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Abstract

This paper examines the dynamic characteristics of the inflation rate in Tunisia over the last two decades, and particularly following the onset of the Arab Spring in 2010 which causes distortions in this country's monetary policy. We focus on the two specific dimensions of the Tunisian inflation rate: inflation regimes and persistence. We tackle this issue by adopting an evolutionary spectral approach, initially proposed by Priestley and Tong (1973). Our main findings indicate a stable inflation regime in the last ten years, with an average inflation rate of around 5.5%. It is also found that the Tunisian inflation experienced a high degree of inertia which reflects its gradual responses to shocks. We also discuss the policy implications of these results, which typically require policymakers to implement sound institutional reforms to reduce inflation.

Keywords: Inflation, structural break, spectral analysis, inflation persistence.
JEL Classification: C52, E52, E63.

1. Introduction

After the onset of unrest movements in the Middle East and North Africa (MENA) region, which began in December 2010 and have become known as the Arab Spring, the economic conditions in many MENA countries have become worse. In Tunisia, where the first protests occurred on the 17^{th} of December 2010, distortions in monetary and macroeconomic indicators such as the decrease in the economic activity and the devaluation of the real exchange rate led to a growing inflation uncertainty. Since January 14, 2011, the prices of necessary goods including particularly foods and energy rose and were very volatile. This rising volatility of the price level created economic uncertainty and led to reduce economic growth because it typically discourages investment, consumption, and industrial competition. Some policymakers from the Tunisian Central Bank, specialists from the Tunisian ministry of economics and academic researchers from the African Development Banks urged appropriate reforms and actions in order to stabilize inflation and get the country out of the difficult economic situations. The focus was put on the adjustments of interest rate as well as on the liquidity level that the Central Bank should inject into the banking sector. However, the policies would not be efficient without a good understanding of inflation characteristics and forecasting.

From a theoretical point of view, studies such as Svenson (1997, 1999), Clarida et al. (1999), and Stock and Watson (2002) show that the inflation is determined by the output gap, the interest rate with a dynamic system involving an Euler equation, a short Phillips curve, and a central bank loss function. Within this system, a good understanding of inflation dynamics requires, on the one hand, the identification of different regimes of the inflation rate and, on the other hand, the nature of inflation persistence which has been observed as a stylized fact in previous empirical studies. These elements (i.e., inflation regimes and persistence) are critical to inflation forecasting exercise and, therefore, appropriate monetary policy decision-making.

This study analyses the above-mentioned dimensions of the inflation rate in Tunisia as far as the empirical results may provide insights to monetary policies. For this objective, we rely on the evolutionary spectral analysis, initially proposed by Priestley (1965) to gauge the inflation characteristics. We then derive a new time-varying measure of inflation persistence from the evolutionary co-spectral analysis in the sense of Priestley and Tong (1973). The evolutionary spectral analysis is a time-frequency approach, and contrary to time series models, it allows for a representation of nonstationary series without any risk of misspecification. It is also advantageous in that it does not depend on any preliminary modeling and does not have an “end-point problem” since no future information is used, implied or required as in band-pass or trend projection methods.[1] Compared to other alternative decomposition approaches such as the Fourier transformation and wavelets, the most prominent contribution of the evolutionary spectral analysis is effectively the relevant decomposition of time series with respect to two dimensions – a frequency dimension (short-term, medium-term, and long-term horizons) and temporary dimension – without facing problems of periodicity and non-stationarity of frequency components as well as having to use any pre-treatment technique. In fact, the traditional Fourier transformation provides no information about the time evolution of signal’s spectral characteristics, while the windowed Fourier transformation treats a signal under fixed time-frequency window with constant intervals in the time and frequency domains. As a result, the Fourier transformation requires that the time series under consideration is periodic and stationary, and does not evolve over time. This is unrealistic as most economic and financial variables exhibit quite complicated patterns over time such as trends, abrupt changes, and volatility clustering (Fan and Gençay, 2010). On the other hand, the wavelet approach is suitable for capturing the time-varying characteristics of studied variables over different time scales, but its implementation typically depends on the data length and the type of wavelet filters.

At the empirical level, we only focus on the inflation characteristics in the short and medium terms as the monetary authority in Tunisia does not follow long-run objectives. In the related literature, several studies have investigated inflation dynamics in various countries (Pétursson, 2004; Levin et al., 2004; Kontonikas, 2004; Genc et al., 2007; Ftiti and Essaadi, 2008), none of them has addressed this issue in the context of Tunisia with the objective of providing useful information for inflation forecasting. The existing studies are limited to the analysis of the Tunisian monetary policies using a VAR approach and impulse responses functions (e.g., Boughrara and Smida, 2004). It is also worth noting that most previous studies have employed a VAR framework to identify monetary policy shocks and their impacts on the real and financial sectors (e.g., Huh, 1996; Bernanke and al., 1999; Bernanke and Mihov, 1998; Lane and Van Den Heuvel, 1998; Da Silva and Portugal, 2000; Honda, 2000; Jarociński, 2010; Mallick and Sousa, 2011;

1 In traditional time series modeling, empirical methods often require the estimation of the mean and volatility processes (e.g., ARMA, GARCH, etc.) which depend on distributional assumptions. Misspecification can also arise when, for example, a time series is not stationary and subject to structural change.

Cover and Olson, 2013; Bekaert et al., 2013; Calza et al., 2013; Gupta et al., 2013; Fernald et al., 2014).

Using monthly data for the consumer price index (CPI) in Tunisia from July 1987 to December 2014, we find evidence of a stable inflation regime over the last ten years which also cover the recent global financial crisis and the Arab Spring revolutions. The average monthly inflation rate of this stable regime is around 0.39%. We also document that the Tunisian inflation experienced a high degree of inertia which reflects its gradual responses to shocks.

The rest of this paper is organized as follows. Section 2 presents the empirical methodology to identify different regimes of the Tunisian inflation rate and to measure inflation persistence. Section 3 describes the data. Section 4 reports the obtained results. Section 5 concludes the paper.

2. Methodology

Given the objective of our paper, we rely on frequency analysis to identify the inflation regimes in Tunisia as well as measure inflation persistence. Specifically, the time-frequency approach of Priestley (1966, 1996) is used, in the first stage, to examine the dynamics of the Tunisian inflation rate over different time scales (short-term versus medium-term frequencies).[2] Two short-term frequencies (4-month and 1-year horizons) and one medium-term frequency (3-year horizon) are selected for our analysis. We then use the Bai and Perron (2003)'s test to determine the potential of endogenous break points in these frequency components of the inflation rate.

In the second stage, we develop a time-varying measure of inflation persistence in a time-frequency domain based on the co-evolutionary spectral analysis of Priestley and Tong (1973). This stage enables us to discern short-term and medium-term persistence.

2.1 Identification of inflation regimes

Following the evolutionary spectrum theory developed by Priestley (1965, 1996), the time-dependent process of the Tunisian inflation rate can be defined as follows:

$$inf_t = \int_{-\pi}^{\pi} A_{\text{inf}}(w,t)\ e^{iwt}\ dZ_{inf}(w) \qquad (1)$$

where, for each time scale w, the sequence $A_{\text{inf}}(w,t)$ is a time-dependent and nonstationary pro-

[2] The long-run frequency is not considered as the monetary authority in Tunisia does not follow long-run objectives.

cess. If the process under consideration is stationary, the sequence in Eq. (1) is only frequency-dependent and equal to A_{inf}. The sequence $A_{\text{inf}}(w,t)$ has a generalized Fourier transform. The modulus of this sequence has an absolute maximum at the origin. $\{dZ_{inf}(w)\}$ is an orthogonal process on $[-\pi,\pi]$ with $E\{dZ_{inf}(w)\}=0$. Without loss of generality, the evolutionary spectral density of the process inf_t is defined by $h_t(w)$ as follows:

$$h_t(w) = \frac{dH_t(w)}{dw} \tag{2}$$

where, $dH_t(w) = \left|A_{inf}(w,t)\right|^2 d\mu_{\text{inf}}(w)$. The variance $\sigma^2_{inf,t}$ of inf_t at time *t* depends on the evolutionary spectral density $h_t(w)$ through the following equation:

$$\sigma^2_{inf,t} = \text{Var}(inf_{t)} = \int_{-\pi}^{\pi} h_t(w)\ d(w) \tag{3}$$

Following Priestley (1965), the estimation of the evolutionary spectrum is performed by using two windows $\{g_u\}$ and $\{W_v\}$. Without loss of generality, $h_t(w)$ is constructed as follows:

$$\widehat{h_t}(w) = \sum_{v\in Z} w_v\ |u_{t-v}(w)|^2 \quad \text{with} \quad u_t = \sum_{v\in Z} g(u) X_{t-v} e^{-iw(t-v)} \tag{4}$$

where $\{g(u)\}$ and $\{w_v\}$ are defined as follows :

$$g(u) = \begin{cases} \dfrac{1}{2\sqrt{h\,\pi}} & if\ \ |u| \leqslant h \\ 0 & if\ \ |u|\ > h \end{cases}$$

$$w_v = \begin{cases} \dfrac{1}{T'} & if\ \ |v| \leqslant \dfrac{T'}{2} \\ 0 & if\ |v|\ > \dfrac{T'}{2} \end{cases}$$

In this paper, we also opt for $h=7$ and $T'=7$ as in Priestley (1995), Artis et al. (1992) and Ahamada and Boutahar (2002). According to Priestley (1988), we have $E\left(\widehat{h_t}(w)\right) \approx h_t(w)$, $var\left(\widehat{h_t}(w)\right)$ decreases when T' increases for $\forall(t_1,t_2)$ and $\forall(w_1,w_2)$ if at least one of the following conditions *(i)* and *(ii)* is satisfied.

$$(i) \quad |t_1 - t_2|\ \geq T' \qquad\qquad (ii) \quad |w_1 \pm w_2| \geq \frac{\pi}{h}$$

We now turn to investigate whether endogenous structural breaks are present in the time-frequency dynamics of the Tunisian inflation rate under monetary policy shocks. The empirical literature has proposed a number of techniques to detect multiple breakpoints. While some tests allow one to detect a single breakpoint at a specified known break (Chow, 1960; Brown et al., 1997), others enable the identification of structural breaks at unknown break dates (Andrews et al., 1996; Liu et al., 1997; Bai and Perron, 1998, 2003). For instance, Andrews *et al.* (1996) considers multiple structural changes but requires a well-known variance. The Liu and al. (1997)'s test examines multiple unknown breakpoints but considers only the pure structural change case where all parameters are subject to shifts. Bai and Perron (1998, 2003) propose another test focusing on the instability problem over time. The Bai-Perron test is typically based upon an information criterion in the context of a sequential procedure, and allows not only for the identification of multiple structural breakpoints in a linear regression model but also the estimation of their timing. Our study uses the Bai-Perron test to determine possible shifts in the dynamic of inflation series at different time scales according to the following standard linear regression model:

$$y_t = x_t' \beta_j + u_t \quad for \;\; t = T_{j-1} + 1, \dots\dots, T_j \; and \; j = 1, \dots, m+1. \qquad (5)$$

where y_t is the observation of the dependent variable; x_t is a $(k * 1)$ vector of regressor; β_j is $(k * 1)$ vector of regression coefficients; and u_t is the error term. The parameter *m* is the number of breaks.

The Bai Perron model allows coefficients to change over time. Therefore, the null hypothesis is presented as follows:

$$H_0: \beta_i = \beta_0 \quad for \quad i = 1, \dots, n.$$

The null hypothesis consists to test if the regression coefficients remain constant. For the hypothesis H_0, the objective of Bai and Perron test is to estimate the both unknown regression coefficients and the break dates.[3] To do this, some restrictions are imposed on the possible values of the break dates. They define a set for some arbitrary small positive value of ε

$$\lambda_\varepsilon = \{(\lambda_1, \dots, \lambda_m);\; |\lambda_{i+1} - \lambda_i| > \varepsilon, \lambda_1 > \varepsilon, \lambda_m > 1 - \varepsilon\}$$

Bai and Perron (1998, 2003) define the asymptotic critical values of ε which depend on the maximum number of breaks. For example, $\varepsilon = 0.05\ (M = 9), \varepsilon = 0.10\ (M = 8),\ \varepsilon = 0.15\ (M = 5), \varepsilon = 0.2\ (M = 3), \varepsilon = 0.25\ (M = 2)$. In our study, we follow Bai and Perron

[3] Bai-Perron test considers the break points $(T_1, \dots, T_m)$ as unknown and for $i = 1, \dots, m$, and $\lambda_i = \frac{T_i}{T}$ with $0 < \lambda_1 < \cdots < \lambda_m < 1$.

(1998) by imposing $\varepsilon = 0.15$ $(M = 5)$.

2.2 Measure of inflation persistence

Priestley and Tong (1973) propose a time-varying measure of co-spectral density function between a bivariate process $\{X_t, Y_t\}$. In this paper, we suggest an extension of this theory in order to propose a new measure of time-varying autocorrelation function in frequency domain. This new measure is called Evolutionary Auto-Spectral Density Function (EASDF) and obtained by substituting Y_t in the co-spectral density function of Priestley and Tong (1973) by X^d, with d being a delay factor. $X^d = X_{t-i}$ for i = 1,…, q, where q represents the lags of inflation rate we consider when estimating inflation persistence. Formally, X_t and $X^d$are defined as follows:

$$X_t = \int_{-\pi}^{\pi} A_{t,X}\ e^{iwt}\ dZ_X(w)\ and\ X^d = \int_{-\pi}^{\pi} A_{\mathrm{x^d}}\ e^{iwt}\ dZ_{X^d}(w) \qquad (6)$$

with $E[dZ_x(w_1)dZ_X^*(w_2)] = E\left[dZ_{x^d}(w_1)dZ_{X^d}^*(w_2)\right] = E\left[dZ_x(w_1)dZ_{X^d}^*(w_2)\right] = 0$

and for $w_1 = w_2$, we have $E[|dZ_x(w_1)|^2] = d\mu xx(w_1)$; $E\left[\left|dZ_{x^d}(w_1)\right|^2\right] = d\mu x^d x^d(w_1)$; $E\left[dZ_X(w_1)dZ_{X^d}^*(w_1)\right] = d\mu xx^d(w_1)$

By virtue of the Cauchy-Schwarz inequality, we can write that:

$$\left|dH_{t,XX^d}\right|^2 \leq dH_{t,XX}\ dH_{t,X^dX^d} \text{ for all } t \text{ and } w$$

$$dH_{t,XX^d} = h_{t,XX^d}$$

where h_{t,XX^d} may then be termed “the evolutionary auto-spectral density function”. In time domain, this time-varying function is equivalent, according the Fourier transformation, to the dynamic auto-correlation function. In this paper, h_{t,XX^d} represents the time-varying inflation persistence measure.

The estimation of the evolutionary auto-spectral density function needs two filters. For the discrete univariate process, Priestley (1966) gives two relevant windows. These are relevant filters and they have been tested in past studies such as Ahamada and Boutahar (2002), Ftiti (2010) and Bouchouicha and Ftiti (2012). For the discrete bivariate process, Priestley and Tong (1973) propose the same filters:

$$g_u = \begin{cases} \frac{1}{2\sqrt{h\,\pi}} & if\ |u| \le h \\ 0 & \end{cases}$$

$$w_v = \begin{cases} \frac{1}{T'} & if\ |v| \le \frac{T'}{2} \\ 0 & if\ |v| > \frac{T'}{2} \end{cases}$$

Then, the estimation of the evolutionary auto-spectral density function is as follows:

$$\hat{h}_{t,XX^d} = \sum_{v \in Z} W'_T(v) U_x(w, t-v) U_X^d(w, t-v)$$

with

$$U_x(w,t) = \sum_{u \in Z} g(u) X(t-u) e^{iw(t-u)} du$$

$$U_X^d(w,t) = \sum_{u \in Z} g(u) X^d(t-u) e^{iw(t-u)} du$$

We also set $h = 7$ and $T' = 20$ as in Artis et al. (1992), Priestley (1995), Ahamada and Boutahar (2002), Essaadi and Boutahar (2008), and Ftiti and Essaadi (2008).[4] According to Priestley (1988), we have $E(\hat{h}(w)) \approx h_t(w)$, and $Var\left(\hat{h}(w)\right)$ decreases when $T^{'}$ increases for $\forall(t1, t2), \forall(w1, w2), cov(\hat{h}_{t1}(w1), \hat{h}_{t1}(w2)) = 0$, if at least one of the conditions (*i*) or (*ii*) described previously is satisfied.

3. Data

We use monthly data for the Tunisian consumer price index (CPI) from July 1987 to December 2014. The data are extracted from the Datastream database. The inflation rate inf_t is then computed as:[5]

$$inf_t = 100 * \text{Ln}\left(\frac{CPI_t}{CPI_{t-1}}\right)$$

[4] This choice of values is justified by the fact that they respect the conditions (i) and (ii).
[5] The CPI index is seasonally adjusted according the CENSUS method X12 (see, Darne, 2000 for more details).

When carrying out the evolutionary spectral estimation, we lose ten observations at the beginning and at the end of the sample period. Therefore, the estimated evolutionary spectral density function of inflation dynamics will range from June 1988 to December 2014.

Table 1: Dynamics of the Tunisian inflation rate (1987-2014)

Mean (%)	0.323
Maximum (%)	1.146
Minimum (%)	-0.440
Standard deviation	0.243
Skewness	0.488
Kurtosis	3.810
JB	22.095 (0.00)
ADF	-5.881 (0.00)
PP	-14.988 (0.00)
ZA [break date: 2001]	-16.173 (0.00)

Notes: this table reports the descriptive statistics of the Tunisian inflation rate over the study period. JB refers to the Jarque-Bera test for normality. ADF, PP, and ZA are the empirical statistics of the Augmented Dickey-Fuller (1979), Phillips-Perron (1988), and Zivot and Andrews (1992) unit root tests. The ZA test accounts for the potential of structural break occurring in the dynamics of the inflation rate. P-values of the statistical tests are in parenthesis.

Table 2: Autocorrelation function (ACF) of inflation series (1987-2014)

Lag	1	2	3	4	5	6	7	8
ACF	0.988	0.976	0.964	0.952	0.940	0.928	0.916	0.905
Lag	9	10	11	12	13	14	15	16
ACF	0.893	0.882	0.870	0.858	0.847	0.836	0.832	0.814
Lag	17	18	19	20	21	22	23	24
ACF	0.803	0.792	0.781	0.770	0.759	0.748	0.738	0.727
Lag	25	26	27	28				
ACF	0.717	0.707	0.696	0.686				

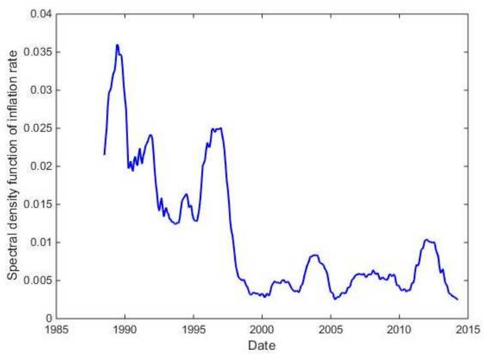
Table 1 shows that the Tunisian inflation rate fluctuates around an average of 0.323% over the study period, with a maximum value of 1.146%. The series is skewed to the right and has a leptokurtic behavior. This evidence of non-normal distribution is confirmed by the results of the Jarque-Bera test for normality. We also examine the stationary property of the inflation rate using three commonly-used unit root tests. The results indicate that the null hypothesis of unit root can be rejected at the 1% level, suggesting that the inflation rate under consideration is suitable for further statistical analysis.

Table 2 displays the autocorrelations of the Tunisian inflation rate until lag 28, ranging from 0.686 to 0.988. It can be seen that the inflation dynamics in Tunisia is characterized by a high level of persistence over time. Our proposed approach in this paper can thus model this long-range dependence feature.

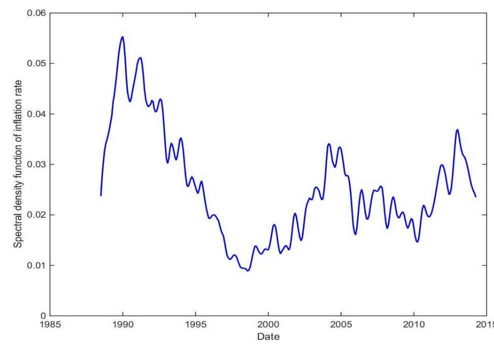
4. Results

Figure 1 depicts the spectral density function of the Tunisian inflation rate over different frequencies ranging from the short-term horizon (4 months and 10 months) to the medium-term horizon (3 years). While the timescale dynamics are not alike across frequencies as the short-term dynamics contain more variability than the medium-term dynamics, and they seem to experience different regimes over time. This evidence leads us to apply the Bai and Perron (1998, 2003) structural change test to these estimated timescale spectral density functions in order to endogenously determine different breakpoints.

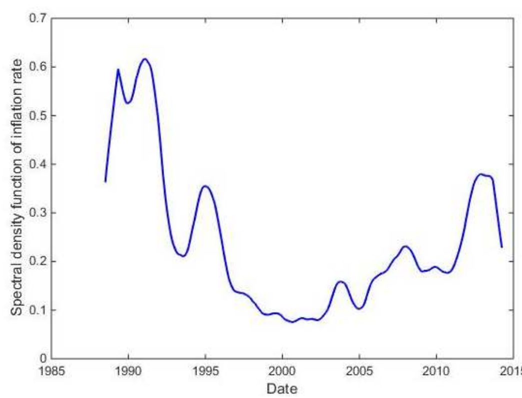
Figure 1: The spectral density function of inflation rate in Tunisia



Short-term horizon ($\frac{10\pi}{20}$, 4 months)



Short-term horizon ($\frac{4\pi}{20}$, 10 months)



Medium-term horizon ($\frac{\pi}{20}$, 3 years)

Table 3 shows a common structural break for all the frequencies under consideration (short-term and medium-term horizons) which occurred around the year 1990. This breakpoint can be explained, on the one hand, by the reforms undertaken by the Tunisian government at the end of the 1980s and, on the other hand, by the adoption of a new monetary targeting (M2) policy. In 1985-1986 the Tunisian economy have experienced a balance payment crisis. As results, Tunisian policy-makers have constraint to adopt stabilization and structural adjustment programs of economic and financial liberalization. The primary goal of these programs is to move for an open and market-oriented system. The main reforms of these programs consist on the reduction progressively of credit allocation controls by abolishing credit ceilings and preferential interest rates. The interest rate becomes freed under central bank control. Further, a capital market has been started and a lot of non-bank financial institutions have been created. Since this date and in line to developed countries, we find a legal financial system and an informal one.

Table 3: Results of the Bai-Perron structural change test applied to spectral density functions (1987-2014)

	Short-run: 4 months	Short-run: 10 months	Medium-run: 3years
Break 1	1990 M10**	1990 M03***	1990 M02**
Break 2	1997 M05*		1997 M03**
Break 3	2012 M02	2011 M01**	2010 M12**

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

Our analysis identifies a second break in 1997 for the short-run frequency (4 months) and the medium-term frequency (3 years). This detection is quite meaningful as the first inflation regime (1987-1996) is characterized by a gradual decrease of inflation rate compared to the increasing tendency of inflation since the end of the 1990. Some studies including Elkaram (1990) show that the monetary policy adopted in late of the 1980s had been disarmed because the Tunisian central bank is unable to act on interest rate owing to its hesitation between stabilisation and expansionary policies. Differently, since the beginning of the 1990s, the Tunisian central bank adopted a prudential regulation policy for banker sector, which thus leads to reduce the liquidity level in the economy and in turn the inflation rate. In 1996, the Tunisian central bank gradually decreased the interest rate and revised its monetary aggregate targeting in several times (1996, 1997, and 1999). These actions seem to explain the structural break identified by our analysis in 1997.

The Bai-Perron test shows another common breakpoint at the end of 2010 and the beginning of 211, which seemed to coincide with the Tunisian revolution (January 2011). Figure 1 indicates that the Tunisian inflation rate had an increasing tendency after the break, which is explained by the economic downturn and political instability generated by the Tunisian revolution. Indeed, the popular uprising since 2011 together with increased insecurity has severely affected many economic sectors (such as phosphate, industrial production...) and led foreign investors to liquidate their investment projects and leave the country. Massive capital outflow caused the Tunisian currency to devaluate, which consequently induced a rise in inflation.

Overall, according our methodology, we identify four inflation regimes in the case of Tunisia.

Before 1990, Tunisian experienced a higher inflation regime with bad macroeconomic indicators. Boughrara and Smida (2004) obtain similar results. The second regime spanned from 1991 to 1996, which is characterized by a decreasing tendency of inflation rate, even though some macroeconomic indicators such as the objective of monetary aggregate are not achieved. The third regime between 1997 and 2010 reveals a more stable inflation rate with a slightly increasing tendency but the inflation level is still lower than that in the first regime. In this period, the main objective of the Tunisian monetary authority was to insure the financial stability of capital markets. The last regime characterized the bad economic situation of Tunisia with a high level of inflation, a high unemployment rate and a slow growth.

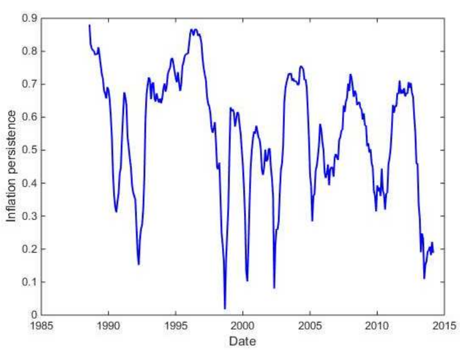
Turning out to the persistence issue of inflation rate over time, it is firstly important to note that the concept of persistence has important implications on the macroeconomic and monetary theories. In the earlier 1970s, many models have established the role of persistence notion (mainly inflation and output persistence) regarding expectations (e.g., Fischer, 1977; Gray, 1976; Taylor, 1980; Calvo, 1983; Rotemberg, 1982, 1983). A deep understanding of inflation persistence is particularly important for monetary policy analysis.[6] It allows one to not only determine the nature of inflation responses on shocks but also to analyze the effect of monetary policy on the inflation. The past literature on inflation persistence distinguishes between the reduced form of inflation persistence and the structural inflation persistence. The reduced form of inflation persistence refers to an empirical property of an observed inflation measure, without economic interpretation, whereas the structural form consists in specifying the economic sources of inflation persistence. Many measures of inflation persistence have been proposed in the literature and they generally involve testing the presence of unit root, the autocorrelation function of inflation series, the sum of the autoregressive coefficients for inflation, and the dominant root of the univariate autoregressive inflation process. Our present study considers the reduced form of inflation persistence and proposes a time-varying measure of inflation persistence by adopting the evolutionary auto-spectral analysis as presented in Section 2.

Figure 2 clearly shows that the Tunisian inflation exhibits a higher degree of persistence over time, regardless of the frequencies (short-term vs. medium-term horizon). Figure 2 shows two

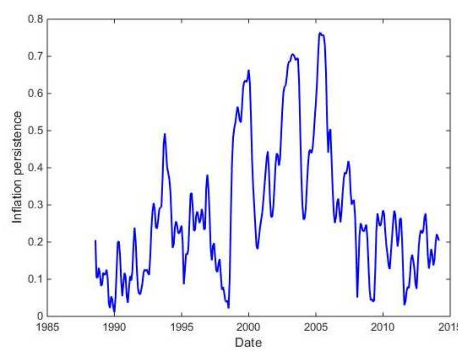
[6] Persistence of inflation assumes that inflation is positively correlated with its own lags, an assumption that holds up well over most of post-war history. More generally, a time series may be deemed persistent if the absolute value of its autocorrelations is high, so that a strongly negatively auto-correlated series would also be characterized as persistent.

different patterns of Tunisian inflation persistence: pre-revolution (pre-2011) and post-revolution. Concerning the first period, while the periods of high persistence are relatively the same across frequencies, the inflation persistence is higher for the short-term horizon than for the medium-term horizon on average. Indeed, the short-term horizons (4- and 10-month frequencies) show an average inflation persistence of 70% and 50%, respectively. As for the medium-term horizon, it experiences an average persistence of 50%. Although the degree of persistence decreases over time, our results show that the medium term persistence is still higher, around 50% on average. This finding thus suggests that the Tunisian inflation only responds gradually to shocks affecting it. This evidence is therefore very important for the conduct of monetary policy because it requires the monetary authority to implement institutional reforms if they want to reduce the inflation rate. Otherwise, monetary policy actions that consist in interest rate cut-offs may not be effective when the inflation has a higher degree of inertia.

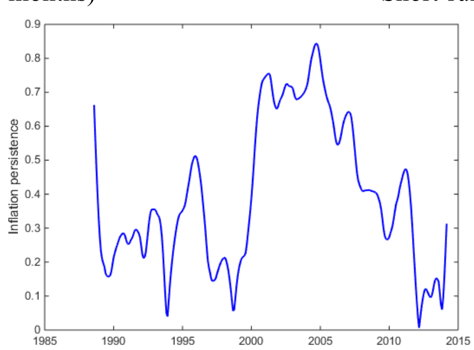
Figure 2: Patterns of the Tunisian inflation persistence



Short-run horizon (2 months)



Short-run horizon (10 months)



Medium-term horizon (3 years)

For the post-revolution period, we observe a dramatic decrease of persistence level for all studied frequencies. Improvements in economic situation and stability following the establishment of a technocratic government has helped to reduce uncertainty and restored consumer and investor confidence, which in turn makes the inflation rate more responsive to the implementation of economic and monetary policies.

5. Conclusion

The great challenge faced by the Tunisian monetary authority is to stabilize the general price level in order to promote growth. This objective can only be reached if we have an in-depth understanding of the inflation dynamics over time and across horizons. In this paper, we investigate the Tunisian inflation rate on two dimensions. We first examine whether the Tunisian inflation rate has a regime-switching behavior and then whether the Tunisian inflation is persistent over time. In particular, we propose a new time-varying measure of inflation persistence which relies on the use of the evolutionary auto-spectral analysis. This approach is interesting in that it allows us to look at the persistence patterns of inflation rate in a time-frequency domain.

We mainly find that the Tunisian inflation rate is subject to a regime-switching behavior in views of three distinct regimes which correspond to monetary reforms and policy actions of the Tunisian central bank. On the other hand, our measure of inflation persistence shows that the Tunisian inflation is highly persistent over time and generally more persistent in the short term than the medium term. This gradual response of inflation rate to economic shocks affecting it implies that policymakers should undertake institutional reforms (e.g., independence of central bank, creation of a disinflationary environment in order to prepare the adoption of the inflation targeting policy, capital account liberalization, banking liberalization, and interest rate deregulation) to reduce inflation rate because a policy consisting in lowering policy interest rate takes time to be effective given the high degree of inflation’s inertia. For instance, following the increasing tendency of inflation rate in Tunisia since 2011, the Tunisian central bank has managed multiple revisions of interest rates but these actions did not bring the expected results.

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