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Coordination and Strategic Interactions Between Monetary and Fiscal Policies in Tunisia

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ABSTRACT
This paper proposes a new method of explaining the persistence of inflation in Tunisia due to coordination failures arising from the lack of a strategic interaction between monetary and fiscal policies. We have performed an structural vector autoregressive (SVAR) approach to estimate monetary and fiscal policies reaction functions based on monthly data for the period 2001M1-2018M1, and carried out additional estimations using Markov-Switching (MS) model. The results support the argument for the validity of the fiscal theory of the price level in Tunisia (PM/AF). In addition, the results of the MS model showed opposite signs for the policies in the low inflation regime (first regime). Then, the same signs in the high inflation regime (second regime). This leads us to state that in a low inflation regime, the two policies have a strategic substitutability behaviour, which is the case of a coordination between the two policies. Whereas in a highly inflation regime, there is a strategic complementarity between the two policies. Thus, the persistence of inflation in Tunisian, especially after the revolution, can be understood as a lack of coordination between monetary and fiscal policies.

Keywords: Inflation, Monetary Policy, Fiscal Policy, Coordination, SVAR, Markov-Switching.

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Introduction

Coordination and interaction between monetary and fiscal policies are an ongoing subject in macroeconomics because achieving the objectives of macroeconomic policy is significantly determined by the relevance of both policies. The main focus of monetary policy is price stability, whilst fiscal policy deals mainly with debt and output stabilisation. While both the fiscal and monetary authorities pursue their policies according to their objectives, these policies sometimes conflict depending on the state of the economy and their priorities. Under the assumption that there is a tight interaction between fiscal and monetary policies, the macroeconomic effects of each policy are influenced by this interaction [1].

In many studies, coordination is identified as one policy dominating the other [2]. In this light, a large part of the research on policy coordination has focused on determining which policy prevails. If fiscal policy dominates monetary policy, the fiscal variables are defined independently of the debt level and the intertemporal budget constraint. In fact, in a monetary dominance regime, fiscal variables are defined to satisfy the intertemporal budget constraint. Furthermore, coordination can be characterised as a combination of active and passive policies because equilibrium can only be obtained if one policy is active and the other passive [3].

Our study uses the extensive literature that examines the interaction between monetary and fiscal policies in determining inflation dynamics [2-15].

In this paper, we suggest a new method of explaining inflation persistence that is due to coordination failures that arise from the lack of a strategic interaction between monetary and fiscal policies. Furthermore, the lack of coordination between monetary and fiscal policies can lead to explosive dynamics for inflation, output and debt [14]. This scenario

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is not well understood because the lack of policy coordination has not been thoroughly explored in the context of empirical models. To achieve this, we will adopt two models, namely the SVAR model and the Markov-Switching model. The database covers the period 2001M1-2018M1 for the case of Tunisia.

This paper is structured as follows. Section 2 focuses on the empirical literature on the strategic interactions between monetary and fiscal policy. In section 3, we present the econometric methodology and results. Section 4 summarizes the findings and concludes the paper.

Empirical Literature Review

The macroeconomic theory highlights the impact of the interaction of monetary and fiscal policies on inflation and macroeconomic variables. Indeed, regime switching can influence price stability. Leeper provided the first analysis of the interaction between monetary and fiscal policies in the framework of dynamic stochastic general equilibrium (DSGE) [3].

Guo et al. examined the role of the mix of fiscal and monetary policy rules in the determining of inflation dynamics with monetary and fiscal policy reaction functions and Markov switching vector autoregression (MSVAR) methods based on quarterly data for the period 1992-2007 [9]. Their results indicate that fiscal and monetary policies in China can be adequately described by simple rules and that significant regime changes occurred around 1998. Fiscal policy tended to be active and countercyclical in the pre-1998 period, and then became passive and more countercyclical. Thus, a monetary policy was described as passive and pro-cyclical before 1998, and then became active and counter-cyclical. The combination of fiscal and monetary policy rules can better explain inflation dynamics than the monetary policy rule alone. Therefore, price stability not only requires an appropriate monetary policy, but also an appropriate fiscal policy.

Bianchi estimated full estimation of the information of a Markov switching DSGE model with policy rule coefficients that change between three states [16]. The findings reveal that an M-regime was implemented starting in the 1990s, an F-regime was implemented in the 1970s, and an unbounded solution regime was implemented in the 1980s.

Tarawalie et al. examined the level of coordination between fiscal and monetary authorities in the WAMZ countries (Ghambia, Ghana, Guinia, Nigeria, Sierra Leone, and liberia) and its implications for inflation and budget deficit criteria [17]. To attain this objective, they employed the Set Theoretic Approach (STA) and vector autoregressive modeling (VAR) to estimate the level of policy coordination in the area. The empirical analysis uses annual data for the 1980-2011 period. The findings of the theoretical models indicate that the explicit policy coordination scores in the WAMZ countries are less than 50.0%, with Gambia achieving a coordination score of 46.6%, Ghana (34.5), Guinea (31.8), Liberia (37.9), Nigeria (46.6) and Sierra Leone (41.3). Moreover, the WAMZ countries' monetary authorities have tended to adopt relatively more prudent policies than the fiscal authorities, with the exception of Guinea where both policies are at par in terms of prudence. The results of the impulse response also suggest that there is a small response to shocks caused by different variables. The variables take a long time to converge to their long-run equilibrium path.

The work by Bianchi and Melosi is also linked to studies that investigate the effect of monetary authorities decisions on inflation and inflation expectations, as for example Del Negro and Eusepi and Melosi [18-20]. They used a large-scale DSGE model augmented with stochastic volatilities and showed that the real risk of inflation may be higher than it appears. Basically, this risk is related to the way policy makers will behave in the future. At that point, only an immediate change in fiscal and monetary policies can to reduce the inflationary spiral.

Using Bayesian methods, Davig and Doh estimate a Markov-switching New Keynesian model (MSNK) that allows for changes in policy reaction coefficients and shock volatilities [21]. Using US data, they found that more aggressive monetary policy regime was in place post-Volcker disinflation and pre-1970 than in the period of the Great Inflation of the 1970s. Their results also show that a regime of low volatility was in place for most of the sampling period after 1984. Moreover, a move to an aggressive monetary regime or a low volatility regime will shift the weight of the more persistent to the less persistent shocks, leading to lower inflation persistence. The timing of the regimes from the estimated MSNK model yields a statistically significant "low-high-low" inflation persistence pattern which is consistent with empirical reduced form models.

Cevik et al. studied the interactions between fiscal and monetary policy for some former European emerging economies in transition over the period 1995Q1-2010Q4 by applying a Markov switching model [22]. The authors considered the monetary policy rule proposed by Taylor and the fiscal policy rule implied by Davig and Leeper to explain for the interactions between the monetary and fiscal policies [23, 24]. The results suggest that the monetary and fiscal policy rules exhibit properties of change between active and passive regimes. Moreover, the empirical results indicate that the Czech Republic, Estonia, Hungary and Poland followed both active and passive monetary policies, while Slovenia and the Slovak Republic followed passive monetary rules. As regards fiscal policy, Estonia, Hungary, Poland and Slovenia seem to have alternated between active and passive fiscal regimes, while Czech and Slovak fiscal policies can be characterised by a single fiscal regime. The policy mix and the interactions between monetary and fiscal policies give a diverse picture across the countries in the sample.

Abdel-Haleim explored the extent of coordination between monetary and fiscal policies in Egypt over the period (1974-2015) [25]. Quantifying the extent of coordination depends on the appropriate policy mix responding effectively to different shocks. Indeed, the results confirm that coordination between policies in Egypt was absent or weak during the studied period. The lack of coordination stems from high fiscal deficits that put pressure on monetary policy to achieve its price stabilisation objective, despite the moderation of the fiscal dominance of monetary policy through the phasing out of monetisation.

Kliem et al. investigated the impact of the interaction between fiscal and monetary policies on the low-frequency relationship between the fiscal stance and inflation using cross-country data from 1965 to 1999 [26]. In a first step, they contrasted the monetary-fiscal relationship for Germany, the US and Italy with evidence from simple regression models and a time-varying VAR. Thus, they found that the low-frequency relationship between the fiscal stance and inflation is weak during the periods of an independent central bank and a responsible fiscal policy but more pronounced in periods of a nonresponsible fiscal policy and an accommodating monetary authority. As a second step, they use an estimated DSGE model to provide a structural interpretation of the low-frequency measure and to illustrate how fiscal actions affect inflation in the long run. The DSGE results imply that the interaction between monetary and fiscal policies and the accompanying variations in the propagation of structural shocks may explain the changes in the low-frequency relationship between fiscal position and inflation.

Piergallini studied the dynamics of the aggregate equilibrium in a macroeconomic model in which both monetary and fiscal policies are non-linear and coherent with the empirical evidence [2]. The non-linear monetary policy, where the nominal interest rate exhibits an increasing marginal response to inflation, in interaction with the non-linear fiscal policy. Therefore, the primary budget surplus has an increasing marginal response to the debt and results in four equilibria in a state of equilibrium. Thus, each steady state presents in its neighbourhood a pair of "active" / "passive" monetary / fiscal policies in Leeper style. The obtained results showed that with non-linear adjustments of interest rates and primary surpluses of the type documented empirically, neither the monetary nor the fiscal variables are viable for identifying inflation rate.

Bianchi and Ilut have recently applied a New Keynesian model (MSDSGE), to US data, which is similar to the model used by Clarida et al. and Lubik and Schorfheide [27-29]. The results showed that the rise and fall of US inflation can be attributed to a change in the power balance between the monetary and fiscal authorities. When the fiscal authority is the dominant authority, budget imbalances produce persistent increases in inflation and the monetary authority loses its ability to control inflation. As such, the effects of these shocks last as long as agents expect the fiscal authority to dominate in the future. Therefore, if the monetary authority attempts to reduce inflation with no support from the fiscal authority, inflation hardly moves at all. However, as the fiscal authority adapts the central bank's behaviour, the inflation rate falls rapidly, the economy goes into recession and the debt-to-GDP ratio starts to rise.

Marodina and Portugalb studied the non-linearity of exchange rate pass-through for the Brazilian economy during the period of floating exchange rates (2000-2015) with an MS-DSGE model [30]. They found proof of two distinct regimes for exchange rate pass-through and volatility of inflation shocks. Indeed, the existence of regime changes in the volatility of inflation shocks can be linked to the heteroscedasticity of inflation itself (Brunner and Hess, 1993) or to the theoretical arguments (Ball and Cecchetti, 1990). They point out that higher levels of inflation lead to greater volatility and uncertainty about future inflation expectations. This means that unanticipated inflation shocks increase uncertainty about future inflation and lead to higher inflation volatility in subsequent periods.

Jevđović and Milenković empirically checked the prevailing policy regime (monetary versus fiscal dominance) across five emerging European economies (Hungary, Romania, Bulgaria, Serbia and Macedonia) [31]. The results suggest that monetary policy was subordinated to fiscal policy during the analysis period in all examined economies and that the fiscal regime prevailed.

Bianchi and Melosi studied the issue of coordination between monetary and fiscal authorities at the zero bound [14]. The lack of coordination between monetary and fiscal authorities can lead to an explosive inflation dynamics and large output losses. This contribution is different from the Bianchi and Melosi reports in several dimensions [32]. First, they focused on the absence of inflation in the Great Recession recovery. Second, they have explicitly addressed the risks associated with a lack of coordination between monetary and fiscal authorities. Finally, they showed how policymakers can trigger an increase in inflation expectations and stimulate economic activity through a coordinated policy strategy.

Bazzaoui and Nagayasu examined the relationship between fiscal policy and inflation for 44 countries from 1960 to 2020 [33]. The study was conducted using a panel VAR approach while taking into account the different monetary policy frameworks and levels of fiscal space across countries. The results indicate that budget deficits are less likely to cause inflation when monetary policy is based on targeting inflation. Overall, they found little empirical evidence of fiscal policy effect on the price level. In contrast, the choice of monetary policy framework has a significant impact on the sensitivity of inflation to shocks. While the recent pandemic has significantly affected budget deficits around the world, the paper's conclusions are not altered if the year 2020 is excluded from the study.

Chibi et al. analysed the dynamic interaction between monetary and fiscal policies in Algeria during the period 1963-2017 [34]. First, the nature of fiscal policies in Algeria was examined using a structural vector autoregression model. The obtained results show that there is evidence of a non-Ricardian fiscal policy in Algeria (validity of the fiscal theory of the price level). In addition, the paper analyses the interactions between monetary and fiscal policies by applying a Markov regime-switching model to estimate the time-varying parameters of the relationship. The results revealed that monetary and fiscal policies in Algeria interact contractively for most of the period under consideration. With these findings, they identified a game where the fiscal authority plays first (or is active), while the monetary authority behaves passively in determining debt levels relative to the prices given by fiscal policy. This favours fiscal dominance.

Ultama et al. explored the interaction between monetary and fiscal policies during periods of economic turbulence in the United States (external shock) in Indonesia based on the hybrid New Keynesian model (HNK) [15]. The HNK model is estimated using the Full Information Maximum Likelihood and time-series data over the period 2001Q1-2014Q4. The results showed that the form of coordination is a policy mix between active monetary policy and passive fiscal policy (AM/PF). The degree of coordination decreases with increasing external shock. The study found no significant coordination relationship between interest rates and government fiscal policy. The study shows that HNK is applicable in Indonesia so that based on the theoretical framework, the model used is valid. The study shows that the US shock positively affected Indonesia, therefore it is reasonable to apply the AM/PF policy mix to achieve an effective policy mix.

Econometric Methodology and Results

Description of Markov-Switching Model

The use of Markov-Switching model has become increasingly popular in economic studies of monetary and fiscal policies. Nevertheless, the Markov-Switching model was developed to determine the regimes adopted by the respective political agents. Hamilton argues that the Markov-switching method assumes an endogenous transition from one regime to another, which implies that the political regimes are determined in the model [35]. It is logical to limit the number of regimes to two that are defined as sustainable or unsustainable [36].

Therefore, our objective is to predict two possible regimes for both the effects of the interest rate (monetary policy) and the budget deficit (fiscal policy) on inflation in Tunisia.To this end, two main methods of estimating transition probabilities can be identified in the case of the Markov-Switching process, namely, the fixed transition probability (FTP) and the time-varying transition probability (TVTP).

We have adopted the second method (TVTP) which is a Filardo extended version of the fixed transition probability (FTP) approach that allows transition probabilities to vary over time [37]. In line with Filardo, we let the transition probabilities vary over time and depend on monetary and fiscal policies variables [37]. In addition, we were inspired by the model proposed by Ball to provide a new interpretation of the two regimes arising from the analysis of the inflation process using Markov-Switching (MS) models [38, 39]. In this way, the regimes could be associated with two types of decision-makers or two types of preferences, as in the Ball's model. We believe that this is a further theoretical argument for the use of MS models in inflation modelling.

The Logistic function for the specification of transition probabilities was firstly established by Filardo and later extended by Filardo, Gray, Beine et al. And Isogai et al. [37, 40-43]:

$$|p_{i,j,t} = Pr[S_t = j | S_{t-1} = i] = \frac{exp(\lambda_{i,j,0} + Z'_{t-1}\lambda_{i,j,1})}{1 + exp(\lambda_{i,j,0} + Z'_{t-1}\lambda_{i,j,1})} \quad (1)$$

Where, $i = 1,2,\cdots,M \,;\, j = 1,2,\cdots,M-1$

And $p_{i,M,t} = Pr[S_t = M | S_{t-1} = i, Z_{t-1}] = 1 - \sum_{j=1}^{M-1} p_{i,j,t} \,,\, i = 1,2,\cdots M$ (2)

Thus, we define the unobserved discrete Marcov variable of first order S_t,, which takes two possible values S_t^π=[1,2] and serves as an indicator of the state of the economy in period t. M is the number of regimes and Z_t is the vector of economic variables which explain the transition between regimes. According to Filardo, the probabilities varying over time for two regimes are described below [37]:

$$Pr[S_t = 1 | S_{t-1} = 1] = \frac{exp\left(\lambda_{10} + \sum_{j=1}^{n} Z'_{t-1}\,\lambda_{1j}\right)}{1 + exp\left(\lambda_{10} + \sum_{j=1}^{n} Z'_{t-1}\,\lambda_{1j}\right)} \quad (3)$$

$$Pr[S_t = 2 | S_{t-1} = 2] = \frac{exp\left(\lambda_{20} + \sum_{j=1}^{n} Z'_{t-1}\,\lambda_{2j}\right)}{1 + exp\left(\lambda_{20} + \sum_{j=1}^{n} Z'_{t-1}\,\lambda_{2j}\right)} \quad (4)$$

The state variable not observed in the model of monetary and fiscal policies rules, st, moves in a first-order Markov switching process as described in Hamilton [39]:

$P[S_t=1|S_{t-1}=1]=p_{11}$

$P[S_t=1|S_{t-1}=2]=1-p_{11}$

P[S_t=2|S_(t-1)=2]=p_(22) (5)

$P[S_t=2|S_{t-1}=1]=1-p_{22}$

$0<p_{11}<1 \,;\, 0<p_{22}<1$

Where p_(11) is the probability of remaining in a low inflation regime, given that the previous regime is characterised by low and stable inflation, and p_(22)is the probability of a high inflation regime preceded by high and volatile inflation. $1-p_{11}$ And $1-p_{22}$ are the fixed transition probabilities of being respectively in the first or second state as suggested by Hamilton and discussed by Kim and Nelson (1999) [39, 44]. It should be noted that the average duration in a regime can also be calculated as $d=1⁄(1-p_{ii}$ [22].

Franses and van Dijk noted that for the $p_{i,j}$, S defines the appropriate probabilities [45]. They have to be non-negative but should also require that $p_{11}+p_{12}=1$ and $p_{21}+p_{22}=1$. The transition probabilities can be written in the form of a transition probability matrix (TPM):

$$\Pi = \begin{bmatrix} p_{11} & 1 - p_{11} \\ 1 - p_{22} & p_{22} \end{bmatrix} (6)$$

However, it is important to note that the transition probabilities from one regime to an other are affected by the λ coefficients (equations 3 and 4). Specifically, the signs of these coefficients are important to consider. For instance, when the coefficient $\lambda_{11}$is positive, the relevant economic fundamental Z_t is a significant factor in raising the probability of remaining in a disinflationary regime (regime 1). But if the coefficient is negative, the associated macroeconomic variable Z_t decreases the probability of staying in a regime of low inflation and raises the probability of overbalancing towards high and volatile inflation (regime 2). Likewise, the coefficient λ_{21} captures the effects of the exogenous variable Z_t on the probability of remaining in a high-inflation regime (regime 2) and potentially on the probability of moving to a low-inflation regime (regime 1) according to whether this coefficient is positive or negative, respectively [46].

Data Analysis

We include six observable variables covering the sample from 2001M1-2018M1: inflation rate (CPI), gross domestic product per capita (GDPC), interest rate (IR), money supply growth (MS), budget deficit (BD) and government debt (GD). All the data used correspond to the statistics of the Central Bank of Tunisia except the data concerning the gross domestic product per capita and the growth of money supply, which come from the World Bank. Indeed, all annual variables except CPI and IR are converted to monthly data using the Cubic Spline Interpolation [47]. The data set includes seasonally adjusted data. Seasonally adjusted series (using the X-12-ARIMA monthly seasonal adjustment method) include an automatic ARIMA model selection procedure based largely on the procedure of Gomez and Maravall performed in TRAMO (1996) [48]. The X-12-ARIMA programme is an improved version of the X-11-ARIMA programme, which provides even more comprehensive tools for detecting and addressing seasonal and timing adjustment problems [49].

Furthermore, the interest rate and money growth are considered as indicators of monetary policy, as well as the budget deficit and government debt which are used to guide fiscal policy [26, 9, 50, 51].

The descriptive statistics for the main variables in the sample are presented in Table 1.

Table 1. Descriptive Statistics

Variables	Units	Mean	Standard Deviation	Min	Max	Observations
CPI	%	123.175	11.757	100.000	145.550	205
GDPC	Dollars	3896.837	419.467	3088.500	4402.000	205
MS	%	9.441	2.978	3.110	18.370	205
IR	%	4.800	0.633	3.500	5.870	205
BD	%	3.761	1.600	1.000	6.900	205
GD	%	53.202	9.273	40.700	76.700	205

This table (Table 1) shows that the CPI exhibited clear variations from the sample over the period 2001M1-2018M1, with an average of 123.175 and a standard deviation of 11.757. The mean GDP per capita is equal to 3896.837 USD. It has a standard deviation of 419.4678, with a minimum of 3088.500 USD and a maximum of 4402 USD. Money supply growth and interest rate are very volatile. The averages of MS and IR are 9.441% and 4.800% respectively, while their standard deviation is 2.978 and 0.633, respectively. In fact, the Tunisian government's fiscal situation is relatively worrying for the period under consideration. The budget deficit is on average 3.761% with a standard deviation of 1.600 and government debt is on average 53.202%, with a standard deviation of 9.273.

Discussion of the Results

Preliminary Tests and SVAR Modelling

Before analysing the supposed regime change character of monetary and fiscal policy interactions in Tunisia, we first conducted preliminary empirical research on the nature of policy in Tunisia, using an SVAR model. In recent years, structural vector autoregression models (SVAR) have become a powerful tool for analysing the monetary transmission mechanism and the sources of business cycle fluctuations [53]. This approach was first employed by Blanchard and Watson (1986), Bernanke (1986) and Sims (1986), notably in the US macroeconometric literature [54-56].

Similarly, the adoption of an SVAR model captures the dynamic interdependence of macroeconomic aggregates within a linear model, where each variable's value is expressed in terms of its own past values and an error term. Consequently, SVAR models are explicit about the contemporaneous relationships between variables in order to ensure identification [57].

The SVAR models are well known as embodying an identification problem, in which additional restrictions $((n^2-n))/2$ derived from the theory must be imposed to recover the SVAR specification. According to Sims, Bernanke and Blanchard and Quah the structural disturbance orthogonalization, which is a common practice in applied macroeconomics, or contemporaneous cross-equation effects, are manifested through the contemporaneous effects matrix among different variables [55, 56, 58]. Hence, the cross-equation error variance-covariance matrix Σ of structural shocks is needed diagonally [59].

There are several ways in which restrictions can be placed in the process of identification. It should be noted, though, that setting arbitrary restrictions such as the Choleski decomposition will distort the predicted dynamic behaviour of the system and may at best lead to erroneous results. Otherwise, the restrictions must be driven by economic theory or so-called 'a priori restrictions' in the literature. These fall into one of three categories: (1) Sims and Bernanke short-run restrictions, also known as contemporary restrictions; (2) Blanchard and Quah long-run restrictions; and (3) Gali type restrictions, which represent a combination of short- and long-run restrictions [55, 56, 58-60].

Indeed, an SVAR model for λt, an observable vector of n variables, can be written as:

$$A(L)y_t=A_0(I_n-B_1L - B_2L^2...B_p L^p)y_t = A_0e_t = Bu_t \quad (7)$$

Where u_t is an nxn column vector of structural shocks; e_t is an n-column vector of reduced-form shocks; A_0 an nxn matrix of contemporaneous effects between variables; and A(L) is a matrix lag polynomials such that :

$$A(L) = A_0 - \sum_{k=1}^{P} A_k L^k \quad (8)$$

We introduced the SVAR model with a vector of six endogenous variables, $y_t=(IR_t,CPI_t,GDPC_t,MS_t,BD_t,GD_t)$,, where IR_t denotes interest rate; CPI_t, consumer price index (Inflation); $GDPC_t$, the real GDP per capita; MS_t, money supply growth; BD_t, budget deficit as % of GDP; GD_t, government debt as % of GDP.

The objective of our preliminary analysis is to determine if the assumptions of the fiscal theory of price level determination are valid or invalid for Tunisia. We adopt an approach in the spirit of Dungey and Fry, Haug et al., Mahfoudh, Cazacu, Coric et al., Afonso and Gonçalves and Chibi et al. [61-67].

Before the SVAR estimation begins, we test the stationarity of the variables using the Perron and Zivot-Andrews unit root tests. One problem for time series observations is the possibility of non-stationary data. If the time series data are not stationary, it will give biased results. As a result, the unit root test must be applied before estimation in order to see whether the time series data are stationary or not.

In our work we adopted the tests of Perron and Zivot-Andrews. Perron and Rappoport and Reichlin were the first to emphasise the importance of structural breaks in the performance and interpretation of unit root tests [68, 69]. Perron suggested that the results of unit root tests may be affected by structural changes in the time series [68].

The unit root test results based on the level data and the first difference data should be presented in Table 2.

Table 2: Results of the Perron and Zivot-Andrews Tests

Tests	Perron									Zivot-Andrews								
Models	**Structural break in the intercept**			**Structural break in the trend**			**Structural break in the intercept and trend**			**Structural break in the intercept**			**Structural break in the trend**			**Structural break in the intercept and trend**		
Variables	Date	t-Stat	CV (5%)	Date	t-Stat	CV (5%)	Date	t-Stat	CV (5%)	Date	t-Stat	CV (5%)	Date	t-Stat	CV (5%)	Date		CV (5%)
CPI	2009 M1	-3.07	-5.23	2014 M1	-3.23	-4.83	2009 M1	-2.71	-5.59	2009 M2	-2.10	-4.93	2013 M1	-3.21	-4.42	2009 M2	-3.74	-5.08
GDPC	2005 M1	-3.15	-5.23	2008 M6	-4.60	-4.83	2008 M2	-4.87	-5.59	2005 M2	-3.19	-4.93	2007 M2	-4.04	-4.42	2006 M2	-4.84	-5.08
MS	2003 M9	-4.50	-5.23	2006 M5	-4.31	-4.83	2005 M12	-4.90	-5.59	-	-	-	2004 M2	-4.11	-4.42	2006 M1	-4.96	-5.08
IR	2013 M12	-2.66	-5.23	2003 M3	-2.69	-4.83	-	-	-	2014 M1	-2.67	-4.93	2012 M4	-2.83	-4.42	2011 M6	-4.18	-5.08
BD	2010 M1	-4.96	-5.23	2009 M1	-3.48	-4.83	-	-	-	2010 M2	-3.46	-4.93	2007 M7	-3.57	-4.42	2010 M2	-4.62	-5.08
GD	2005 M1	-3.58	-5.23	2011 M3	-3.98	-4.83	2005 M1	-4.65	-5.59	2005 M2	-3.52	-4.93	2009 M3	-3.56	-4.42	2005 M2	-4.66	-5.08
ΔCPI	2010 M1	**-7.82**	-5.23	2009 M9	**-8.41**	-4.83	2010 M1	**-7.07**	-5.59	2010 M2	**-6.67**	-4.93	2009 M3	**-6.47**	-4.42	2010 M2	**-8.91**	-5.08
ΔGDPC	2007 M1	**-6.84**	-5.23	2015 M5	**-6.41**	-4.83	2007 M1	**-8.96**	-5.59	2007 M2	**-9.83**	-4.93	2015 M5	**-9.40**	-4.42	2007 M2	**-8.98**	-5.08
ΔMS	2015 M1	**-5.81**	-5.23	2003 M8	**-6.71**	-4.83	2003 M8	**-7.23**	-5.59	2015 M2	**-6.82**	-4.93	-	-	-	-	-	-
ΔIR	2011 M9	**-8.11**	-5.23	2011 M7	**-6.80**	-4.83	2011 M9	**-8.05**	-5.59	2012 M5	**-7.11**	-4.93	2011 M7	**-6.77**	-4.42	2012 M5	**-7.09**	-5.08
ΔBD	2010 M1	**-6.18**	-5.23	2011 M6	**-4.99**	-4.83	2010 M1	**-6.48**	-5.59	2010 M2	**-6.02**	-4.93	2011 M1	**-4.96**	-4.42	2010 M2	**-6.25**	-5.08
ΔGD	2010 M1	**-5.89**	-5.23	2005 M8	**-5.57**	-4.83	2006 M1	**-7.30**	-5.59	2010 M2	**-5.90**	-4.93	2005 M3	**-5.77**	-4.42	2006 M2	**-7.32**	-5.08

Notes: CV = critical value, Δ= is the first difference operator.

The results indicate that the variables are stationary at their first differences (the calculated statistical value < the critical value). Therefore, we determine the appropriate number of lags. In addition, we use the Akaike information criterion (AIC), the Schwarz Bayesian criterion (SBC) and the Hannan-Quinn information criterion (HQ).

Table 3: Choice of Lag Order

Lag order	AIC	SBC	HQ
0	-7.11	-7.01	-7.07
1	-39.07	-38.37	-38.78
2	-45.57*	-44.27*	-45.04*
3	-45.32	-43.42	-44.55
4	-45.10	-42.60	-44.08

Notes : AIC : Akaike information criterion ; SBC : Schwarz Bayesian criterion ; HQ : Hannan-Quinn information criterion

Table 3 shows that the three selection criteria lead us to retain the number of lag 2. Concerning the Trace cointegration test, it allows to see if there are cointegrating relationships between the variables in the model.

Table 4: Trace Test Results

Nombre of cointegration (r)	Eigenvalue	Trace Test	Critical Value (5%)	Probability
0	0.158249	83.43961	101.7752	0.0014
1	0.138249	60.06141	66.97645	0.1016
2	0.073552	36.92110	40.17493	0.1024
3	0.056182	21.48884	24.27596	0.1079
4	0.030944	9.808785	12.32090	0.1272
5	0.016979	3.459270	4.129906	0.1746

The findings in Table 4 suggest that the null hypothesis of no-cointegration (83.43961<101.7752) should be accepted at the 5% threshold. In this case, the appropriate modelling methodology is the VAR or SVAR model.

To this end, we try to estimate the policy reaction function of the central bank and the Tunisian government. We, therefore, identify the main determinants of the two interaction decisions, namely the interest rate equation and the budget deficit equation.

To identify the structural parameters, we impose the following short-run zero-value contemporaneous restrictions for $A_0e_t = Bu_t$:

$$A = \begin{pmatrix} 1 & C_{12} & C_{13} & C_{14} & C_{15} & C_{16} \\ 0 & 1 & 0 & 0 & C_{25} & C_{26} \\ 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & C_{56} \\ 0 & 0 & 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} e_t^{IR} \\ e_t^{CPI} \\ e_t^{GDPC} \\ e_t^{MS} \\ e_t^{BD} \\ e_t^{GD} \end{pmatrix}$$

$$B = \begin{pmatrix} C_{11} & 0 & 0 & 0 & 0 & 0 \\ 0 & C_{22} & 0 & 0 & 0 & 0 \\ 0 & 0 & C_{33} & 0 & 0 & 0 \\ 0 & 0 & 0 & C_{44} & 0 & 0 \\ 0 & 0 & 0 & 0 & C_{55} & 0 \\ 0 & 0 & 0 & 0 & 0 & C_{66} \end{pmatrix} \begin{pmatrix} u_t^{IR} \\ u_t^{CPI} \\ u_t^{GDPC} \\ u_t^{MS} \\ u_t^{BD} \\ u_t^{GD} \end{pmatrix}$$

The restrictions mentioned above are generated by an ordinary theoretical intuition about the relationships between variables within a period of three months:

- The zero restrictions in the first four rows and columns of A and B are standard restrictions in the literature on monetary SVARs [70, 71, 62]. For example, if the central bank modifies the interest rate, this impact is first transmitted to the money market, affecting the reserves of commercial banks and their ability to grant credit, thereby spreading further through the economy. This process usually takes more than three months. In the short term, meanwhile, prices are sticky. Moreover, an inflationary shock does not influence the level of GDP in the three-month interval. These restrictions are in line with well-known models (IS curve, Philips curve, Taylor rule, etc.) [64]. Also, some additional restrictions are necessary.
- The contemporaneous response of both variables of fiscal policy to a GDP innovation, inflation rate and the interest rate are set at 0, as it usually takes the government more than three months for new measures to be approved and implemented and for the decision process to pass the legislative body [72, 73, 64].

The estimation results of the monetary policy reaction function (Table 5) indicate that only inflation (-2.993) and GDP per capita (-4.898) significantly determine the interest rate. This means, during this period, monetary policy did not respond to fiscal policy. Hence, the growth of the money supply and government debt did not play a role in determining monetary policy. During this period, the Tunisian central bank seems to be more concerned about inflation and the level of production than anything else. So, there is a systematic relationship between the interest rate and the inflation rate. The coefficient of inflation is well below unity (-2.993). Consequently, monetary policy in this period tended to be passive.

In contrast, the only important variable in determining fiscal policy is government debt. Over this period, fiscal policy was characterised by a relatively strong and significant response to government debt (the debt ratio is -9.904 (below 0). This indicates that fiscal policy during this period was active. Thus, we can conclude that Tunisia adopts a FTPL regime (PM\AF) which validates the results found by Mahfoudh and the results of Chibi et al. for the case of Algeria and contradicts the results of Ultama et al. for the case of Indonesia (AM/PF) [63, 34, 15]. This is the typical case in most countries. This regime implies that the monetary authorities will not react to inflation and the fiscal authorities will not respond to the level of debt. It also assumes that prices are completely flexible. That is, inflation does not depend only on the monetary policy, but the level of prices is determined by the fiscal policy [27].

As regards the interaction of monetary and fiscal policies, the Tunisian central bank has not considered fiscal policy in the conduct of monetary policy. It is supported by the coefficient on the budget deficit (0.083) which is statistically insignificant in determining interest rate.

Table 5: Estimation of the policy reaction function of the central bank and the government by the SVAR model

Structure of contemporaneous relations matrix						
Variables / Shocks	**Monetary policy IR**	CPI	GDPC	MS	Fiscal policy BD	GD
IR	1	0	0	0	0	0
CPI	**-2.993*** **	1	0	0	0	0
GDPC	**-4.898***	0	1	0	0	0
MS	0.039	0	0	1	0	0
BD	0.083	**-0.067*** **	0	0	1	0
GD	0.148	**-0.278** **	0	0		1
Log Likelihood = 4612.934						
LR test for over-identification: Chi-square (7) 127.7058 Prob. 0.000 Correlation LM test : 7.732 Prob. 1.000						

Notes: *, ** and * are significant at the 1%, 5% and 10% level respectively**

In our analysis, we also focus on the impact of monetary and fiscal policies on the evolution of these six variables in order to study the effectiveness of these economic policies in Tunisia. We also aim to find out from their behaviour whether they are complementary or substitutable. Then, the impulse response functions can be calculated from the estimated SVAR models (structural decomposition of innovations). These models show the effects of the shocks to analyse. Figures 1 and 2 show the impulse response functions for monetary and fiscal shocks.

According to Figure 1 a very restrictive monetary policy has a positive effect on interest rate (IR). Monetary policy shocks exhibit the traditional interest rate mechanism, i.e. an increase in the IR leads to a decrease in real output, although these results are not statistically significant (Keynesian monetary policy). This contradicts Sargent and Wallace's (1975) fundamental assumption that monetary policy is ineffective in stabilising output.

We also note that, following the increase in IR, there is a non-instantaneous and insignificant decrease in the rate of inflation up to month 15. After this period, there is a price puzzle that emerges from many attempts to identify exogenous changes in the monetary policy that is well documented by Sims, Christiano and Eichenbaum and Hanson [74-76]. This means that an increase in interest rates is often accompanied by higher inflation rather than lower inflation, as many theories predict [74]. The effect of this shock will not disappear during twenty months. Therefore, we can conclude that the restrictive monetary policy leads to a short-term inflationary effect on the economic life in Tunisia. In this regard, Afonso and Gonçalves also found the same result for the case of the European Union [66].

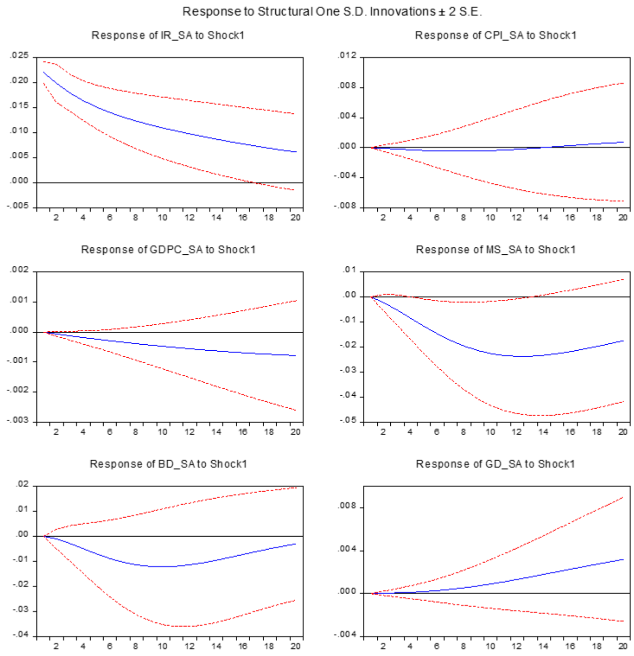

Figure 1: The estimated impulse response functions of a monetary shock: IR shock (Shock 1)

In figure 2, fiscal policy shocks show Keynesian effects, i.e. an expansionary fiscal policy to mitigate the socio-economic impact of the revolution was followed by an increase in real output. Another interesting result is that inflation reacts positively to fiscal policy shocks for 14 months. Then the response becomes negative. This could indicate a mechanism worthy of further study regarding fiscal activity and prices. This policy has generated more or less strong growth (not significant) but has unfortunately aggravated inflation which reached 7.8% in 2018. These results confirm the findings of Ben Slimane and Ben Taher for Tunisia [77].

However, even in this regime (PM/AF), in general, inflation dynamics depend on both monetary and fiscal policy stances, even if the monetary policy is passive.

With respect to fiscal and monetary interactions, a (+) or (-) sign indicates either complementarity or substitutability in the response of the fiscal or monetary policy to a shock. A double sign denotes a non-monotonic response; i.e. +/- shows that there is complementarity after an initial delay, followed by substitutability [78]. In our case, the IR responds positively and significantly to the BD until the second month, then the impact becomes negative for the rest of the period. That is, both fiscal and monetary policies seem to move in the same direction and then in the opposite direction. There is therefore evidence of complementarity and substitutability between fiscal and monetary policies in Tunisia. Overall, it could be said that substitutability between policies dominates complementarity, which is in line with the results of Büyükbaşaran et al. [1].

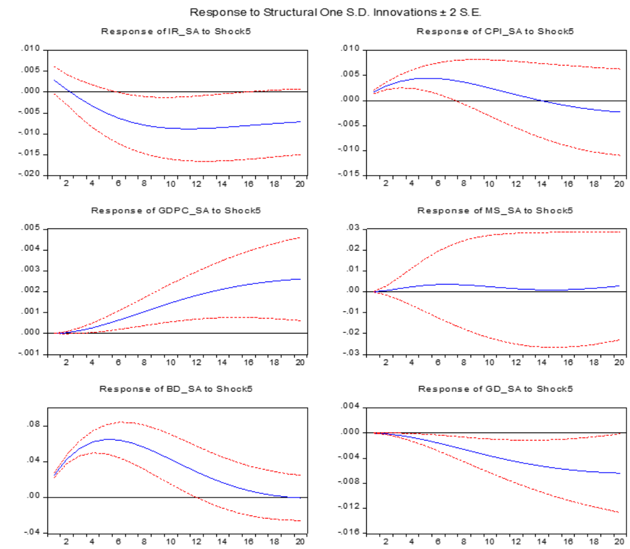

Figure 2: The estimated impulse response functions of a fiscal shock: BD shock (Shock 5)

Nevertheless, the transmission of shocks from these variables to the two policies is better discerned by using variance decomposition analysis.

Table 6: The Variance Decomposition of the Two Policies

Period	Monetary policy: IR						Fiscal policy: BD					
	Shock1 IR	Shock2 CPI	Shock3 GDPC	Shock4 MS	Shock5 BD	Shock6 GD	Shock1 IR	Shock2 CPI	Shock3 GDPC	Shock4 MS	Shock5 BD	Shock6 GD
1	87.52	9.08	0.97	0.00	1.99	0.42	0.00	0.00	0.00	0.00	73.53	26.46
5	82.42	14.03	0.77	0.45	2.02	0.28	0.42	0.83	0.72	0.50	70.94	27.01
10	70.45	17.62	2.39	0.37	8.57	0.58	1.52	0.66	3.84	2.64	64.66	26.65
15	61.89	20.32	3.31	0.74	13.22	0.48	2.20	1.23	7.52	5.92	57.56	25.54
20	55.89	**23.29**	3.38	1.58	**15.28**	0.55	2.21	7.64	**8.32**	8.65	49.98	23.26

From the above table, we can observe that the interest rate variation came by far from an expansionary fiscal policy (15.28%), which confirms that the expansionary fiscal policy in Tunisia provokes a restrictive monetary policy after almost two years, which explains the substitutability between the two policies. We also note that after two years, the shock of private demand (inflation) leads to a variation of 23.29% of the interest rate, which is why this rate is a reduced efficiency tool for the Tunisian monetary authorities to control inflation. As regards fiscal policy, it can be seen that a shock to real output causes a variation of 8.32% in the budget deficit.

Markov-Switching Model Estimation Results

This part also focuses on the identification of macroeconomic variables associated with monetary and fiscal policies in order to highlight the mechanisms underlying inflation dynamics in Tunisia along the study period. To this end, we consider a set of five explanatory variables to explain the changes in inflation regimes. We focus, on macroeconomic policy instruments.

The results of the regression of the Marcov-Switching-TVTP approach based on different variables are presented in Table 7. This table provides the empirical effects of these variables on the inflation dynamics.

The specifications of the transition probabilities presented in the previous subsections are as follows:

$$Pr[S_t = 1|S_{t-1} = 1] = \frac{exp(\lambda_{10} + \lambda_{11}LTI_{t-1} + \lambda_{12}LDB_{t-1} + \lambda_{13}LPIBH_{t-1} + \lambda_{14}LMM_{t-1} + \lambda_{15}LDG_{t-1})}{1 + exp(\lambda_{10} + \lambda_{11}LTI_{t-1} + \lambda_{12}LDB_{t-1} + \lambda_{13}LPIBH_{t-1} + \lambda_{14}LMM_{t-1} + \lambda_{15}LDG_{t-1})} \quad (9)$$

$$Pr[S_t = 2|S_{t-1} = 2] = \frac{exp(\lambda_{20} + \lambda_{21}LTI_{t-1} + \lambda_{22}LDB_{t-1} + \lambda_{23}LPIBH_{t-1} + \lambda_{24}LMM_{t-1} + \lambda_{25}LDG_{t-1})}{1 + exp(\lambda_{20} + \lambda_{21}LTI_{t-1} + \lambda_{22}LDB_{t-1} + \lambda_{23}LPIBH_{t-1} + \lambda_{24}LMM_{t-1} + \lambda_{25}LDG_{t-1})} \quad (10)$$

Table 7: The Estimation Results of the MS-TVTP Model

Variables indépendantes	MS-TVTP model: Dependent variable: CPI					
	Regime 1 (Low inflation)			Régime 2 (High inflation)		
	Coefficients	SD	Pr(>|t|)	Coefficients	SD	Pr(>|t|)
C	-0.4963	0.5324	0 .3512	-0.6167	0.1274	1.2940
IRt-1	-0.0524	0.0134	9.214***	0.4351	0.0367	2.2000***
BDt-1	-0.0094	0.0103	0.3615	0.0024	0.2330	0.81576
GDPCt-1	0.3748	0.0272	2.2000***	0.5333	0.0106	2.2000***
MSt-1	-0.0258	0.0041	3.1200***	0.0864	0.0091	2.2000***
GDt-1	0.5785	0.0728	1.998***	0.0489	0.0218	0.0248
TPM : Regime 1 Regime 2	0.98885 0.01115			0.01813 0.98187		
Average duration (d)	89.686			55.157		
Number of observations	104			101		
AIC	-971.2766					
BIC	-867.5243					
LogLik	479.6383					
R^2	0.9134			0.8834		

Notes : d=1⁄(1-p_ii) ; *** 5% (1.96).

Firstly, this table shows that the data for all specifications can be divided into two main regimes. The first is a low inflation state with low volatility, and the second is a high inflation state with high volatility. We also note that the monthly inflation rate estimated with an inference based on TVTP switching between two different regimes is negative. Thus, the coefficient of the constant (C=-0.4963%) for the first regime is negative but insignificant. As for the second regime it is equal to -0.6167%, which is negative and insignificant (the t-statistics were respectively 0.3512 and 1.2940 which are lower than the critical value of 5% (1.96)).

However, the probability of regime 1 being followed by regime 1 is 0.98885, and the probability of regime 2 being followed by regime 2 is 0.98187. Therefore, both regimes are very persistent. Hence, the probability of the country moving from a "high inflation" state to a "low inflation" state (p_{21}=0.01813) is higher than the probability of moving from a low inflation regime to a high inflation regime (p_{12}=0.01115). Another interesting result is the expected duration of a state, which was estimated as the inverse of the transition probability. The expected duration of a period of low inflation was estimated to be 89.7 months while that of high inflation was about 55.1 months. This implies that, on average, once Tunisian economy enters a period of "high" inflation, it will remain in this state for about 55 months. The duration of low inflation is estimated to be around 90 months. Overall, the relatively low probability of moving from a low to a high inflation regime and the duration of a low inflation regime strengthen the effectiveness of macroeconomic policy in managing inflation in Tunisia.

Thus, based on the results presented in Table 7, the price level decreases in response to a decrease in the interest rate. In fact, the coefficient of the interest rate indexed by regime 1 is negative (λ_{11} = -0.0524) and significant with a t-statistic of 9.214, which is well above the critical value of 5% (1.96). This suggests that the lower the interest rate, the higher the probability of remaining in a low inflation regime. Similarly, for regime 2, an increase in the interest rate increases the probability of remaining in a high inflation regime (λ_{21}= 0.4351). This supports the price puzzle identified by Sims, Christiano and Eichenbaum and Hanson [74-76]. Two explanations can justify this price puzzle for Tunisia which is either the effectiveness of the IR in controlling inflation is reduced or the IR is a wrong indicator of monetary policy. Our results call into question those found by Khemiri and Ben Ali, for the case of Tunisia [46]. Therefore, we resorted to the growth of the money supply as a tool of the Tunisian central bank. Thus, we found, a negative and significant sign for the money supply coefficient (λ_{13}= -0.0258). This shows that the slowdown in money supply growth reduces inflation in Tunisia. Our results are consistent with the monetary theory where inflation is a monetary phenomenon and with those found by Ofori et al. [79]. Moreover, in a high inflation regime, the coefficient (λ_{23}= 0.0864) is positive and significant. Therefore, one can say that the increased growth of the money supply leads to higher prices. This contradicts the results of Diermeier and Goecke for the case of European countries [80].

The response of inflation to the gross domestic product per capita (GDPC) was estimated to be (λ_{13}= 0.3748) which is positive and significant. However, any increase in the GDPC of 1% increases the probability of staying in a low inflation regime by 0.37%. Similarly, the rise in the GDPC increases the probability of remaining in a high inflation regime by 0.53% (λ_{23}= 0.5333). We, therefore, concluded that the increase in the GDPC affects inflation regardless of the regime. For example, Omay et al. and Khemiri and Ben Ali indicate that inflation is affected by the industrial production index via the nominal uncertainty channel [81, 46]. Moreover, several studies show no statistically significant impact of the output gap on inflation, namely Arruda et al. and Sachsida [82, 83].

As regards the response of inflation to fiscal policy, the results showed that it is difficult to establish a relationship between budget deficits and inflation. In this respect, according to Giannitsarou and Scott, widely expected future increases in budget deficits do not necessarily have a substantial impact on inflation [84]. The budget deficit coefficient for both inflation regimes is insignificant although it is negative for the first regime (λ_{12}= -0.0094) and positive for the second regime (λ_{22} = 0.0024). Therefore, the reduction or total elimination of budget deficits can be seen as a stylized fact of successful stabilisation programmes. Our results differ from most of the existing empirical literature on fiscal policy and inflation and are consistent with those of Giannitsarou and Scott and Kwon et al. [84-88]. Consequently, we focus on the role of government debt in determining inflation in Tunisia instead of the budget deficit. The government debt coefficient, positive and significant in regime 1, is equal to (λ_{15}= 0.5785). This implies that government debt plays a major role in determining the price level in Tunisia. Specifically, government debt could affect inflation, as assumed by the FTPL.

It was concluded that price level determination requires not only an appropriate monetary policy regime but also an appropriate fiscal policy regime. These suggest an important role for the combination of fiscal and monetary policy rules in determining inflation dynamics.

Besides, when considering the combination of policy rules, as shown in Figure 3, regime 2 occurred mainly in the periods 2003M1-2009M8 and 2015M7-2018M1, which is consistent with the Tunisian experience.

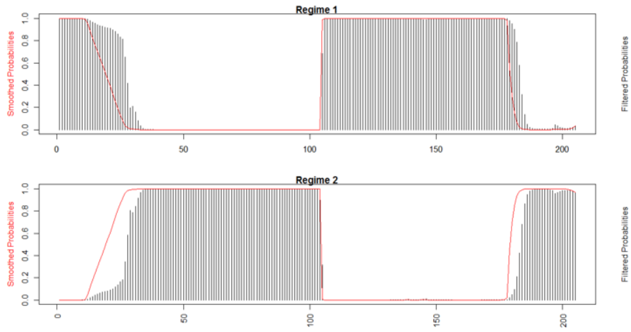

Figure 3: Probabilities of regime 1 and 2 for the Markov-Switching model with fiscal and monetary policy rules, 2001M1-2018M1.

Conclusion

In this paper, we have examined the role of the combination of fiscal and monetary policy rules in determining the dynamics of inflation. The aim is to identify the determining mechanism of price level in Tunisia. More specifically, we estimated the monetary and fiscal policy reaction functions based on monthly data for the period 2001M1-2018M1, and performed additional estimations using the MS model.

Our empirical results on monetary and fiscal policy reaction functions support the argument for the validity of the fiscal theory of the price level in Tunisia (PM\AF). Indeed, the dynamics of inflation according to the SVAR model depend on both monetary and fiscal policy stances, even if monetary policy is passive. Moreover, the monetary and fiscal interaction indicates evidence of complementarity and then strategic substitutability.

On the other hand, a particular analysis was carried out on the introduction of the MS model with two regimes. To thid end, we studied how Tunisian monetary and fiscal policies interacted during the study period. The advantage of this approach over the SVAR models is that it allows to dertermine changes in policies of each regime. However, raising the interest rate would increase the probability of remaining in a high inflation regime, i.e. monetary policy would lose its effectiveness on prices. Consequently, we suggested using the growth of the money supply as an instrument of monetary policy instead of the interest rate. The latter shows that slower money supply growth increases the probability of remaining in a low inflation regime. Also, the findings deny the existence of a relationship between budget deficit and inflation. This highlights the role of government debt in determining inflation in Tunisia. That is, government debt affects inflation and an increase in debt increases the probability of remaining in a low inflation regime.

In general, the results showed opposite signs for policies in the first regime and then the same signs in the second regime. This allows us to argue that in a low inflation regime, the two policies have strategic substitutability, which is the case of coordination between the two policies [89]. In a high inflation regime, however, there is a strategic complementarity between the two policies. Thus, the persistence of inflation in Tunisia, especially after the revolution, can be explained by the lack of coordination between monetary and fiscal policies [90-96].

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