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THE IMPACT OF STRUCTURAL BREAKS ON THE LONG MEMORY BEHAVIOUR OF EXTERNAL VULNERABILITY INDICATOR: EVIDENCE FROM TURKEY

This paper investigates the long memory behaviour of the ratio of international reserves to short term external debt as an external vulnerability indicator by using classical R/S analysis, modified R/S analysis, GPH test and Robinson test over the period from 1990:01 through 2010:02. In the first part of the analysis, we ignore potential structural breaks in the data. In the second part, the structural breaks identified by Bai and Perron multiple structural break test are taken into account. In both cases, the results of the classical R/S analysis, modified R/S analysis and Robinson test show significant evidence of long memory while the results of GPH test indicate no evidence of long memory. Another finding of the paper is that taking into account structural breaks reduces the integration order.

Keywords: long memory, structural breaks, detrending, external vulnerability
JEL Classification: C20, H60

INTRODUCTION

Recent economic crises in emerging economies have led economists and policy makers to find alternative indicators for examining the characteristics of these crises. In this context, special attention has been placed on the international reserves to short term external debt ratio. Furman and Stiglitz (1998) and Radelet and Sachs (1998) examine the importance of this ratio and conclude that international reserves to short term external debt ratio is one of the determining factors of the Asian crises. Bussiere and Mulder (1999), Rodrik and Velasco (1999) and De Beaufort Wijnholds and Kapteyn (2001) support the importance of this ratio as a liquidity indicator of the economy. In addition, the ratio of international reserves to short term external debt is incorporated into the indicator series used in early warning systems of the International Monetary Fund (IMF) (IMF, 2002).

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Since reserves and external debt affect a country's external vulnerability through their impact on the country's ability to discharge external obligations, the international reserves to short term external debt ratio of a country is also used as a vulnerability indicator. A country with a low international reserves to short term external debt ratio is more vulnerable to speculative attacks or external shocks due to the more limited availability. A low ratio indicates that imprudent macroeconomic policies are being pursued and an economic crisis will tend to be more severe (Guzman Calafell and Padilla del Bosque, 2002).

The aim of this paper is to investigate the long memory behaviour of the international reserves to short term external debt ratio in Turkey as an external vulnerability indicator over the period from 1990:01 through 2010:02. It is clear that the period of our analysis covers very important economic and financial crises in the Turkish economy. The analysing of the long memory behaviour of this ratio also gives evidence on the reaction of this indicator to the economic and financial crises. Thus, we can see the usefulness of this indicator in predicting these crises.

It is known that the economic crisis started to affect the Turkish economy with increasing frequency during the 1990s. After the adoption of the Structural Adjustment and Stabilization Program, which is called 24 January 1980 decisions, Turkey liberalized its economy to integrate with the world economy and a lot of new laws were passed to liberalize foreign trade and financial transactions (Çepni and Köse, 2006). Import substitute industrialization strategy was replaced by an export-led growth strategy that relies more in a market-based economy. Policies based on adjustments upon tariffs rather than quantity restrictions were adopted, and also protection rates in imports regime were steadily lowered. Besides, export licenses were abolished, and export liberalization was put in effect as a major policy issue in the Turkish economy politics (Varol, 2003). In 1989, capital account liberalization was completed, which enabled domestic residents and private firms to borrow freely on the international financial markets. This allowed residents to make financial transactions in foreign currencies and nonresidents were allowed to invest freely in the domestic markets. After switching to a free floating regime, Turkey started to encounter more frequent crises. The main reasons for these crises were (Çepni and Köse, 2006): the development of an unsustainable domestic debt and a dynamic and unhealthy financial sector structure with particular problems caused by the state banks and by the failure of structural problems. In early 1994, a post-liberalization financial crisis started, followed by another stabilization program that was supported with an IMF stand-by agreement in April 1994. The contagion effects of the Asian and particularly the Russian crises brought about a protracted crisis in Turkey in 1998 and 1999 (Uygur, 2010). On February 2001, Prime Minister Bülent Ecevit announced that there was a severe political crisis, which ignited an equally serious economic crisis in the highly sensitive markets. On that day in February, overnight rates jumped to unprecedented levels of 6,200 percent. Three days later the exchange rate system collapsed, and Turkey declared that it was going to implement a floating exchange rate system (Özatay and Sak, 2003). Another important crisis which occured over the analysis period is the 2008 global crisis. It has had major negative effects on the economies of all countries in the world. Turkey was also one of the seriously affected countries by the global crisis.

In the light of this information, we investigate the long memory behaviour of the international reserves to short term external debt ratio in Turkey by using classical rescaled-range (R/S) analysis, modified R/S analysis, Geweke and Porter-Hudak (GPH) test and Robinson test. Our analysis covers two parts: in the first part, the long memory behaviour of the external vulnerability indicator is examined by ignoring potential structural breaks. In the second part, we identify possible structural breaks in the data by using the Bai and Perron multiple structural break test. After finding significant structural breaks, we detrend the data and apply long memory methods by using the detrended series. Following this analysis with these two parts provides us to see the impacts of structural breaks on the long memory behaviour of the external vulnerability indicator. In order to support the studies which conclude that the international reserves to short term external debt ratio is an important indicator for examining the characteristics of the crises, we expect that the structural breaks identified in the second part refer to the economic and financial crises of the Turkish economy.

The paper is constructed as follows: Section 2 gives details of the methodology, Section 3 describes the data and presents the empirical results. Finally, Section 4 gives our conclusions.

1. METHODOLOGY

In this section, we briefly describe the methods which are employed for the long memory analysis of the ratio of international reserves to short term external debt. Before giving the methodological descriptions of these methods, we need to describe the long memory property first. The long memory property describes the high-order correlation structure of a series. If a series exhibits long memory, there is a persistent temporal dependence even between distant observations. Such series are characterized by distinct but non-periodic cyclical patterns. A stationary stochastic y_t process is defined as a long memory process if there exists a real number $\alpha \in (0,1)$ and a constant $C > 0$ such that $\lim_{k\to\infty} p(k)\big/\left[Ck^{-\alpha}\right] = 1$ where $p(k)$ is the autocorrelation function. The Hurst exponent H, which represents the long memory property of the time series, is defined as $H = 1-\alpha/2$. Thus, a time series is said to exhibit long memory if there is a slow hyperbolic decay in autocorrelations. The H takes values from 0 to 1 ($0 \le H \le 1$). Long memory occurs when $0.5 < H < 1$. If $0 < H < 0.5$, the series is called anti-persistent. A long memory time series is said to be fractionally integrated where the fractional differencing parameter d is related to the parameter H as $d = H - 1/2$. It is clear that the process is a long memory process when $0 < d < 1$. In the econometric and statistical literature, various tools for detecting possible long memory in time series are developed. In this paper, by using related Gauss program codes, we apply the following methods: classical R/S analysis, the modified R/S analysis, the Geweke and Porter-Hudak test and the Robinson test. It is important to note that the findings on long memory can be vary substantially depending on the method and therefore checking is crucial.

1.1. Classical R/S and Modified R/S Analyses

The R/S statistic which was proposed by Mandelbrot (1972) and originally developed by Hurst (1951) can be defined as follows:

$$\left(R/S\right)_t = \frac{R_t}{S_t} = ct^H \tag{1}$$

where $\left(R/S\right)_t$ is the rescaled range statistic measured over a time index t, c is a constant and H is the Hurst exponent. In a detailed form, R_t and S_t can be written as:

$$R_t = \max_{0 \le j \le t} \left\{ \sum_{j=1}^{t} (y_j - \overline{y}) \right\} - \min_{0 \le j \le t} \left\{ \sum_{j=1}^{t} (y_j - \overline{y}) \right\} \tag{2}$$

$$S_t = \left\{ (1/t) \sum (y_j - \overline{y})^2 \right\}^{1/2} \tag{3}$$

Here, R_t is the range, S_t is the sample standard deviation and $\overline{y}$ is the sample mean. This analysis is relatively easy to implement but it suffers from two disadvantages. One of them is about the statistical properties of the estimators. This analysis does not lead itself to hypothesis testing. And the second is that the results may be biased by the presence of short memory effects in addition to long ones. To overcome these disadvantages, an improved version of the classical R/S statistic (modified R/S statistic) is proposed by Lo (1991) who argues that the classical R/S statistic is extremely sensitive to "short range dependence" and heteroskedasticity. Lo's modified version considers as well this short range dependence by performing a Newey-West correction using the Bartlett window to derive a consistent estimate of the long range variance of the time series. In this improved version, the null hypothesis of short memory (i.e. $d = 0$) is tested against the long memory ($d > 0$) and/or antipersistence ($d < 0$). Lo's modified R/S statistic is as follows:

$$Q_t(q) = R_t / \delta_t(q) \tag{4}$$

Lo (1991) defines R_t as in Equation (2). But, instead of simply using sample standard deviation, S_t, to normalize R_t, Lo uses a weighted sum of autocovariances as can be seen below:

$$\delta_t^2(q) = S_t^{\,2} + 2 \sum_{j=1}^{q} w_j(q) \gamma_j \, . \tag{5}$$

Here, γ_j is the jth order sample autocovariance of y_t, $w_j(q)$ is the Bartlett window weight of $w_j(q) = 1 - \left(\frac{j}{q+1} \right)$, for $q < T$ and S_t^2 is the sample variance. The critical values for this test are given by Lo (1991). The statistic can be further normalized as:

$$V_t(q) = Q_t(q) \big/ \sqrt{t} \tag{6}$$

The advantage of the modified R/S statistic is that it allows us to obtain a simple formula for the fractional differencing parameter d:

$$d = \log\left(Q_t(q)\right) \big/ \log(t) \tag{7}$$

The classical R/S statistic corresponds to $q = 0$, so that Equation (5) does not contain the second term. This second term was suggested by Lo (1991) to take into account short range dependence. Whereas the classical R/S analysis focused on estimating the limit of the ratio $\log\left(Q_t(0)\right)\big/\log(t)$, called the Hurst coefficient, Lo (1991) proposed a statistical hypothesis testing procedure to detect long memory. It should be noted that the asymptotic distribution of the statistic $Q_t(0)$ depends strongly on the correlation structure of the data and is not asymptotically parameter free, so it cannot be used to construct a test (Giraitis et al., 2003). If $q = 0$, Lo's statistic reduces to classical R/S statistic. This statistic is highly sensitive to the order of truncation q but there are no statistical criteria for choosing q in the framework. If q is too small the statistic does not account for the autocorrelation of the process, while if q is too large it accounts for any form of autocorrelation and the power of this test tends to its size. Given that the power of a useful test should be greater than its size, this test is not very helpful. For that reason, Teverovsky et al. (1999) suggest to use this statistic with other tests.

1.2. Geweke and Porter-Hudak Test

As an alternative method, we examine the long memory behaviour in the series by using fractionally integrated process based on a fractionally differencing parameter d. The fractionally integrated time series process, proposed by Granger and Joyeux (1980) and Hosking (1981) can be described by the following stochastic equation:

$$(1-L)^d y_t = u_t, \quad t = 1,2,.... \tag{8}$$

where L is the lag operator and u_t is a stationary ($I(0)$) process. Here, d can take any real value, y_t is known as a fractionally integrated process.

In general y_t is called an $I(d)$ process. The arbitrary restriction of d to integer values gives rise to the standard autoregressive integrated moving average (ARIMA) model. If $d=0$ in Equation (8), $y_t = u_t$ and a "weakly autocorrelated" y_t is allowed for. When $d>0$, y_t is said to be "strongly autocorrelated" or "strongly dependent". When $d=1$, y_t is known as a unit root process. When $d<1$, the process y_t is said to be a mean reverting process. If $0<d<1$, the process is a long memory process. If $0.5<d<1$, the process is nonstationary and exhibits long memory or long range dependence while the process is stationary and exhibits long memory, if $0<d<0.5$. It is important to note that when $d<0.5$, the process is stationary as well as mean reverting and when $0.5 \le d$ the process is non-stationary even if the fractional parameter is significantly less than 1. The process is said to exhibit intermediate memory (anti-persistence) or long-range negative dependence, for $-0.5<d<0$.

In order to estimate the fractional differencing parameter d, Geweke and Porter-Hudak (1983) developed a nonparametric test. They show that differencing parameter d, which is called a long memory parameter and can be estimated consistently from the least squares regression:

$$\ln(I(w_j)) = \theta - d\ln(4\sin^2(w_j/2)) + v_j\ , \quad j=1,\dots,J \tag{9}$$

where θ is a constant, $w_j = 2\pi j/T$ ($j=1,\dots,T-1$) denotes the Fourier frequencies of the sample, $J=f(T^{\mu})$ is an increasing function of T which is the number of observations and $0<\mu<1$. $I(w_j)$ is the periodogram of the series at frequency w_j. In empirical analysis, $J=f(T^{\mu})$ is used with μ ranging from 0.5 to 0.7. Since one can choose different values of μ, different estimates of the fractional parameter for the same process can be obtained. The GPH test is carried out on the first differences of the series to ensure that stationarity and invertibility are achieved. The differencing parameter in the first differenced data is denoted by $\tilde{d}$ in which case the fractional differencing parameter for level series is $d=1+\tilde{d}$. The existence of fractional order of integration can be tested by examining the statistical significance of the differencing parameter. If the OLS estimator $\hat{d}$ is significantly different from zero, then the time series are fractionally integrated, and thus exhibit long memory process.

1.3. Robinson Test

To obtain efficient results in this paper, we also employ different versions of the Robinson (1994) test. The main advantage of this procedure is that it tests unit and fractional roots with a standard null limit distribution, which is unaffected by the inclusion or not of deterministic trends. Robinson (1994) considers the following regression model,

$$y_t = \beta' z_t + x_t, \quad t = 1, 2, ... \tag{10}$$

where y_t is the observed time series for $t = 1, 2, ...T$, $\beta = (\beta_1, ..., \beta_k)'$ is a $(k \times 1)$ vector of unknown parameters, z_t is a $(k \times 1)$ vector of deterministic regressors such as an intercept or a linear trend. And the regression errors x_t can be explained as follows:

$$(1 - L)^d x_t = u_t, \quad t = 1, 2, ... \tag{11}$$

where L is the lag operator and u_t is a stationary process. Here, d can take any real value. If $d > 0$, x_t is said to be long memory (Granger and Joyeux, 1980; Hosking, 1981). The process is nonstationary and exhibits long memory if $0.5 < d < 1$. Clearly, the unit root case corresponds to $d = 1$ in Equation (11). If $0 < d < 0.5$, the process is stationary and exhibits long memory. When $d < 0.5$, the process is stationary as well as mean reverting with the effects of the shocks dying away in the long run. On the other hand, the process is non-stationary even if the fractional parameter is significantly less than 1, when $0.5 \leq d$.

The Lagrange Multiplier (LM) test suggested by Robinson (1994), tests unit roots and other forms of nonstationary hypotheses embedded in fractional alternatives. The null hypothesis of the test is as follows:

$$H_0 : d = d_0 \tag{12}$$

The test statistic can be described by:

$$\hat{r} = \frac{T^{1/2}}{\hat{\sigma}^2}\hat{A}^{-1/2}\hat{a} \tag{13}$$

Here, T is the sample size and

$$\hat{a} = \frac{-2\pi}{T}\sum_{j=1}^{T-1}\psi(\lambda_j)g(\lambda_j;\hat{\tau})^{-1}I(\lambda_j) \; ;$$

$$\hat{\sigma}^2 = \sigma^2(\hat{\tau}) = \frac{2\pi}{T}\sum_{j=1}^{T-1} g(\lambda_j;\hat{\tau})^{-1}I(\lambda_j);$$

$$\hat{A} = \frac{2}{T}\left(\sum_{j=1}^{T-1}\psi(\lambda_j)^2 - \sum_{j=1}^{T-1}\psi(\lambda_j)\hat{\varepsilon}(\lambda_j)' \times \left(\sum_{j=1}^{T-1}\hat{\varepsilon}(\lambda_j)\hat{\varepsilon}(\lambda_j)'\right)^{-1} \times \sum_{j=1}^{T-1}\hat{\varepsilon}(\lambda_j)\psi(\lambda_j)\right)$$

$$\psi(\lambda_j) = \log\left|2\sin\frac{\lambda_j}{2}\right|; \qquad \hat{\varepsilon}(\lambda_j) = \frac{\partial}{\partial\tau}\log g(\lambda_j;\hat{\tau}_j); \qquad \lambda_j = \frac{2\pi j}{T};$$

$$\hat{\tau} = \arg\min_{\tau\in T^*}\sigma^2(\tau)$$

where T^* is a compact subset of the Euclidean space. Here, $I(\lambda_j)$ is the periodogram of $\hat{u}_t$ (obtained from the Equation (11)) where,

$$\hat{u}_t = (1-L)^{d_0}y_t - \hat{\beta}w_t$$

$$w_t = (1-L)^{d_0}z_t$$

$$\hat{\beta} = \left(\sum_{t=1}^{T} w_t w_t'\right)^{-1}\sum_{t=1}^{T} w_t(1-L)^{d_0}z_t \, .$$

Robinson (1994) showed that the test statistic under certain regularity conditions satisfies:

$$\hat{r} \rightarrow_d N(0,1) \text{ as } T \rightarrow \infty. \tag{14}$$

Thus, a one sided $100\alpha\%$ level test of $H_0 : d = d_0$ against the alternative $H_1 : d > d_0$ is given by the rule "Reject H_0 if $\hat{r} > z_\alpha$" where the probability that a standard normal variate exceeds z_α is α and conversely, a one sided $100\alpha\%$ level test of $H_0: d = d_0$ against the alternative $H_1: d < d_0$ is given by the rule "Reject H_0 if $\hat{r} < -z_\alpha$".

2. DATA AND EMPIRICAL RESULTS

This paper examines the long memory properties of the quarterly international reserves to short term external debt ratio (IR/SED) as an external vulnerability indicator over the period from 1990:01 through 2010:02. The source of the data is the Central Bank of the Republic of Turkey (CBRT). Before the analysis, we examine the possible seasonal effect on the series by using seasonal dummy variables. Since the dummy variables are not statistically significant, we use seasonally unadjusted data in the analysis. The illustration of the data can be seen in Figure 1.

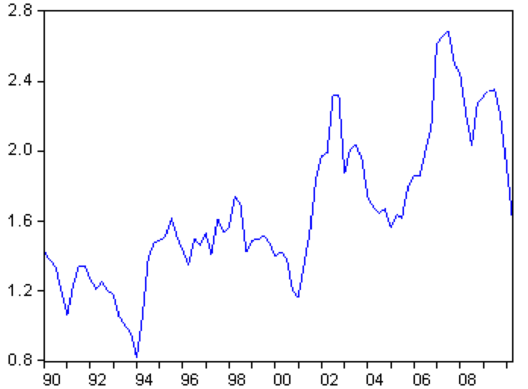

Figure 1. The plot of the ratio of international reserves to short term external debt

Source: The Central Bank of the Republic of Turkey

As can be seen from the figure, there are possible structural breaks in the structure of the international reserves to short term external debt ratio. Several works, including Granger and Hyung (1999) and Diebold and Inoue (2001), argue that structural breaks or regime switching can generate spurious long memory behaviour in the observed time series. In other words, the long memory property in the data may be due to the presence of structural breaks or regime switches. This is called "the spurious long memory process". In order to avoid the spurious long memory problem, we need to take into account potential structural breaks. In the first part of the analysis we apply the long memory methods by ignoring structural breaks. In the second part we identify the structural breaks by using Bai and Perron (1998, 2003) multiple structural break test, obtain detrended series and repeat the analysis by using new series. As a first step, we apply Augmented Dickey Fuller (ADF), Philips and Perron (PP) and Kwiatkowski-Phillips-Schmidt and Shin (KPSS) unit root tests in order to check for stationarity. These tests differ in the null hypothesis. The null hypothesis of the ADF and PP tests is that a time series contains a unit root, I(1) process, while the KPSS test has the null hypothesis of stationarity, I(0) process. The results of unit root tests under the different null hypothesis are characterized by four possible outcomes (Barkoulas et al, 1997): 1) When we reject the null hypothesis of the ADF and PP tests and we cannot reject the null hypothesis of the KPSS test, a time series is stationary. 2) Conversely, failure to reject a unit root by ADF and PP tests and the rejection of stationarity by KPSS test supports that a time series is nonstationary. 3) Failure to reject a unit root and stationary null hypotheses shows that the series are not sufficiently informative with respect to the low frequency properties. 4) Rejection of null hypotheses indicates that the series are not well represented as either I(0) and I(1), which indicates that the series appear to be a long term dependence process. The results of the ADF, PP and KPSS unit root tests are reported in Table 1.

Table 1

Unit root test results

	ADF	PP	KPSS
IR/SED	-2.122	-1.883	0.960***
Δ IR/SED	-6.770***	-6.782***	0.072

The test statistics are obtained under the case with an intercept and a linear trend.

*** denotes that the null hypotheses are rejected at the 1% significance level.

Source: author's own

As can be seen from the table, the results of unit root tests correspond to the second case, implying that the IR/SED ratio is nonstationary in level but stationary in the first difference, that is, it is integrated of order one, I(1). In the second step, we examine the long memory behaviour of the ratio of IR/SED by performing different methods: the classical R/S analysis, the modified R/S analysis, the Geweke and Porter-Hudak test and the Robinson test. The results of the classical R/S and modified R/S analyses are tabulated in Table 2.

Table 2

The results of classical R/S and modified R/S analyses

	Statistics
Classical R/S	3.593***
Modified R/S	2.583***

The null hypothesis of short memory process is rejected if the R/S statistic does not fall within the confidence interval [0.809,1.862] at the 5% significance level and [0.721,2.098] at the 1% significance level.

*** indicates the statistical significance at the 1% level.

Source: author's own

The results in the table show that the null hypothesis of short memory is rejected, implying that there is significant evidence of long memory in the series. Then, we also perform the GPH test in order to check the long memory behaviour. The reason is that the GPH test may make available a further powerful test for long memory than the classical and modified R/S analyses. Since the IR/SED ratio is found to be nonstationary according to the unit root tests, the GPH test is carried out on the first difference of the series. Table 3 reports the GPH test results for $\mu = 0.50, 0.55, 0.60, 0.65$ and 0.70.

Table 3

The results of GPH test

μ	$\tilde{d}$	t - statistics
0.50	-0.197	-0.620
0.55	-0.151	-0.550
0.60	0.059	0.241
0.65	0.240	1.165
0.70	0.127	0.703

The null hypothesis $H_0 : \tilde{d} = 0$ is tested against the alternative hypothesis $H_1 : \tilde{d} \neq 0$.

Source: author's own

The results from the GPH test indicate that there is no significant evidence of long memory. According to Agiakloglou et al. (1993), the GPH method is appropriate for stationary long memory process with $-0.5 < d < 0.5$ and the estimator is not invariant to first differencing, so that there might be bias due to over-differencing of the data. Since the GPH test has some problems, we consider using the Robinson (1994) tests which have several distinguishing features compared with other procedures. It is also known that they have a standard null limit distribution and they are the most efficient ones when directed against the appropriate alternatives. Following this way, we test $H_0 : d = d_0$ for different values of d_0 in the case of a linear trend with white noise and AR(1) disturbances. The one-sided test statistics given by $\hat{r}$ in (13) are reported in Table 4. For a given d_0, significantly positive values of $\hat{r}$ are consistent with an order of integration higher than that, whereas significantly negative ones are consistent with smaller orders of integration.

Table 4

The results of Robinson test

d_0	**White noise disturbances**	**AR(1) disturbances**
0	6.75	25.87
0.05	6.41	24.02
0.10	6.07	22.26
0.15	5.74	20.58
0.20	5.41	18.97
0.25	5.08	17.43
0.30	4.76	15.95
0.35	4.43	14.55
0.40	4.10	13.21
0.45	3.78	11.93
0.50	3.45	10.73
0.55	3.12	9.60
0.60	2.80	8.53
0.65	2.47	7.53
0.70	2.14	6.59
0.75	1.81	5.70
0.80	1.48**	4.85
0.85	1.16**	4.03
0.90	0.84**	3.24
0.95	0.53**	2.46
1.00	0.22**	1.68
1.05	-0.08**	0.91**
1.10	-0.37**	0.14**

** indicates nonrejection values of the null hypothesis at the 95% significance level
Source: author's own

The results in the table indicate that $H_0 : d = d_0$ cannot be rejected for $d_0 = 0.80, 0.85, 0.90, 0.95, 1, 1.05$ and 1.10 in the case of white noise disturbances and for $d_0 = 1.05$ and 1.10 in the case of AR(1) disturbances.

These results mean that the IR/SED ratio follows a non-stationary process with long memory. In the second part of the analysis, the potential structural breaks in the IR/SED ratio are considered following Granger and Hyung (1999) and Diebold and Inoue (2001). They argue that the long memory property in the data may be due to the presence of structural breaks or regime switches. In order to determine the structural breaks, Bai and Perron (1998, 2003) multiple structural break test which allows to test for multiple breaks at unknown dates is used. Before reporting the obtained results we need to describe the multiple structural break test of Bai and Perron. They follow the multiple structural break model with m breaks ($m+1$ regimes) as below:

$$
\begin{aligned}
y_t &= x_t^{'}\beta + z_t^{'}\delta_1 + u_t, \quad t = 1,...,T_1 \\
y_t &= x_t^{'}\beta + z_t^{'}\delta_2 + u_t, \quad t = T_1+1,...,T_2 \\
&\cdots\cdots\cdots\cdots\cdots\cdots \\
y_t &= x_t^{'}\beta + z_t^{'}\delta_{m+1} + u_t, \quad t = T_m+1,...,T.
\end{aligned}
\tag{15}
$$

where y_t is the observed dependent variable at time t, x_t is $(p\times 1)$ and z_t is $(q\times 1)$ and β and δ_j ($j = 1,...,m+1$) are the corresponding coefficient vectors, and u_t is the disturbance term at time t . Here, T is the sample size and $T_1 < T_2 < ... < T_m < T$. The estimation method is based on the least squares principle. For each m -partition ($T_1,...,T_m$), denoted $\{T_j\}$, the associated least squares estimate of δ_j is obtained by minimizing the sum of squared residuals $\sum_{i=1}^{m+1}\sum_{t=T_{i-1}+1}^{T_i}(y_t - z_t^{'}\delta_i)^2$. Bai and Perron (1998, 2003) suggest several statistics for the consistent estimation of the number and location of breakpoints ($T_1,...,T_m$) and the parameters ($\delta_1^{'},...,\delta_{m+1}^{'}$):

- $SupF_T(k)$ test, i.e., a $SupF$ - type test of the null hypothesis of no structural break *versus* the alternative of a fixed number k of breaks.
- Two maximum tests of the null hypothesis of no structural break *versus* the alternative of an unknown number of breaks given some upper bound, i.e.,$UD_{\max}$ test, an equal weighted version, and $WD_{\max}$ test, with weights that depend on the number of regressors and the significance level of the test.

- The $SupF_T(l+1|l)$ test, i.e. a sequential test of the null hypothesis of l breaks *versus* the alternative of ($l+1$) breaks.

The asymptotic distributions of these three tests are derived in Bai and Perron (1998) and asymptotic critical values are tabulated in Bai and Perron (1998, 2003) for $\varepsilon = 0.05$ $(M = 9)$, 0.10 $(M = 8)$, 0.15 $(M = 5)$, 0.20 $(M = 3)$, and 0.25 $(M = 2)$ (For more details, see Bai and Perron (1998, 2003). In our analysis, the maximum permitted number of breaks is set at $M = 5$ and a trimming $\varepsilon = 0.15$ is used to determine the minimal number of observations in each segment [$h = [\varepsilon T]$ with the sample size T]. The findings are tabulated in Table 5.

Table 5
The multiple structural break test results of Bai and Perron

Tests	Hypothesis	Statistics
$SupF_T(k)$ Test:	H_0: 0 break *vs* H_1: 1 break	6.12
	H_0: 0 break *vs* H_1: 2 breaks	5.44
	H_0: 0 break *vs* H_1: 3 breaks	6.55^{**}
	H_0: 0 break *vs* H_1: 4 breaks	5.30^{**}
	H_0: 0 break *vs* H_1: 5 breaks	5.37^{***}
UD_{max} Test:	H_0: 0 break *vs* H_1: an unknown break	6.55
WD_{max} Test:	H_0: 0 break *vs* H_1: an unknown break	11.79^{**}
$SupF_T(l+1 \mid l)$ Test:	$SupF_T(2 \mid 1)$ $SupF_T(3 \mid 2)$ $SupF_T(4 \mid 3)$ $SupF_T(5 \mid 4)$	7.06 7.03 1.53 0.00
Number of Breaks:	BIC: 3 LWZ:3	
Break Dates:		1994:02, 2001:03, 2006:03

*** and ** denote that the test statistics are significant at 1% and 5% levels, respectively. The number of breaks are chosen by the Bayesian information criterion, BIC (Yao, 1988); and the Schwarz modified criterion, LWZ (Liu, Wu and Zidek, 1997).

Source: author's own

The results in the table show that $SupF_T(k)$ tests are significant for k = 3, 4 and 5. Although the $UD_{\max}$ statistic is insignificant, the $WD_{\max}$ statistic is highly significant which implies that at least one break is present in the data. On the other hand, all the $SupF_T(l+1|l)$ statistics are found insignificant but BIC and LWZ criterions refer to three possible breaks around 1994:02, 2001:03 and 2006:03. It is clear that these breaks occur during the economic, financial and global crises periods in Turkey. This finding supports that the IR/SED ratio can be used as an important indicator for the crises. By using the obtained breaks, we detrend series and repeat the classical R/S analysis, the modified R/S analysis, the Geweke and Porter-Hudak test and the Robinson test in order to examine the impact of structural breaks on the long memory behaviour of the IR/SED ratio. The results of classical R/S and modified R/S analyses for detrended IR/SED ratio can be seen in Table 6.

Table 6
The results of classical R/S and modified R/S analyses for detrended series

	Statistics
Classical R/S	3.500^{***}
Modified R/S	2.550^{***}

The null hypothesis of short memory process is rejected if the R/S statistic does not fall within the confidence interval [0.809,1.862] at the 5% significance level and [0.721,2.098] at the 1% significance level.

*** indicates the statistical significance at the 1% level.

Source: author's own

As reported in the table, the classical R/S and modified R/S statistics do not fall within the confidence interval at the 1% significance level. These results indicate that the null hypothesis of short memory is rejected and significant evidence of long memory is found. It can be said that there is no significant impact of structural breaks on the long memory behaviour of the IR/SED ratio according to classical R/S and modified R/S analyses. In order to compare the difference, we also repeat the GPH test by using detrended series and report the results in Table 7.

Table 7
The results of GPH test for detrended series

μ	$\tilde{d}$	t - statistics
0.50	0.132	0.771
0.55	0.074	0.268
0.60	-0.034	-0.141
0.65	-0.136	-0.660
0.70	-0.102	-0.562

The null hypothesis $H_0 : \tilde{d} = 0$ is tested against the alternative hypothesis $H_1 : \tilde{d} \neq 0$.

Source: author's own

As can be seen from the table, the GPH test results for μ = 0.50, 0.55, 0.60, 0.65 and 0.70 show that the null hypothesis of short memory cannot be rejected and significant long memory cannot be found. These results are consistent with the previous GPH test results which do not take into account the impacts of possible structural breaks. Since Agiakloglou et al. (1993) argue that GPH test has some problems, we also consider to apply Robinson (1994) test for detrended IR/SED ratio and tabulate the results for different values of d_0 with white noise and AR(1) disturbances in Table 8.

Table 8

The results of Robinson test for detrended series

d_0	**White noise disturbances**	**AR(1) disturbances**
0	10.50	55.50
0.05	9.77	48.01
0.10	8.99	40.78
0.15	8.18	34.07
0.20	7.33	28.08
0.25	6.47	22.90
0.30	5.60	18.51
0.35	4.75	14.86
0.40	3.93	11.84
0.45	3.14	9.32
0.50	2.40	7.20
0.55	1.71	5.36
0.60	1.07**	3.72
0.65	0.48**	2.23
0.70	-0.05**	0.83**
0.75	-0.53**	-0.47**
0.80	-0.96**	-1.67
0.85	-1.35**	-2.77
0.90	-1.69	-3.76
0.95	-2.00	-4.63
1.00	-2.28	-5.40
1.05	-2.52	-6.07
1.10	-2.74	-6.66

** indicates nonrejection values of the null hypothesis at the 95% significance level.

Source: author's own

According to the Robinson test results, it can be seen that the non-rejection values take place at $d_0 = 0.60, 0.65, 0.70, 0.75, 0.80$ and 0.85 for white noise disturbances. On the other hand, $H_0 : d = d_0$ cannot be rejected at $d_0 = 0.70$ and 0.75 for AR(1) disturbances. The null hypothesis of unit root ($d_0 = 1$) is rejected in both cases. These results indicate that the detrended IR/SED ratio follows a nonstationary process with long memory. If we compare these results with the previous Robinson test results without structural breaks, it is seen that the nonrejection values are smaller than before. It can be said that taking into account the impact of structural breaks significantly reduces the integration order. On the other hand, it is clear that the identified structural breaks refer to the economic, financial and global crises periods of the Turkish economy. Since the structure of international reserves to short term external debt ratio is very sensitive to the structural breaks, this ratio can be used by policy makers in predicting economic and financial crises for Turkey.

CONCLUSIONS

In this paper, we examine the long memory behaviour of the international reserves to short term external debt ratio as an external vulnerability indicator over the period from 1990:01 through 2010:02 by using different methods, namely: the classical R/S analysis, the modified R/S analysis, the GPH test and the Robinson test. The analysis is conducted in two parts. In the first part, the long memory methods are applied by ignoring the potential structural breaks in the data. As a second part, the possible strutural breaks are identified by using Bai and Perron (1998, 2003) multiple structural break test. After finding three significant structural breaks around 1994:02, 2001:03 and 2006:03, we obtain detrended series and repeat the long memory methods. The obtained results from classical R/S analysis, modified R/S analysis and Robinson test indicate that the ratio of international reserves to short term external debt has long memory in both cases. On the other hand, the GPH test results show no evidence of long memory for these two cases. Another finding of the paper is that taking into account structural breaks reduces the integration order of the external vulnerability indicator. Our results support the usefulness of the international reserves to short term external debt ratio in predicting economic, financial and global crises periods. Policy makers can use this ratio as an indicator for the Turkish economy.

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