. Therefore, the first differencing of these time series became necessary to stabilize them. Consequently, these variables are integrated in order one.

Several interesting results were obtained from our analysis. First, we estimated the long-run relationship between the consumer price index and the monetary aggregate M3, the industrial production index, the money market rate and the nominal exchange rate. The results of the ordinary least squares technique and the residual stationary tests validated the existence of a co-integration relation. These latter results in a long-run equilibrium in an error correction model (ECM), since the speed of adjustment of the consumer price index is significantly negative.

The results of estimates using Johansen's maximum likelihood technique (1990) are expected, significant and consistent with the economic theory. In addition, this technique permitted to observe the co-integrating vector from the traces and maximum eigenvalues tests and to assess the performance of the inflation targeting policy by the low exogeneity test in a vector error correction model (VECM).

Using the two-step procedure of Engle and Granger (1987), we found that the inflation targeting policy is effective since the coefficient of the lagged LCPI takes a negative and significant sign within an ECM. Although this technique is weak in statistical terms, it still has a good interpretation in economic terms. We corrected this technique by the multivariate method of Johansen (1991). We concluded from the weak exogeneity test that the long-term adjustment does not lead to a long-run equilibrium for the lagged residual of the relation between the inflation rate and its explanatory variables. Hence, the correction of the method of Engle and Granger leads to disequilibrium in the inflation targeting policy.

Then, we studied the cyclical effect of the performance of the inflation targeting policy in Tunisia using two-regime Markov chains. We evaluated this performance by the nonlinear MCO method of the long-run relationship. Our findings revealed that the average duration of a period of recession is longer than that of a period of expansion.

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1. Faculty of Law, Economics and Management of Jendouba, Tunisia [↑](#footnote-ref-1)
2. Higher Institute of Management of Tunis [↑](#footnote-ref-2)
3. Higher Institute of Management of Tunis [↑](#footnote-ref-3)
4. The sacrifice ratio is defined as the cumulative loss in terms of output (percentage of GDP) following a permanent decline in inflation of 1%. [↑](#footnote-ref-4)
5. In Tunisia, the interest rate is indexed to the money market average. [↑](#footnote-ref-5)
6. Since 1999, The Tunisian monetary authorities had decided to target the monetary aggregate M3 instead of M2 as an intermediate objective to control inflation. [↑](#footnote-ref-6)
7. The test was written in RATS language; source: Estima. [↑](#footnote-ref-7)
8. The test of the likelihood ratio permits to determine the optimal number of lags of the vector autoregressive process. The statistic of this test follows a Chi-square distribution with k degrees of freedom. [↑](#footnote-ref-8)
9. However, some authors such as Sichel (1994) argue that MS models with only two regimes lack flexibility to accurately capture evolutions. [↑](#footnote-ref-9)