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Abstract

This paper investigates the dilemma of long memory versus a switching regime for the Tunisian real exchange rate (TRER). Empirically, three long memory tests are implemented to examine the long-range dependence in the processes of Tunisian REER. All long memory tests that we used are based on the frequency approach (log-periodogram estimation). While, we apply the Bai Perron (1998, 2003) test in order to detect structural changes in the studied series. In order to discriminate between true long memory or spurious long memory on presence of structural change, we adopt a recent test developed by Perron and Qu (2010). The empirical results show strong evidence in favour of a short memory process with level shifts and not a true long memory process. The presence of structural break traduces that policymakers in Tunisia are unable to react on the exchange rate system through central bank interventions. Therefore, we recommend to Tunisian policymakers to follow a strategy of the market-orientation to make the relevance reforms, such as revisit the weight of dinar with euro, in order to ensure a better managing of the foreign exchange risk and fluctuations.

Keywords: Real Exchange Rate, Long Memory, Structural Breaks, Spurious, and Tunisia.

JEL Classifications: C22, F31.
I. Introduction

After the Arabic spring many economic indicators has been downward in concerning countries. Especially in the case of Tunisia, many macroeconomic indicators have been distorts such as growth rate, inflation, persistent macroeconomic disequilibrium and especially a continuously depreciation of the exchange rate. Indeed, since 2000 decade, the Tunisian exchange rate has decreased by 69% (32%) compared to the euro (dollar). Nevertheless, this depreciation did not favour the Tunisian exportation behind the European recession (2011-2012) and the showdown of the Tunisian economic activity since 14 January 2011 up to now. But, it has enhanced the importation by 22% (2012) and it has created an environment of inflationary pressures.

Understanding the dynamic of exchange rate, the theory of purchasing power parity (PPP) is one of the central building blocks in the literature of international economics allowing this objective.¹ This concept was originally proposed in the 16th century, but since the 1970s, this theory has been a subject hotly debated (Taylor, 2006). The research on PPP is extensive (see Taylor and Taylor, 2004 and Taylor, 2006 for recent surveys on PPP), but the conclusions are mixed, depending on the different countries, data periods or econometric methods. For a long time, these mixed empirical evidence on PPP was attributed mainly to the argument that, owing to the very low adjustment² to PPP, the sample for the recent floating period was too short to detect a statistically significant mean reversion (Froot & Rogoff, 1995; Juselius & MacDonald, 2004). However, some recent papers (Taylor, Peel, & Sarno, 2001; Kilian & Taylor, 2003; Dufrenot et al., 2008; and Aloy et al., 2013) show that nonlinear mean reverting processes can well characterize major real exchange rates (RER).

Different explanations of the nonlinearity behavior of the RER are proposed in the literature. Benning and Protopapadakis (1988), Dumas (1992), Michael, Nobay and Peel (1997), suggested that transaction cost is the main cause of nonlinearity behaviour of exchange rate. However, Krugman (1991) explained this by the existence of the target zone. Teräsvirta (1994), De Grawe et Grimaldi (2001) showed that the major sources of nonlinearities in the exchange rate dynamic consist on the heterogeneity of participants in the foreign exchange market and heterogeneity in investors. Indeed, investors have different objectives arising from different investment horizons, geographical location, and various types of risk profiles and institutional constraints. It is argued that this type of heterogeneity explains why investors often respond differently to the same set of news, and it is shown to generate a nonlinear exchange rate process.

A large volume of literature deals with the nonlinear behaviour of the exchange rate. Most of these studies have shown significant nonlinear character of real exchange rate (Dufrénot et al., 2006; Aloy, et al., 2011; Caporale and Gil-Alana, 2013). However, only a few attempts have analysed the

¹ There is other theories aim to explain the dynamic of exchange such as: balance of payments and financial globalisation.
² Cheung & Lai (1998) and Rogoff (1996) have explained the slow adjustment through the intertemporal smoothing or cross-country wealth distribution.
dynamics and distributional characteristics of the real exchange rate. In other words, they’re a fewer number of studied (Aloy et al., 2011) has been focused on what kind of nonlinearity supporting by real exchange rate: if the exchange rate follow a long memory process or regime-switching model. Indeed, the long memory phenomenon is related to the intersection of several phenomena: self-similarity of distributions, fractional integration or short memory aggregation for some process. However, shift level phenomenon is related to breaks explained by exogenous shock, structural change and policy change. So far, modelling nonlinearity properties of exchange rate is still of major interest in the economic literature as central bank interventions, theory of heterogeneous investors (noise trader vs. arbitrage traders, chartists vs. fundamentalists, and transaction costs). One should note that long memory and structural breaks are at the heart of the debate regarding times series modelling. While persistence deals with exponential decays in the autocorrelation function of a time series, long memory processes requires models accommodating persistence over long horizons. But, a presence of structural breaks may reduce the persistence of time series and hinder the prediction process.

In our acknowledgement, there are no studies that interested on real exchange rate (RER) modelling for the case of Tunisia. We think, that understanding the behaviour of exchange market can offer solutions to Tunisian policy-makers to deal with the continuously showdown of the RER when some instruments are no longer operational such as devaluation, since the dinar is partially convertible since 1992. The objective of this paper is to assess whether the dynamic of Tunisian exchange rate follow a long memory process or a process with switching regime. Indeed, a long memory process implies that Tunisian RER is sensitive to productivity chock. However, if it follows a nonlinear model, through a regime-switching model, therefore his dynamics is explained by the heterogeneity of investors. These distinguish between these two kind of models is useful for policymakers to manage the RER.

In this article, we extend the existing literature on the dynamics of real exchange rate by examining the relevance of structural breaks and long memory in modelling the Tunisian real effective exchange rate (REER). Empirically, three long memory tests are implemented to examine the long-range dependence in the processes of Tunisian REER. All long memory tests that we used are based on the frequency approach (log-periodogram estimation). While, we apply the Bai Perron (1998, 2003) test in order to detect structural changes in the studied series. In order to discriminate between true long memory or spurious long memory on presence of structural change, we adopt a recent test developed by Perron and Qu (2010).

Our results show that Tunisian REER has short long memory process with shifts change. So, the dynamic of the Tunisian real exchange rate is not generated by a fractional model such ARIFMA but threshold models are more suitable. Our research thus constitutes a good venue for understanding the distributional characteristics of Tunisian real exchange rate and has important implications for economic and policy decisions. First, the strong evidence of short memory in presence of structural
breaks we found in Tunisian real exchange rate implies that the linear models are misspecified and cannot be properly used for policy analysis and forecasts. Moreover, we conclude that the productivity shocks (such occurred in Tunisia since 14 January 2011) had a short memory effect on the exchange rate. Finally, the dynamic of Tunisian exchange rate is governed by heterogeneity of investors and policy change. Indeed, the presence of different regimes of Tunisian exchange rate implies that monetary policy did not able to affect its dynamic through central bank interventions.

The remainder of paper is structured as follows. Section 2 briefly reviews related literature. Section 3 provides empirical methodology. Results are summarized and discussed in Section 4. A conclusion follows and points out directions for future research.

I- Literature review

At the beginning of 80s, the debate on studying the exchange rate fluctuation has been grown. Since the seminal paper of Meese and Rogoff (1983), numerous studies have attempted to forecast exchange rates, using both purely statistical models and models based on macroeconomic fundamentals. Nevertheless, the research has not yet reached a consensus as to whether any of these models is able to outperform a drift less random walk consistently when forecasting exchange rates out of sample. The structure of the fundamental-based models is typically represented by a linear function consisting of two parts: a constant term (characterizing the permanent disparity between the two economies in question during the sample period) and a function of macroeconomic fundamentals (characterizing a temporary time-varying imbalance). One route that researchers commonly take is to employ various macroeconomic models to model the fundamental-based part in an attempt to outperform the random walk. However, Meese and Rogoff (1983) proved the difficulty of macroeconomic models in predicting exchange rates. Indeed, they showed that a simple random walk provided regardless of the forecast horizon considered better than all models of exchange time, the PPP. Flood and Rose (1999) proved the difficulty in explaining exchange rate's volatility. So, a common consensus emerges in 90s to specify the exchange rate dynamic through empirical models. However, many studies have proved the inadequacy of the linear models in modelling the exchange rate dynamic and their mean reverting behaviour. Consequently, since the end of 1990s, there is a common consensus to specify exchange rate through nonlinear models. The related literature can be presented through two strands:

The first strand of literature, model the dynamic of exchange rate through a long

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3 In addition, the structure of the fundamental-based models is very similar between studies.
memory process. This kind of model has increased much attention in the recent literature. The long memory character describes the high-order correlation structure of a series. If a series exhibits long-term dependence (long memory), there is persistent temporal dependence even between distant observations. These kinds of series are characterized by distinct but non-periodic cyclical, patterns. A long memory property causes nonlinear dependence in the first moment of the distribution. Therefore, long memory allows potentially predictable component in the series dynamics.

The based theoretical leading to specify the exchange rate dynamic through a long memory process is explained through the productivity shocks. Indeed, persistent misalignments of the exchange rates from their equilibrium value (i.e., the Purchasing Power Parity or an equilibrium value according to the fundamentals) are due to the long-memory property of the adjustment process towards equilibrium. Fractionally integrated processes can give rise to long memory. In this case it would be more natural to study the RER misalignment using fractional cointegration models (see, for example, Dufrénot et al., 2006). This is a relevant point since it is well known that long-memory and nonlinearity can easily be confused. Many studies have been interested to analyze the impact of productivity shocks on the exchange rate dynamic. For example, Baum, Barkoulas, and Caglayan (1999) estimated fractional ARIMA (ARFIMA) models for RER in the post-Bretton Woods era and found almost no evidence to support long run PPP. Additional studies on exchange rate dynamics using fractional integration are those by Crato and Ray (2000), Wang (2004), Dufrenot, et al. (2008) and Aloy, Booutahar et al., (2011) among others. Some recent studies have specify the dynamic if exchange rate through long memory process but for high frequency data such as (Caporale and Gil-Alana, 2013).

Additionally, the long-memory property may be relevant in the RER debate (PPP), since temporal aggregation (Taylor, 2001) or cross-sectional aggregation (Imbs et al., 2005) are found to induce a positive bias in the computed aggregate half-lives. The empirical literature (such as, for instance Diebold and Inoue, 2001) has shown that aggregation may be a source of long-term dependence. Fractional integrated processes exhibit a hyperbolic reversion towards the long-run equilibrium after a shock.

The second strand of literature suggested that the nonlinearity in the dynamic of the exchange rate is attributed mainly to investors heterogeneous (noise trader vs. arbitrage traders, chartists vs. fundamentalists, and transaction costs). Indeed, Taylor et al. (2001) considered that nonlinear RER dynamics could be formally derived in the context of international arbitrage costs. Alternative sources of nonlinearity consist on the interaction of

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4 An extensive mathematical research over the last years has focused on important issues such as stock exchange and foreign currency, in economics and finance.
heterogeneous agents (Kilian and Taylor, 2003, Taylor, 2006) and the influence of official intervention in the foreign exchange market (Menkhoff and Taylor, 2007; Reitz and Taylor, 2008). From the PPP theory, bands of transactions and iceberg costs are related to regime switching models. This kind of models (regime switching models) has been successfully applied to exchange rates by various researchers. Many empirical studies have been proved the successful of Markov switching models on studying the exchange rate dynamics (Evans and Lewis, 1995; Dewachter, 2001; Chen and Lee, 2006...). Engel (1994) showed that a Markov switching model fits well in-sample for many exchange rates, but is not able to generate forecasts which are superior to the random walk according to either the MSPE or mean absolute error (MAE) criteria. Bergman and Hanson (2005) characterise the exchange rate dynamic through a 2 state Markov Switching AR (1). They show that the RER switches between two states and exhibits mean reversion within each regime where the drift is negative (positive) when the deviation from parity is positive (negative). So, the authors reject a single regime autoregressive model for the RER and that only the drift is state dependent. Nikolsko-Rzhevskyy and Pordan (2012) use two-state Markov switching stochastic segmented trend model and they show evidence of both short-run (one month) and long run (up to one year) predictability for monthly exchange rates over the post-Bretton Woods period. Authors proved evidence of consistent multi-horizon predictability. The model strongly outperforms the random walk for 9 out of 12 exchange rate series at short horizons; for 7 of the 12 exchange rates.

Other kind of regime switching models (Threshold autoregressive models-TAR models-) is used in literature to specify the dynamic of Exchange rate. Taylor et al. (2001) employ an ESTAR models to four-major RER against the dollar. They find that the studied exchange rates are well characterized by nonlinear mean reverting processes. Kilian and Taylor (2003) confirm the nonlinear behaviour of exchange rate through ESTAR model. Wu and Hi (2009) extend the analysis of Kilian and Taylor (2003) and include the Harrod/Balassa/Samuelson effect by allowing for deviation from purchasing power parity return to an equilibrium trend. They provide further evidence of nonlinear exchange rate. In this context, Nam (2011) develops a recent analysis via a discrete threshold models. He shows that the nominal exchange rate adjusts only outside a threshold band. Through these results, The PPP is restored mainly through adjustments in relative prices within the band.

Previous studies, cited above, employ different methodologies to modelling the dynamic of the exchange rate dynamics: long memory process and regime switching models. However, an important criticism has been addressed to the first kind of model (long memory process).

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5 Under the assumption of spatially separated countries and shipping cost,
6 To discriminate between different models, the authors use the out-of-sample forecasting behaviour and they employ the mean squared error to distinguish between out-of-sample forecasting forecasts.
Indeed, the long memory property in the data may be due to the presence of structural breaks or regime switches, which is called “the spurious long memory process”. Several recent works including Granger (1999), Diebold and Inoue (2001), Granger and Hyung (2004), Smith, 2005, and Aloy et al. (2013) show that the presence of structural breaks or regime switching can generate spurious long memory behavior in the RER series. In other words, it is very useful to distinguish between switch regimes and the long memory process to avoid any wrong political implications. Therefore, the aim of this paper is to propose a new methodology allowing distinguishing between regime switching and spurious long memory in the dynamic of the exchange rate.

II- Empirical methodology

In this part of analysis, we present our empirical strategy to deal with structural break or long memory for the Tunisian reel effective exchange rate (REER). Therefore, we present the main test that we used to detect long memory behaviour in times series analysis. Then, we present the followed test that we used to detect structural change. At the end, we present the test to discriminate between long memory and structural break.

1- Procedures to detect long memory behavior

The empirical literature has proposed, in the last two decades, varieties of approaches (semi-parametric and nonparametric) for the estimation of the fractional long memory parameter $d$. However, semi-parametric estimators appear to be particularly more attractive than others because they are agnostic about the short-run dynamics of the process and hence robust to misspecification of these dynamics. Nevertheless, these kinds of estimators are still being developed. Therefore, there are few semi-parametric estimators that can be used to satisfactorily assess the non-stationary case. In our analysis, we retain: the Geweke and Porter-Hudak (1983) method and its extension proposed Andrews-Guggenberger (2000), the Robinson’ test of fractional integration (1995), and the Shimotsu (2006) technique extension for the Exact Local Whittle (ELW) of Shimotsu & Phillips (2005). We start by a brief description of the long memory process. Then, we present the used methods of long memory parameter estimation.

1.1 Definition of long memory process

The empirical literature proposed several ways in order to defined the long memory process. From a time domain point of view, the long memory is defined as a decay rate of long lag-autocorrelations. However, from a frequency domain, it is defined as rates of explosions of low frequencies spectra.
Concerning the time domain, let $X_t$ a covariance stationary process. This series exhibits a long memory process if its autocorrelations $\rho(k)$ has a slow decay and persistence such that

$$\sum_{k=-n}^{n} |\rho(k)| \rightarrow \infty \text{ as } n \rightarrow \infty$$

In order to respect the stationary and invertible conditions, the autocorrelation function of $X_t$ must satisfy the hyperbolic decay

$$\rho(k) \sim ck^{2d-1} \text{ as } n \rightarrow \infty$$

where $c \neq 0$ and the long memory parameter $d \in \left[0, \frac{1}{2}\right]$. If $d < 0$, $\{X_t\}$ has “intermediate” or “anti-persistent” memory since $\sum_{k=-\infty}^{\infty} |\rho(k)| < \infty$.

Concerning a frequency domain, the long memory is defined when we evaluate the spectral density function at frequencies that tend to zero. In the case of spectral density function $f(w)$ respecting the following property $f(w) \sim c|w|^{-2d}$ as $w \rightarrow 0^+$, the process $\{X_t\}$ exhibits long memory where the stationary range of $d$ is similar to the time domain definition above.

The first long memory process introduced in the literature is the popular Auto-Regressive Fractionally Integrated Moving Average (ARFIMA) model developed independently by Granger and Joyeux (1980) and Hosking (1981). This model satisfies the above conditions (cited in the definition above). We say that a process $\{X_t\}_{t=1}^{T}$, with $t \in \mathbb{Z}$ follows an ARFIMA $(p; d; q)$ process if it takes the form,

$$(1 - L)^d \phi(L)X_t = \theta(L) \epsilon_t \quad (1)$$

Where, $L$ is the lag operator i.e. $LX_t = X_{t-1}$; $\epsilon_t$ is a Gaussian strong white noise $N(0, \sigma^2)$. $\phi(L)$ and $\theta(L)$ are the autoregressive and the moving average polynomials of order $p$ and $q$ respectively, whose roots lie outside the unit circle.

If $d \in (-\frac{1}{2}, 0)$ the ARFIMA$(p; d; q)$, the model (1) is an invertible stationary process with intermediate memory. However, if $d \in (0, \frac{1}{2})$, the model (1) is stationary and invertible and has an autocorrelation function $\rho(k)$ which exhibits a slow decay when the lag $k$ increases, see Beran (1994). In this latter case we say that we are in presence of a stationary long memory behavior.

The testing procedure for the presence of a long memory behavior consists on testing the null hypothesis:
\begin{align*}
H_0 : d &= 0 \\
H_1 : d &
eq 0
\end{align*}
where \( d \) is the long memory parameter introduced in (1). Thus, under the null hypothesis \( (H_0) \) we have a short memory behavior and a long-range dependence under the alternative \( (H_1) \).

1.2 The Geweke and Porter-Hudak (1983) method (GPH)

Geweke and Porter-Hudak (1983) proposed a first semi-parametric method based on spectral analysis, to estimate the long memory parameter \( d \). This technique becomes extensively used behind its computational simplicity. It is based on log-periodogram regression estimator.

Let \( y_t \) be the Tunisian REER. The GPH estimator for the long memory parameter \( d \) for \( y_t \) can be determined a log-periodogram defined as follows:

\[
\ln\{I(w_j)\} = \alpha + \beta \ln\left\{ 4\sin^2\left(\frac{w_j}{2}\right) \right\} + \varepsilon_t \quad j = 1, \ldots, m.
\]

Where, \( w_j = \frac{2\pi j}{T} \); \( \varepsilon_t \): is the residual term ; \( w_j \) represents the Fourier frequencies \( m \) \((m \in N)\). For consistency, it is required that \( m \) grows slowly with respect to the sample size. It is suggested to set \( m = \sqrt{T} \) (see Banerjee and Urga, 2004). \( I(w_j) \) is the periodogram that defined as follows:

\[
I(w_j) = \frac{1}{2\pi T} \left| \sum_{t=1}^{T} y_t e^{-w_j t} \right|^2
\]

Where \( y_t \) is assumed to be a covariance stationary times series. The estimated \( d \) of parameter integration says \( \hat{d}_{GPH} \) is \((-\hat{\beta})\). The ordinate least-square estimator of \( d \) is asymptotically normal with standard error equal to \( \pi(6m)^{-1/2} \) (Geweke and Porter-Hudak, 1983; Robinson, 1995). Agiakloglou et al. (1992) show that this method is biased and inefficient when the error term is an \( AR(1) \) or an \( MA(1) \) and in addition this estimator does not possess good asymptotic properties. Note that this method is only robust for \(|d| < 1/2\).

1.3 The Robinson’s test of fractional integration (1995)

Robinson (1995) proposes a frequency domain test based on spectral analysis in order to resolve the above limit of GPH test. Indeed, he developed a proof of consistency and asymptotic normality for \( 0 < d < 1/2 \).

Consider the following process

\[ y_t = \mu_t + \varepsilon_t \quad t = 1, \ldots, T \]

Where \( \{\varepsilon_t\} \) has a spectral density given by

\[ f_\varepsilon(v) = Gv^{-2d} \quad \text{as} \quad v \to 0 \]

For \( 0 \leq d < \frac{1}{2} \) and \( G \in (0, \infty) \) As the process \( X_t \) has \( n \) mean changes (level shifts)
\[
\mu_t = \mu_1 + \sum_{i=1}^{n} \lambda_i 1\{k_i < t \leq k_{i+1}\} = \mu_1 + \sum_{i=1}^{n} \lambda_i 1\{[T \tau_i] < t \leq [T \tau_{i+1}]\}
\]

Where, \(1\) is the indicator function, \(n\) is the number of breaks, \(k_i, i = 1, \ldots, n\), are the true change points, and \(\lambda_i = \mu_{i+1} - \mu_i k_i\) denote the magnitudes of changes. Also let \(\tau_i = \frac{k_i}{T}, \tau_0 = 0\), and \(\tau_{n+1} = 1\).

In order to estimate the fractional long memory parameter \(d\), Robinson (1995) proposes a local Whittle estimator, \(d_r\). This estimator is based only on the spectral density in the neighborhood of frequency zero and ignores a correct specification of short-run dependence. When there is no mean change, \(\mu_t = \mu_1\), Robinson (1995) proves that:

\[
\frac{1}{m^2} \left( d_r - d_0 \right) \Rightarrow N(0, 1/4)
\]

Where the bandwidth \(m\) is an integer less than \(\left\lfloor \frac{T}{2} \right\rfloor\) to avoid aliasing effects; \(d_0\) is the true value of \(d\), with the only additional requirement that \(m \rightarrow \infty\) slower than \(T\); \(\frac{1}{m} + \frac{m}{T} \rightarrow 0\) as \(T \rightarrow \infty\). The bandwidth can be selected by minimizing mean squared error for the local Whittle or Geweke Porter–Hudak estimators (Henry, 2001). We suggest using the local Whittle method and incorporating potential breaks into the models for long-memory tests.

### 1.4 Andrews and Guggenberger (2003) test (AG test)

Andrews and Guggenberger (2003) proposed other procedure in order to resolve the limit of GPH technique. They develop a bias-reduced log-periodogram estimator \(\hat{d}_r\), eliminating the first- and higher-order biases of GPH estimator. This estimator is similar to the GPH estimator except that one includes frequencies to the power \(2k\) for \(k = 1 \ldots r\), for some positive integer \(r\), as additional regressors in the pseudo-regression model that yields the GPH estimator. The reduction in bias is obtained using assumptions on the spectrum only in a neighbourhood of the zero frequency. Andrews and Guggenberger (2003) follow the study of Robinson (1995b) and Hurvich, Deo, and Brodsky (1998) in order to establish the asymptotic bias, variance, and mean-squared error (MSE) of \(d_r\), determine the asymptotic MSE optimal choice of the number of frequencies, \(m\), to include in the regression, and establish the asymptotic normality of \(d_r\). This approach shows that the bias of \(d_r\) goes to zero at a faster rate than the GPH estimator.

### 1.5 The Shimots (2006) test

Shimotsu (2006) has developed a new semi-parametric method. There is an extend version of the Exact Local Whittle (ELW) estimator proposed by Shimotsu and Phillips (2004, 2005). This later is a semi-parametric method giving a good estimation for the fractional memory parameter in terms of consistency and limit distribution, except the case of unknown mean. Therefore, Shimotsu (2006)
extend the ELW test to Feasible Exact Local Whittle (FELW) and showed that this estimator is consistent and has $N(0, 1/4)$ limit distribution for $d \in (-\frac{1}{2}, 2)$.

Shimotsu (2006) considers a class of fractional process $y_t$ generated by the following model

$$(1 - L)^d y_t = \mu_t \cdot I\{t \geq 1\}, \quad t = 0, \pm 1, \pm 2, ...$$

(2)

Where $\mu_t$ is an I(0) process with mean zero and spectral density $f_\mu$ satisfying $f_\mu(\lambda) \sim G$ for $\lambda \sim 0$, and $I\{.\}$ is an indicator function.

The estimated value $\hat{d}_{ELW}$ is obtained as follows:

$$\hat{d}_{ELW} = \arg\min_{d \in [d_1, d_2]} R(d)$$

Where $d_1$ and $d_2$ are the lower and upper bounds of the admissible values of $d$ such that: $-\infty < d_1 < d_2 < 1$ and $R(d) = \ln G(d) - 2d \frac{1}{m} \sum_{j=1}^{m} \ln(w_j)$.

Where $m$ is the truncation parameter, and $G(d) = \frac{1}{m} \sum_{j=1}^{m} I_{\Delta(d,y)}(w_j)$

$I_{\Delta(d,y)}(w) = \frac{1}{2\pi T} \sum_{t=1}^{T} \Delta^d y_t e^{i\lambda w} \delta_{\Delta(d,y)}(\lambda)$

the periodogram of $\Delta^d y_t = (1 - L)^d y_t$. Under certain assumptions, the ELW estimator $\hat{d}_{ELW}$ satisfies:

$$\sqrt{m}(\hat{d}_{ELW} - d) \rightarrow \mathcal{N}(0, 1/4).$$

Where $m$ is chosen so that it satisfies as $T \rightarrow \infty$

$$\frac{1}{m} + \frac{m^{1+2\beta}(\log m)^2}{n} \rightarrow 0, \quad \forall \gamma > 0$$

$\beta$ represents the degree of approximation of the spectral density of $u_t, f_u(\lambda)$, around the origin by $G$. Shimotsu (2006) further developed these results so that it can accommodate an unknown mean and a linear time trend. Hereafter, following Shimotsu (2006), we call that estimator the FELW estimator. He suggested estimating unknown mean by:

$$\tilde{u}(d) = \bar{X} \cdot I(d < c) + X_t \cdot I(d \geq c); \quad c \in \left(\frac{1}{2}, \frac{5}{8}\right)$$

Where $\bar{X}$ is the sample average and applied the ELW estimator to the variable $X_t - \tilde{u}(d)$. Shimotsu (2006) proved that the FELW estimator has the same asymptotic distribution as the ELW estimator, under assumption for $d \in (-\frac{1}{2}, 2)$ excluding arbitrary small intervals around 0 and 1. This result gives a basis for constructing valid asymptotic confidence intervals for $d$, which can be supplemented with finite sample confidence intervals based on simulation.

2- Procedures to detect structural break

2.1 Brief definition of structural break

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7 For consistency and asymptotic normality see assumptions (1-5) and (10-50) in Shimotsu and Phillips (2005), page X
Economic structural change refers to a long-term shift in the fundamental structure of an economy, which is often linked to permanent changes in resources, growth, economic development and policy makers’ decisions… A structural change will shift the parameters of an entity, which can be represented by significant changes in time series data. Hansen (2001) proposed a conventional definition of the term structural change: “The term “structural change” is far from having a univocal meaning in the field of economics. In econometric theory, the issue of structural change refers to the behavior of the parameters of a model in the course of time. The usual assumption of stationarity is commonly made, according to which one or all of the relevant parameters of the econometric model are constant over time. However, structural breaks can occur, such that one of these parameters changes at some time in the sampled period. The econometrics of structural change allows to identifying structural breaks in time series by providing a rich set of tests”.

Many studies related to structural changes have paralleled developments in the analysis of unit root models. Indeed, the structural change are easily confused with unit root (see Perron, 1989; Zivot and Andrews, 1992). This confusion is explained by the using of similar tools, such as: functional central limit theorems or invariance principles. In last decade a great attention has been attributed on the confusion between long memory and structural change. Granger and Hyung (2004), and Choi and Zivot (2007) suggested that the observed long memory property in time series could be explained by the presence of structural breaks. To investigate this conjecture for the Tunisian REER, we use the pure multiple mean break method proposed by Bai and Perron (1998, 2003).

2.2 Bai Perron (1998, 2003) test

Empirical literature has proposed many test for structural change such as the seminal work of Chow (1960), the CUSUM test of Brown and al. (1994) focusing on testing structural change at a single specified known break date. Recently, there was a development of methods that allow to estimate and to test for structural change at unknown break dates. The most popular one is the test of Bai and Perron (1998-2003). This test has some advantages such allowing the detection of multiple break points in a time series and their occurrence endogenously. We use GAUSS software and we obtain the estimate by running the code created by Bai and Perron (1998, 2003b).8

We consider the following multiple linear regression with l breaks ((l + 1) regimes):

\[ Y_t = x_t' \beta + z_t' \delta_1 + u_t \quad \text{if} \quad 1 \leq t \leq T_1 \]
\[ Y_t = x_t' \beta + z_t' \delta_2 + u_t \quad \text{if} \quad T_1 \leq t \leq T_2 \]
\[ \vdots \]
\[ Y_t = x_t' \beta + z_t' \delta_{l+1} + u_t \quad \text{if} \quad T_l \leq t \leq T \]

8 The code is available on the Perron home page: http://people.bu.edu/perron/.
In this model, $Y_t$ is the observed dependent variable, $x_t(p \times 1)$ and $z_t(q \times 1)$ are vectors of covariates, and $\beta$ and $\delta_j$ ($j = 1, \ldots, l + 1$) are the corresponding vectors of coefficients, $u_t$ is the disturbance. The break points are explicitly treated as unknown. When $\beta$ is not subject to shifts, the model is called a partial structural change model and by imposing $p = 0$ we obtain the pure structural change model in which all the coefficients vary with the break points.

Note that in this structural change model, all the coefficients are subject to change over time. The hypothesis that the regression coefficients remain constant is: $H_0: \beta_i = \beta_0$ for $i = 1, \ldots, K, \ldots, n$, against the alternative that at least one coefficient varies over time. The break points $(T_1, \ldots, T_m)$ are explicitly treated as unknown and for $(i = 1, \ldots, m)$, we have $\lambda_i = \frac{T_i}{T}$ with $0 < \lambda_1 < \cdots < \lambda_m < 1$.

To test for the presence of structural changes we need to estimate the unknown coefficients $(\beta^0, \delta^0_1, \ldots, \delta^0_{l+1}, T^0_1, \ldots, T^0_l)$. These parameters are obtained by minimizing the sum of squared residuals from (3). Then, to determine if structural changes occur, Bai and Perron (1998) suggest to use the following two statistics: $UD_{\text{max}} = \max_{1 \leq s \leq L} \text{statistic} \leq LsupF(l) \text{statistic}$

Where $L$ denote the maximum number of breaks allowed and the $WD_{\text{max}} = \max_{1 \leq l \leq L} \text{statistic} \leq Lw_{l}supF_{l}(l)$, where the weights are such that the marginal p-values are equal across values of $l$. Moreover, the authors propose a Sup F type test of no structural change ($l = 0$) against the alternative hypothesis of ($l = i$) with $i = 1, \ldots, l$ breaks.

To determine exactly the number of breaks, Bai and Perron (1998, 2003) propose a sequential procedure which is based on the three tests presented above and a $\text{Sup} F(l + 1 / l)$ test for the null hypothesis of $l$ states against the alternative of $l + 1$ states, for more details see Bai and Perron (1998, 2003).

3- Perron and Qu (2010) Test

According to Perron (1990), a confusion between levels shifts and the long-term dependence can occurs. Indeed, he showed that a unit root could be confused with a structural change that leading to spurious long memory.

A growing strand of literature has tried to address the issue by developing tests that distinguish between true and spurious long memory (Berkes et al., 2006, Ohanissian et al. 2008, Perron and Qu, 2004; Perron and Qu, 2010). Shimotsu (2006) examined the null hypothesis of long memory against the alternative of a structural change. Two and four subsamples are considered since augmenting the number of hypothetical subsamples does not increase the power of the test. A recent study of Perron and Qu (2010) has been developed to give further insight on this same issue.

In our work we adopt the Perron and Qu (2010) test for some reasons; There is a frequency test based on log periodogram estimator proposed by GPH which is in harmony of our above tests used to detect long memory (all are based on log periodogram). In their proposed test, Perron and Qu (2010)
show how the distribution of this estimator is highly dependent on the number of frequencies used, particularly when the data generating process is a stationary short memory process contaminated by shifts changes in level. This test aims to distinguish structural change from long memory.

Let $d_a$ (respectively $d_b$) denote the log periodogram estimate of the memory parameter when $m_a = \lfloor T^a \rfloor$ (respectively $m_b = \lfloor T^b \rfloor$) frequencies are included in the regression. Under the null hypothesis of a stationary Gaussian fractionally integrated process, if $0 < a < b < 1$ and $a < 4/5$, Perron and Qu (2010) demonstrate that the test statistic follows a Gaussian process under the null:

$$\sqrt{\frac{24[T^a]}{\pi^2}} \frac{d}{(d_a - d_b)} \rightarrow N(0,1)$$

They fixed this statistic to $a = 1/2$ and $b = 4/5$ in order to distinguish whether the process is a true long memory and not a short memory process with level shifts. Perron and Qu (2010) prove that their test is not sensitive to the value of $d$ even if $d > 1/2$. In addition, they show that it is consistent against a short memory process with level shifts or a long-memory process with a strongly mean reverting component.

III- Empirical results

1- Data and preliminary tests

For empirical analysis, we consider the Tunisian Real Effective Exchange Rate. This series is collected from the International Monetary Fund on their database: International Financial Statistics (IFS) during the period 1975:M1 to 2012:M11.

![Fig.1 The Tunisian Real Effective Exchange Rate from 1987 to 2012.](image)

As defined in Eq. (1), the Tunisian Real Exchange rate measures the deviation from PPP: under long-run PPP, the logarithm of the REER must display reversion towards zero (after appropriate scaling).

The period of the study is 1975: M1 to 2012: M11. The choice of this period is justified by many reasons. Firstly, this period covers different varieties of policy exchange rate adopted by Tunisia such as fixed exchange rate policy and floating exchange regime. Secondly, this period covers the Tunisian
policy maker’s decision about the devaluation of exchange rate in 1986. Thirdly, the choice of our data span covers many crises.

Table 1: Summary statistics, unit root and ARCH tests of the Tunisian Real Exchange Rate.

<table>
<thead>
<tr>
<th>Statistics</th>
<th>Value</th>
<th>Test</th>
<th>Statistics</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>4.881787</td>
<td>ADF</td>
<td>-2.848 [S]</td>
</tr>
<tr>
<td>Median</td>
<td>0.0740</td>
<td>PP</td>
<td>-2.656 [S]</td>
</tr>
<tr>
<td>Standard errors</td>
<td>0.284880</td>
<td>KPSS</td>
<td>0.333 [NS] (0.09)***</td>
</tr>
<tr>
<td>Skwness</td>
<td>0.548044</td>
<td>Modified PP</td>
<td>-1.321 (-2.98)**</td>
</tr>
<tr>
<td>Kurtosis</td>
<td>1.78162</td>
<td>DF-GLS</td>
<td>-1.325 [NS] (-2.59)*</td>
</tr>
<tr>
<td>Jarque-Bera</td>
<td>50.92021</td>
<td>ERS</td>
<td>2.384</td>
</tr>
<tr>
<td>P-value</td>
<td>0.0089</td>
<td></td>
<td>(21.59)***</td>
</tr>
</tbody>
</table>

Notes: (.) indicates the calculated statistic of series in difference, [S] indicate a stationary series and [NS] indicates a non-stationary series. * Indicates rejection of the null hypothesis at the 10% level. ** Indicates rejection of the null hypothesis at the 5% level. ***Indicates rejection of the null hypothesis at the 1% level.

Table 1 reports evidence for excess kurtosis (leptokurtic distribution) and asymmetric properties. Clearly, Tunisian REER have fatter tails and longer right tails than the normal distribution. The Jarque–Bera test (JB) confirms these findings since it rejects normality at 1% confidence level. Before conducting further modeling, it is crucial to determine whether the Tunisian REER is stationary or not. In our analysis, the ADF and PP tests support that the Tunisian REER is \( I(0) \). However, these tests are overly likely to reject \( I(1) \) when it is true. To deal with this limit, we adopt four more robust unit root tests: The KPSS test of Kwiatkowski–Phillips–Schmidt–Shin (1992), the ERS test of Elliot et al (1996), the modified PP test of Perron and Ng (1996) and the test of Ng and Perron (2001) using the GLS detrending technique in order to create efficient version of the modified PP test. All unit root tests reject the stationary hypothesis. Therefore, the Tunisian REER is a non-stationary process. Moreover, it is \( I(1) \).

Fig.2 Autocorrelation and Partial autocorrelation functions for Tunisian REER.

Fig.2 displays the distributional characteristics of the Tunisian REER autocorrelation function. This
figure shows that the autocorrelation function of the REER is higher and has no particular form. This is suggestive of their short memory property. Overall, our findings shed light on a very persistent behaviour in REER. It is consistent with the common characteristics of exchange rate widely supported in previous analysis, such as: Maican and Sweeney (2013), Beckmann (2013), Aloy et al. (2011), Choi et al. (2010), Dufrénot et al. (2008), Soofi and al. (2006)…. In addition, it is well argued in the previous literature that these characteristics are suggestive of long memory dynamics, and that they can be spuriously generated when structural breaks are ignored in economic modelling of time series analysis. For example, Diebold and Inoue (2001) emphasize that infrequent stochastic breaks can create strong persistence in the autocorrelation structure of time series.

2- Long memory results

In our analysis we apply four long memory tests (presented above), the GPH test, Robinson (1995) test, Andrews and Guggenberger (2003) test, and Shimotsu (2006) test. One problem in calculating the statistical value of different long memory tests concerns the choice of bandwidth $m$ in finite samples. Since there are no satisfactory analytical results for deciding on the appropriate value of $m$ in finite sample, we choose $m$ based on previous literature and according an own analysis based on simulation. Robinson (1995) suggest that $m = T^{0.5}$ in order to test the fractional long memory parameter. For GPH, AG, FELW tests, the choice of $m$ is different in previous literature. In most previous works $m$ is choose usually as equal for $m = T^{0.5}, T^{0.6}, T^{0.7}, T^{0.8}$ (Simotsu, 2006; Charfeddine and Guégan, 2011; Aloy et al, 2011;…). To deal with this problem, we simulate $y_t = (1 - L)^{-d} \epsilon_t$ for $\epsilon_t$ Gaussian white noise with sample size of (25 years) 455 and compute the sample mean squared error (MSE) for the several choice of $m$, that is $m = T^{0.5}, T^{0.6}, T^{0.7}T^{0.8}$. The simulation results indicate that the sample MSE is minimized when $m = 0.5$ or most of the cases. Therefore, in our analysis, we fixed the optimal value of $m$ to 0.6 as the bandwidth. However, for robustness analysis, we report the result for other different values of $m$.

The obtained results are presented in table 2. We postulate the null hypothesis $H_0: d = 0$ against the alternative hypothesis $H_1: d \neq 0$. Indeed, we conclude for long memory behaviour if $0 < \hat{d} < \frac{1}{2}$, otherwise we reject the long memory behaviour.

Unanimously, the used tests show evidence of long memory patterns for Tunisian REER as the null hypothesis of no persistence is always rejected at levels ranging from 1% to 10% ($0 < \hat{d} < \frac{1}{2}$). For different value of $m$, the long memory behaviour is not rejecting. However, we note that, for some value of $m$ (such $m = 0.5$) the parameter $d$ is higher than 0.5, in the case of GPH technique and non significant and near from zero in the case of AG test. These unusual values of $d$ can be explained through several points. Firstly, they can arise from the bias inherent in the GPH and AG estimators. Secondly, given by Granger and Hyung (2004) and Shimotsu (2006) these values of $d$ can be the result of spurious long memory process. In other word, it may be the result of various kinds of
misspecifications and/or the presence of structural breaks. In this scheme of things, a greater accumulation of misspecifications naturally would lead to greater spurious long memory. According to Lamoureux and Lastrapes (1990), the observed long-memory features of Tunisian REER can arise from external shocks that create a structural shift in the this process.

Table 2: Results of long memory tests for the Tunisian REER.

<table>
<thead>
<tr>
<th>Test</th>
<th>$T^m (m)$</th>
<th>0.5</th>
<th>0.6</th>
<th>0.7</th>
<th>0.8</th>
</tr>
</thead>
<tbody>
<tr>
<td>GPH</td>
<td>0.55</td>
<td>0.29</td>
<td>0.21</td>
<td>0.10</td>
<td></td>
</tr>
<tr>
<td>t-static</td>
<td>2.80***</td>
<td>2.07**</td>
<td>2.35**</td>
<td>1.76*</td>
<td></td>
</tr>
<tr>
<td>Robinson</td>
<td>0.44</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td></td>
</tr>
<tr>
<td>t-static</td>
<td>4.03***</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td></td>
</tr>
<tr>
<td>AG</td>
<td>0.02</td>
<td>0.478</td>
<td>0.38</td>
<td>0.20</td>
<td></td>
</tr>
<tr>
<td>t-static</td>
<td>0.06</td>
<td>2.43**</td>
<td>2.219**</td>
<td>2.22**</td>
<td></td>
</tr>
<tr>
<td>Shimotsu</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>t-static</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: * Indicates rejection of the null hypothesis at the 10% level. ** Indicates rejection of the null hypothesis at the 5% level. *** Indicates rejection of the null hypothesis at the 1% level.

3- Structural break results

The results from the Bai and Perron (1997, 2003) test regarding the number and estimated break dates are reported in Table 3. The result shows that the Tunisian RER exhibits four regimes. The first break is identified at 1980:M12, the second one at 1986:M6, the third one is observed at 1992:M12, and the last one at 2004:M9. These breaks are associated with some crisis occurring in Tunisia and also in the world.

The break of 1980:M12 is explained by political destabilization in Tunisia starting in 26 January 1980. At this date, the Forces Progressive National Front party -formed in Tripoli and founded by former supporters of Salah Ben Youssef apart Tunisian politics- occupied the city of Gafsa.\(^9\) In addition to this political crisis, generated serious economic troubles, a popular uprising takes place in December 1983 known as “bread revolt”. This uprising is a result of the increase of cereal price by 87.5%. During this period, Tunisian economy is characterized by a great recession, great inflation, and problems in balance of payment. In order to promoting competiveness and exportations, Tunisian authorities decided, in the beginning of 1985, to broad their currencies baskets. Thus, the exchange rate of the dinar was henceforth also binds to the currencies of competing countries. In addition, at the end of 1985, they decided to change the weights of currencies. However, these corrections were

\(^9\) An armed commando is formed by dozens of men who were conscripted in Libya. This act leads France, USA and Morocco to send some the warships and some plane in order to repress these army opponents. As consequence, the Tunisian President “Bourguiba” breaks up the diplomatic relationship with Libya.
ineffective (crisis in the balance of payments) forcing the authorities to devalue the currency\textsuperscript{10} by 10\% -in 1986 (Hanna, 2001). This devaluation is accompanied by a new monetary policy based in floating exchange rate regime policy adopted in 1986. Theses actions explain our second break in Tunisian REER identified 1986:M6. The third break point in Tunisian exchange rate is identified in 1992:M12. This break point is explained by the lifting of some restrictions on payments for current transactions (announcement of the current convertibility of the Tunisian dinar in December 1992) and the creation of an interbank foreign exchange market in 1994. Since 2001, Tunisia has broadened the fluctuation range of the nominal exchange rate. This policy was be adopted following the IMF recommendations which aim to made the monetary policy (based on floating exchange rate policy) more flexible (and Faniza al, 2003) in order to improve the competitiveness. This leads to our identified last structural breaks in the Tunisian REER occurring in 2004:M9.

The discrimination between long memory and structural breaks is, however, not an easy task. Several studies have examined the nature and causes of times series persistence for economic and financial series, but the results remain inconclusive. For instance, Bhardwaj and Swanson (2006) find that the long memory models give better out-of-sample forecasts than ARMA, standard GARCH and related models. The long memory models are also found to outperform models with occasional breaks in out-of-sample analysis (Granger & Hyung, 2004).

Table 3: Results of Bai and Perron (1998, 2003) test for the Tunisian REER

<table>
<thead>
<tr>
<th>Break 1</th>
<th>Break 2</th>
<th>Break 3</th>
<th>Break 4</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>5.31</td>
<td>5.25</td>
<td>4.80</td>
</tr>
<tr>
<td>Sup(0/1)</td>
<td>(0/1)</td>
<td>(0/2)</td>
<td>(0/3)</td>
</tr>
<tr>
<td>Statistic of Sup</td>
<td>41.4702***</td>
<td>40.6723**</td>
<td>325.8518***</td>
</tr>
<tr>
<td>Sup F(1+1</td>
<td>l)</td>
<td>(1/0)</td>
<td>(2/1)</td>
</tr>
<tr>
<td></td>
<td>41.4702**</td>
<td>65.3263**</td>
<td>22.400**</td>
</tr>
<tr>
<td>WDmax</td>
<td>688.2743***</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

\textsuperscript{10}The devaluation of the dinar reached to 45\% in 1988 in nominal level and 25\% in real level.
4- Long memory versus structural break results

It is essential to test for the relevance of long memory against structural breaks. For doing so, we rely on the procedure proposed by Shimotsu (2006), which examines the null hypothesis of long memory against the alternative of a structural change. Two and four subsamples are considered since augmenting the number of hypothetical subsamples does not increase the power of the test. A recent study of Perron and Qu (2010) has been developed to give further insight on this same issue. The results of Perron and Qu (2010) are presented in table 4.

Table 4 shows strong evidence in favour of a short memory process with level shifts and not a true long memory process. Perron and Qu test p-values is less than 10%, thus we reject the alternative hypothesis of true long memory. So, the REER is not a process generated by a fractional model such ARFIMA but threshold models are more suitable. Our results are consisting with the finding of Perron and Qu (2010): they prove a sharp decrease of the estimated value of \( d \) when \( m \) increase implies a spurious long memory. Indeed, in our results (table.2) show that the estimated parameter of long memory process \( (d) \) decrease sharply when \( m \) increase.

<table>
<thead>
<tr>
<th>Statistics of Perron and Qu (2010) test</th>
<th>4.32</th>
<th>0.0000152</th>
</tr>
</thead>
</table>

The empirical literature relating to the PPP hypothesis emphasizes two puzzles, namely the controversial findings about the mean reversion of RER towards the equilibrium and the high degree of persistence of the RER. In our framework, firstly, we prove that both long run and short-run PPP does not hold in the case of Tunisia. This is explained by the rejection of long memory process of REER (short-run PPP rejection) and in the long run through the non-stationary of REER (Long rejection of PPP). As consequence, this finding implies that the linear models are misspecified and cannot be properly used for policy analysis and forecasts by Tunisian policymakers.

Secondly, we show that non-stationary of the Tunisian RER is explained by its nonlinearity propriety traduced by several structural breaks. The presence of structural break traduces that policymakers in Tunisia are unable to react on the exchange rate system through economic policies or via central bank interventions. This finding is important for economic implications. Indeed, the Tunisian authorities, in order to control the exchange rate system, must follow a strategy of the
market-orientation to make the relevance reforms. As economic implications, we questioning all measures implemented by Tunisian authorities in 2012 and 2013 in order to manage the exchange rate. Indeed, in order to promote competitiveness and reliance exportations, Tunisian policymakers hold up the depreciation of dinar. However, these measure seem to be ineffective because they was not accompanied by an improvement of the Tunisian external position. In addition, these measures have generated more distortions because we observe the increase of the gap between the nominal exchange rate and the effective one. These distortions are explained by many facts: the Tunisian exportation has evolved on lower growth rate (10% in 2012) mainly because of slowdown in Europe (-0.5%) and partially braking local activity. Therefore, the dinar’s depreciation has increased the importation by 21.6%, in 2012, and therefore has accurate the inflationary pressures through the mechanism of pass-through to Taylor.

According to our analysis, it is more relevance to Tunisian policymakers to make out some structural reforms to stop the dinar depreciation and they must not act via central bank interventions. In terms of macroeconomic policy, some instruments are no longer operational such as devaluation, since the dinar is partially convertible since 1992. We suggest to Tunisian policy makers to revisit the weight of dinar with euro, in the way of diversification allowing a better managing of the foreign exchange risk and fluctuations.

IV- Conclusion

In recent years, exchange rate modelling has been the subject of several empirical and theoretical investigations by both academicians and practitioners. The issue of exchange rate analysing have become increasingly important, especially for market participants, policymakers and researchers. Moreover, this growing interest on this issue is also motivated by several recent crises; especially for the case of spring Arabic countries where their macroeconomic indicators have downwards in last years. For the case of Tunisia, exchange rate has been depreciated by more than 69% regarding euro negatively affecting competiveness, balance payment and the economic growth. As a consequence, it may be interesting for academicians and practitioners to investigate the properties of Tunisian exchange rate.

In this article, we extend the existing literature on the dynamics of real exchange rate by examining the relevance of structural breaks and long memory in modelling the Tunisian real effective exchange rate. The finding of the presence of a long-memory compound in Tunisian REER has several economic implications on central bank interventions, economic policy. However, the finding of short long memory process with presence of structural break has several economic implications on exchange rate system as reforms and economic policy that must be undertaken. Empirically, three long memory tests are implemented to examine the long-range dependence in the processes of Tunisian
REER. All long memory tests that we used are based on the frequency approach (log-periodogram estimation). While, we apply the Bai Perron (1998, 2003) test in order to detect structural changes in the studied series. In order to discriminate between true long memory or spurious long memory on presence of structural change, we adopt a recent test developed by Perron and Qu (2010).

We prove strong evidence in favour of a short memory process with level shifts and not a true long memory process. So, the REER is not a process generated by a fractional model such ARIFMA but threshold models are more suitable. The presence of structural break traduces that policymakers in Tunisia are unable to react on the exchange rate system through central bank interventions. Therefore, we recommend to Tunisian policymakers to follow a strategy of the market-orientation to make the relevance reforms: It is more relevance to Tunisian policymakers to make out some structural reforms to stop the dinar depreciation. In terms of macroeconomic policy, some instruments are no longer operational such as devaluation, since the dinar is partially convertible since 1992. We suggest to Tunisian policy makers to revisit the weight of dinar with euro, in the way of diversification allowing a better managing of the foreign exchange risk and fluctuations.

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