Foreign exchange reserves and exchange rates in Turkey: Structural breaks, unit roots and cointegration

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Accepted 27 April 2007

Abstract

This paper investigates the relationship between exchange rates (nominal and real) and foreign exchange reserves in Turkey, using monthly data over the period 1982:1–2005:11. Unit root and cointegration tests, which allow for structural breaks are used. The results indicate that there is a long-run relationship between foreign exchange reserves and exchange rates. The results also suggest that the direction of both long and short-run causality is from foreign exchange reserves to real effective exchange rate. As for the relationship between nominal exchange rate and foreign exchange reserves, the results suggest that in the long-run nominal exchange rate Granger cause foreign exchange reserves.

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JEL classification: E44; F31

Keywords: Exchange rates; Foreign exchange reserves; Structural breaks

1. Introduction

Although critics argue that large foreign exchange reserve holding is a costly self-insurance strategy, the precautionary holdings of substantial foreign exchange reserves in emerging countries are motivated by many factors. Countries would hold more foreign exchange reserves in order to smooth the vulnerability impacts of the domestic and foreign shocks, make exchange rate arrangements to correct international payment difficulties, implement the foreign debt service, increase the international credibility of the country and cope with the speculative attacks.2 The cost of holding more reserves can be ignored compared to the consequences of currency crises in emerging economies.3 The demand for reserves mainly depends on the exchange rate regime, the size of economy, trade and financial openness, financial deepening, capital controls, the level of external debt and political instability (see Lane and Burke, 2001; Gupta and Agarwal, 2004).

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2 The total amount of global reserves is $2.2 trillion at the end of May 2002 and countries in East Asia held around 38% of the world total reserves (see Aizenman and Marion, 2003a,b). By the end of 2003, nearly 25% of total international reserves are accumulated by China, Korea, Malaysia and Thailand (see Garcia and Soto, 2004).

3 In these economies, the sudden changes in exchange rates have triggered currency and banking crises by affecting the financial and macroeconomic variables in recent years.

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doi:10.1016/j.econmod.2007.04.010
On January 24, 1980, the Turkish government announced an economic stabilization and reform program, which included both trade and financial liberalization. The main goal of the financial and trade liberalization program was to promote financial market development and to move towards a liberalized trade regime on the basis of export-led growth strategy with a flexible exchange rate, respectively. Adoption of an actively managed, flexible exchange rate regime was an important step undertaken in the early 1980s, and it has remained a central instrument in the trade liberalization program. International capital movements and foreign exchange operations were subject to tight controls before 1980, they were entirely liberalized and the TL (Turkish Lira) became convertible in 1989.

As seen in Fig. 1, for most of the 1980s there was a little change in Turkish foreign exchange reserves. However, reserves showed an upward trend that fuelled by significant increases in capital inflows after the liberalization of foreign exchange operations and international capital movements. The high reserves reinforced the efficiency of monetary and exchange rate policies and therefore, served as an intervention and credibility tool. To finance the budget deficits and to boost the competitiveness of exporters, a high real interest rate and an undervalued Turkish Lira were the dominant policies by the monetary authority in the early 1990s. These policies have accelerated the short-term capital inflows. The huge amount of inflows then has caused the overvaluation of the Lira within financial and fiscal fragilities. The authorities lost the control in financing fiscal deficit and the first major financial crisis, which was also an early warning signal for preceding financial crises in Turkey, occurred in January 1994 and the second occurred in February 2001. Both crises were followed by rapid reserve depletions, large amount of unexpected devaluations and finally switch to a new exchange rate regime (see Togan and Erasel, 2004; Berument and Dinçer, 2004; Özla and Yeldan, 2004; the Central Bank of the Republic of Turkey (hereafter CBRT), 2005). Fig. 1 also shows the impact of financial crises on the real exchange rate and foreign exchange reserves.

4 Several exchange rate regimes have been adopted in Turkey since the stabilization program of 1980. The exchange rates were determined under the flexible exchange rate system in the form of a “crawling band” for the period 1980–1981 and in the form of a “managed float” between 1981 and 1999. The managed float was transformed into a “dirty float” in 1994. A “crawling peg” regime was adopted under a standby agreement with the IMF in 1999 to control the chronic and high inflation rates. The regime had a short life due to the financial crisis of 2001. Exchange rates have been determined under a “flexible exchange rate” regime since February 2001.

5 Before 1980, foreign exchange reserves were used to meet the short-term needs under “fixed exchange rate” regime with strict foreign exchange controls. With the stabilization program, the aim of coherence between inflation expectation and exchange rate enforced the accumulation of reserves. In this period, reserves were fuelled by the sizeable increase in accumulation of banks, institutions and households and mainly helped to manage the current account purposes until 1989.

6 The crisis of 1994 began in the financial sector and later spread to the real sector. At the end of 1994, the Turkish economy shrunk by 6% and the inflation rate hit three digit levels. As a result, the value of US dollar nearly doubled against TL and about half of the foreign exchange reserves were eroded in managing the crisis. As for the 2001 crisis, the problem started with the disinflation program of 1999. After a relatively small crisis in November 2000, the Turkish financial system got into a deepening crisis period that reached to its peak with the abandonment of the “crawling peg” regime in February 2001. A rapid reserve depletion continued in the whole year, even though a regime of free float was adopted instantaneously.

7 Both real exchange rate and foreign exchange reserves are in natural logarithm.

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Fig. 1. Turkey’s foreign exchange reserves and real exchange rate: Jan. 1982–Nov. 2005.
Studies on the Turkish currency have been relatively few compared to the currencies of developed countries. The existing studies have mostly focused on exchange rate uncertainty (see Aysoy and Balaban, 1996; Aysoy et al., 1996; Akçay et al., 1997; Demiröz, 2001; Selçuk, 2004), the impact of exchange rate uncertainty on international trade (see Doğanlar, 2002; Vergil, 2002; Kasman and Kasman, 2005), the relation between exchange rate (or exchange rate uncertainty) and macroeconomic variables (see Levy, 1983; Salman and Salih, 1999; Kale, 2001; Berument and Günday, 2003; Berument and Paşaoğulları, 2003; Berument and Dinçer, 2004), the exchange rate determination (see Dülger and Cin, 2002) and exchange market pressure and foreign reserves (see Parlaktuna, 2005). The relationship between real exchange rate and foreign exchange reserves, however, has not been investigated for Turkey. Therefore, the main objective of this paper is to fill the gap by undertaking a study on the relationship between exchange rates (real and nominal) and foreign exchange reserves for the period 1982–2005 for Turkey, using cointegration and Granger causality testing.

The paper is structured as follows: Section 2 discusses econometric methodology. The data set and empirical results presented in Section 3. Finally, Section 4 contains concluding remarks.

2. Methodology

The aim of this paper is achieved in three steps. In the first step, detailed unit root tests are performed to establish the order of integration of the data series. The main idea of this exercise is to check whether, in the presence of a structural break in the data, the series are integrated of order one. Zivot and Andrews (1992) unit root test is used. In the second step, we test whether there is a long-run relationship between exchange rates and foreign exchange reserves. To accomplish this, the Gregory and Hansen (1996) cointegration test that takes account of the possibility of instability in long-run relations is used. This methodology allows a more general specification of long-run relations than conventional cointegration tests (Johansen, 1988 and Engle and Granger, 1987) do, while also allowing the estimation of the date of the structural change. In the third step, the Granger causality test is used to check the causation.

2.1. Unit roots

Prior to modeling the time series data, we determine the order of integration of the variables and ensure that it is equal for all series. The classical unit root test, namely Augmented Dickey and Fuller (Dickey and Fuller, 1979) (ADF) unit root test, is used to test for the nonstationarity of the series. Perron (1989) shows that in the presence of a structural break, conventional testing procedures (ADF and Phillips and Perron (1988) unit root tests) may erroneously fail to reject the null hypothesis that a series is integrated of higher order. To allow for the possibility of a structural break, we use the Zivot and Andrews (1992) (ZA hereafter) one break test. The main idea behind this exercise is to check whether, in the presence of a structural break in the data, the time series are integrated of order one or otherwise. Hence, the following testing equation is used.

\[ y_t = \mu + \beta t + \delta y_{t-1} + \gamma DU_t + \theta DT_t + \sum_{i=1}^{k} \eta_i \Delta y_{t-i} + e_t \]  

where

\[ DU_t = \begin{cases} 
1, & \text{if } t > TB \\
0, & \text{otherwise}
\end{cases} \]

and

\[ DT_t = \begin{cases} 
 t - TB, & \text{if } t > TB \\
0, & \text{otherwise}
\end{cases} \]

DU and DT are dummy variables that show a break in mean and slope, respectively. TB is the date of the endogenously determined break. The unit root hypothesis that \( \delta = 1 \) is considered, and the test statistic is the minimum \( t \) ratio over a grid search of all possible break dates in the sample.\(^8\)

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\(^8\) The Eqs. (1) and (2), (3) and (4) are estimated for the entire sample by trimming end points as suggested by Zivot and Andrews (1992). A trimming value of 0.10 is adopted, and the model with the minimum \( t \) statistic associated with the null hypothesis chosen. For each possible break date, \( k \) is estimated and an upper bound of \( k_{max} = 12 \) is used initially. If the last lag included is significant, \( k = 12 \) is set, otherwise \( k \) is reduced by 1 until the last lag becomes significant. If no lags are significant, the \( k = 0 \) is used.
2.2. Cointegration

For the series that are integrated of the same order, the next step is to perform cointegration tests. Bivariate cointegration tests are carried out using the Engle and Granger (1987) two-step estimation technique. In this method, a prerequisite for cointegration is that the nonstationary series have the same order of integration, that is, the number of times a variable has to be differenced to become stationary. The traditional cointegration tests have limitations especially when dealing with a long data span when the data generating process may be affected by major economic events such as financial and economic crises, shifts in industrial structure and productivity growth. These events could alter the equilibrium relationship. Several studies have shown the sensitivity of the outcome of the traditional tests to structural breaks. To examine further the robustness of our results to structural breaks, we apply the Gregory and Hansen (1996) cointegration tests that account for an endogenously-determined break. It is important to consider the possibility of a shift in the equilibrium relations that occurs at an unknown point in time. The results of Gregory and Hansen (1996) approach could be especially insightful when the null hypothesis of no cointegration is not rejected by the traditional tests. When it is rejected by the Gregory and Hansen (1996) test, we receive an important indication that a cointegration relation in fact exists, with the parameters of the cointegration relations being subject to change.

The Gregory and Hansen (1996) test assumes the null hypothesis of no cointegration against the alternative hypothesis of cointegration with one structural break. The timing of the structural change is not known a priori but is determined by the data. Gregory and Hansen (1996) suggest three alternative model specifications of structural change in the spirit of ZA (1992) depending on whether the change affects the intercept or the slope or both. The simplest case of structural change is a level shift in the cointegrating relationship. The second case is to include a time trend into the model in the spirit of ZA (1992) depending on whether the change affects the intercept or the slope or both. The simplest case of structural change is a level shift in the cointegrating relationship. The second case is to include a time trend into the shift model. The third case is the regime shift model, wherein it allows both the intercept term and slope term to change.

Model 1: Level shift (C)

\[ y_{1t} = \mu_1 + \mu_2 \varphi_{1t} + \alpha_1^T y_{2t} + \epsilon_t \quad t = 1, \ldots, n. \]  

(2)

where \( \mu_1 \) represents the intercepts before the shift and \( \mu_2 \) represents the change in the intercept at the time of the shift. \( y_{1t} \) and \( y_{2t} \) are of \( I(1) \) and \( y_{2t} \) is a variable or a set of variables.

Model 2: Level shift with trend (C/T)

\[ y_{1t} = \mu_1 + \mu_2 \varphi_{1t} + \beta t + \alpha_1^T y_{2t} + \epsilon_t \quad t = 1, \ldots, n. \]  

(3)

where \( t \) represents a time trend.

Model 3: Regime shift (C/S)

\[ y_{1t} = \mu_1 + \mu_2 \varphi_{1t} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t} \varphi_{1t} + \epsilon_t \quad t = 1, \ldots, n. \]  

(4)

where \( \mu_1 \) and \( \mu_2 \) are as in the first model, \( \alpha_1 \) represents the cointegrating slope coefficients before the regime shift, and \( \alpha_2 \) represents the change in the slope coefficients. The dummy variable that captures the structural change is defined by:

\[ \varphi_{1t} = \begin{cases} 0, & \text{if } t \leq [\eta \tau] \\ 1, & \text{if } t > [\eta \tau] \end{cases} \]

where the unknown parameter \( \tau \in (0,1) \) is a relative timing of the change point. The [] denotes integer part. Specifically, for each \( \tau \), estimation of the above models (depending upon the alternative hypothesis) by OLS, yields the residuals \( \hat{\epsilon}_{1t} \). The nonstationarity of the obtained residuals, expected under the null hypothesis, is verified by the ADF and PP tests. For each \( \tau \), \( k \) is chosen by the procedure in Perron and Vogelsang (1992). They start backwards from the maximum \( k_{max} \) equal to 12 until the last lag of the first difference included is significant at the 5% level using normal critical values.

The null hypothesis of Gregory and Hansen tests is that the residuals contain a unit root and hence there is no cointegration. The alternative hypothesis is that residuals do not contain a unit root and hence there is cointegration with a single unknown break. Similar to ZA (1992)’s approach, Gregory and Hansen (1996) tests choose the break point that gives the least support for the null hypothesis of a unit root in the residuals and hence no cointegration. Setting the test
statistics (denoted as $\text{ADF}^*_{(Z^*, Z^*)}$) to the smallest value of the ADF ($Z^*, Z_t^*$) statistics in the sequence, we select the value that constitutes the strongest evidence against the null hypothesis of no cointegration.\(^9\)

2.3. Causality

The third step is to examine causal relationship among the variables. In the presence of long-term equilibrium relations between foreign exchange reserves and exchange rates, an analysis of long and short-term interactions is in order. Hence, we use Granger causality tests to determine the direction of the short and long-run relationship. Granger (1986) and Engle and Granger (1987) provide a test of causality, which takes into account the information, provided by the cointegrated properties of variables. The model can be expressed as an error correction model (ECM) as follows:

$$
\Delta L_y = \alpha_1 + \text{lagged}(\Delta L_y, \Delta L_x) + \beta_1 E_{C_{t-1}} + \epsilon_1
$$  \hspace{1cm} (5)

$$
\Delta L_x = \alpha_2 + \text{lagged}(\Delta L_y, \Delta L_x) + \beta_2 E_{C_{t-1}} + \epsilon_2
$$  \hspace{1cm} (6)

where $y$ and $x$ denote foreign exchange reserves and real (nominal) exchange rates, respectively. $\beta_{EC_{t-1}}$ contains cointegrating terms, reflecting the long-run equilibrium relationship among variables. From Models 5 and 6, the short-run dynamics is provided by the lagged values of the difference terms. Since the Granger causality tests are very sensitive to the lag length selection, the lag lengths are determined Akaike’s (1969) minimum final prediction error (FPE) criterion.

3. Empirical results

3.1. Data

We use monthly real and nominal exchange rates and foreign exchange reserves for the period 1982:1–2005:11. The real exchange rate is defined as the trade-weighted real effective exchange rate. The data are taken from the electronic data delivery system of the Central Bank of the Republic of Turkey.\(^{10}\)

3.2. Unit root tests

The Engle and Granger (1987) and Gregory and Hansen (1996) cointegration tests require same degree of integration of nonstationary series. To ascertain the order of integration, we first used the augmented Dickey–Fuller (ADF) unit root test. The test results are reported in Table 1. It is evident from the table that the calculated ADF statistics are less than their critical values in all cases, suggesting that the variables are not level stationary.

\(^9\) We follow Gregory and Hansen (1996) to compute the ADF and PP test statistics for each breakpoint in the interval, 0.15\(T\) to 0.85\(T\) (where \(T\) is the number of observations). We will choose the breakpoint associated with the smallest value at which the structural break occurred.

\(^{10}\) To analyze the relationship between foreign exchange reserves and exchange rates, the number of variables in the model held to a minimum. The reason is that the more variables included in the model the higher the likelihood to obtain more than one cointegrating relationship in the long run. Interpretation of more than one cointegrating relationship can be confusing. Montiel (1999) discusses that factors such as productivity growth, government spending, changes in the international environment and changes in commercial policies are important determinants of the real exchange rate. Jin (2003), however, argues that real foreign exchange reserves can capture the impact of the variables identified by Montiel (1999). Following Jin (2003), Narayan and Smyth (2004) examine the relationship between the real exchange rate and foreign exchange reserves for China.
Perron (1989) states that, in the presence of a structural break, the conventional ADF test has low power to reject a unit root if the stationary alternative is true. To capture a possible structural break during the sample period, the Zivot and Andrews (1992) procedure is used, which treats the presence of a structural break in the series under investigation endogenously. The results of this procedure are reported in Table 2. The ZA test produces no additional evidence.

Table 2
Unit root test (with break): Zivot and Andrews (1992) unit root test

<table>
<thead>
<tr>
<th></th>
<th>Lex</th>
<th>Lreer</th>
<th>Lres</th>
</tr>
</thead>
<tbody>
<tr>
<td>TB</td>
<td>2001:01</td>
<td>1985:11</td>
<td>1994:05</td>
</tr>
<tr>
<td>δ</td>
<td>−0.021 (−2.626)</td>
<td>−0.094 (−4.856)</td>
<td>−0.108 (−4.337)</td>
</tr>
<tr>
<td>θ</td>
<td>−0.002 (−4.219)</td>
<td>0.000 (0.926)</td>
<td>−0.001 (−2.692)</td>
</tr>
<tr>
<td>γ</td>
<td>0.021 (1.601)</td>
<td>−0.026 (−2.276)</td>
<td>0.073 (3.115)</td>
</tr>
<tr>
<td>β</td>
<td>0.001 (2.796)</td>
<td>−0.000 (−0.451)</td>
<td>0.002 (3.662)</td>
</tr>
<tr>
<td>k</td>
<td>12</td>
<td>1</td>
<td>12</td>
</tr>
</tbody>
</table>

Note: Critical values at 1% and 5% significance level are −5.57 and −5.08 respectively (Zivot and Andrews, 1992). k is the lag length used in the test for each series. t statistics of the related coefficients are given in parenthesis.

Table 3
Engle–Granger cointegration test

<table>
<thead>
<tr>
<th>Models</th>
<th>ADF</th>
<th>k</th>
</tr>
</thead>
<tbody>
<tr>
<td>lreer=α+β lres+ε</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Model C</td>
<td>−3.701</td>
<td>2</td>
</tr>
<tr>
<td>Model C/T</td>
<td>−4.387</td>
<td></td>
</tr>
<tr>
<td>Model C/S</td>
<td>−6.086*</td>
<td></td>
</tr>
<tr>
<td>lres=α+β lreer+ε</td>
<td></td>
<td>5</td>
</tr>
<tr>
<td>Model C</td>
<td>−5.442* (1)</td>
<td></td>
</tr>
<tr>
<td>Model C/T</td>
<td>−5.378* (1)</td>
<td></td>
</tr>
<tr>
<td>Model C/S</td>
<td>−5.488* (1)</td>
<td></td>
</tr>
<tr>
<td>lex=α+β lres+ε</td>
<td></td>
<td>1</td>
</tr>
<tr>
<td>Model C</td>
<td>−3.308 (8)</td>
<td></td>
</tr>
<tr>
<td>Model C/T</td>
<td>−4.156 (9)</td>
<td></td>
</tr>
<tr>
<td>Model C/S</td>
<td>−3.299 (7)</td>
<td></td>
</tr>
<tr>
<td>lres=α+β lex+ε</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Model C</td>
<td>−6.256* (4)</td>
<td></td>
</tr>
<tr>
<td>Model C/T</td>
<td>−6.333* (4)</td>
<td></td>
</tr>
<tr>
<td>Model C/S</td>
<td>−6.326* (1)</td>
<td></td>
</tr>
</tbody>
</table>

Note: lreer, lres, and lex represent natural logarithm of reel effective exchange rate, foreign exchange reserves, and nominal exchange rate, respectively. The 1% critical value is −3.500 (MacKinnon, 1991).

Perron (1989) states that, in the presence of a structural break, the conventional ADF test has low power to reject a unit root if the stationary alternative is true. To capture a possible structural break during the sample period, the Zivot and Andrews (1992) procedure is used, which treats the presence of a structural break in the series under investigation endogenously. The results of this procedure are reported in Table 2. The ZA test produces no additional evidence.
against the unit root null hypothesis relative to the ADF test. In other words, the results from the ZA test confirm the results from the ADF test that all series are I(1). An a priori belief would be that the estimation procedure would find a break in the trend function consistent with the timing of the financial and economic crises in the last two decades. Foreign exchange reserves and nominal exchange rate show an estimated break consistent with this belief that shows a break in 1994:05 and 2001:01, respectively.

3.3. Cointegration tests

Having established that all variables are I(1), we now proceed to the cointegration analysis. We first test for cointegration without allowing for a structural break. Since the application of Gregory and Hansen (1996) test is a direct extension to the Engle and Granger (1987) test, we first decided to apply the Engle and Granger (1987) test in a bivariate setting. The test results are reported in Table 3. In all models, we clearly cannot reject the null of no cointegration at 1% level. Hence, exchange rates (real and nominal) and foreign exchange reserves are not cointegrated.

To allow for possible changes in the cointegrating vector over the estimation period, we apply the Gregory and Hansen (1996) procedure, which investigates the presence of cointegration under possible structural shift. Gregory and Hansen (1996) state that their testing procedure is of special value when conventional cointegration tests fail to reject the null hypothesis of no cointegration, but their test rejects it. This implies that structural change is present in the pattern of long-run co-movements between exchange rates and foreign exchange reserves. The test results are reported in Table 4. Panel A investigates the relationship between real effective exchange rate and foreign exchange reserves, and real effective exchange rate is the dependent variable. The first and second models reveal that cointegration is not present. The third model (C/S), which takes into consideration the simultaneous presence of both a mean break and a slope break (regime shift), however, reveals that cointegration is present. Hence, the Gregory–Hansen tests support a cointegrating relationship between the two variables that is not detected by the conventional cointegration test.

Panel B also examines the relationship between foreign exchange reserves and real effective exchange rate but now foreign exchange reserve is the dependent variable. The three models reveal that cointegration is present. Taken together, the Gregory–Hansen tests support a cointegrating relationship between the two variables that is not detected by the conventional cointegration tests, with a possible shift in mean and/or slope.

As for the relationship between foreign exchange reserves and nominal exchange rate, Panel C and D in Table 4 report the empirical results on the relationship between these variables. The test results indicate that the null of no cointegration cannot be rejected against the alternative of the presence of cointegration with unknown regime shifts when nominal exchange rate is used as the dependent variable. When foreign exchange reserves is used as the dependent variables, all three tests reveal that if the regime shifts are allowed for, there is evidence of a long-run relationship between foreign exchange reserves and nominal exchange rate.11

Table 5
Results of Granger causality tests: Real exchange rate and foreign exchange reserves

<table>
<thead>
<tr>
<th>Dependent variable</th>
<th>$F$-statistics</th>
<th>$t$-statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta \text{lreer}$</td>
<td>3.548*</td>
<td>-0.026***</td>
</tr>
<tr>
<td></td>
<td>(0.008)</td>
<td>(0.0929)</td>
</tr>
<tr>
<td>$\Delta \text{lres}$</td>
<td>0.054</td>
<td>0.003</td>
</tr>
<tr>
<td></td>
<td>(0.816)</td>
<td>(0.580)</td>
</tr>
</tbody>
</table>

*, **, *** indicate significance level at 1%, 5%, and 10% respectively. Numbers in parenthesis denote p values. The lag length for the models is determined as 4 and 1 respectively according to AIC.

The Gregory and Hansen (1996) cointegration test produces different break dates than those from the Zivot and Andrews (1992) unit root test though the majority of the breaks are either 2001, 1995, 1997 or 1986. The different break dates are due to the fact that unit root test searches for a break in a series while the cointegration test searches for a break in the residual of two series.

11
3.4. Causality

On the basis of the Gregory and Hansen (1996) test for cointegration, the Granger causality test requires a vector error correction model (VECM). The presence of cointegration Granger causality requires the inclusion of an error correction term in the stationary model in order to capture the short-run deviations of series from their long-run equilibrium path (see Eqs. (5) and (6)). The results of the causality tests for real exchange rate and foreign exchange reserves are reported in Table 5. It can be seen that the lagged error correction term is significant at 10% level with a negative sign, suggesting that in the long run, foreign exchange reserves Granger-cause the real effective exchange rate. The significance of the lagged error correction term displaying the appropriate negative sign supports the cointegration findings and implies a valid equilibrium relationship between the variables in the cointegrating equations. The error correction coefficient is very small (−0.026), suggesting that the adjustment of real exchange rate to changes in the foreign exchange reserves is slow. The error-correction term with foreign exchange reserves as the dependent variable, however, is insignificant. Hence, there is long-term one-way causality between two variables, the direction being from foreign exchange reserves to real effective exchange rate. The short-run dynamics can be seen by looking at the coefficients of lagged differenced terms. The coefficients of foreign exchange reserves are significant, suggesting that growth of foreign exchange reserves in the short run has significant influence on real effective exchange rate. The coefficients of real effective exchange rates, however, are insignificant.

Table 6 reports results of the causality analysis of nominal exchange rate and foreign exchange reserves. On the basis of Gregory and Hansen (1996) cointegration test, the Granger causality test requires a VAR model when nominal exchange rate is the dependent variable and a vector error correction model (VECM) when foreign exchange reserves is the dependent variable. The error-correction term with foreign exchange reserves as the dependent variable is highly significant. Hence, in the long-run the nominal exchange rate Granger causes foreign exchange reserves with causality running through the error correction term. Again, the error correction coefficient is very small (−0.05), suggesting that adjustment is slow. As for the short-run dynamics, the results indicate that changes in foreign exchange reserves are significant in influencing nominal exchange rate in the short run.

4. Conclusion

The objective of this paper was to investigate the long-run relationship between exchange rates (real and nominal) and foreign exchange reserves for Turkey, using conventional, ADF and Gregory and Hansen (1996) cointegration tests. The empirical results from conventional cointegration test suggest that there is no long-run relationship between exchange rates and foreign exchange reserves. Gregory and Hansen (1996) cointegration test that consider the presence of possible structural breaks, however, provided evidence of significant long-run relationship between foreign exchange reserves and real exchange rate; foreign exchange reserves and nominal exchange rate when the former is the dependent variable. Granger causality test was also used to check the causation. The results indicate that there is a unidirectional Granger causality running from foreign exchange reserves to real exchange rate in either short run or long run. Foreign exchange reserves Granger causes nominal exchange rate in the short run while a reverse causation occurs in the long run.

Theoretically, reserve holdings should be relatively low in a freely floating exchange rate regime. The financial crises of 1994 and 2001 that hit the Turkish economy heavily, motivated the monetary authority to hold more reserves. Therefore, foreign exchange reserves have become an important tool to decrease the vulnerability effects of exchange.
rate uncertainty and the likelihood of future financial crises. Hence, the ongoing reserve accumulation of the Central Bank of the Republic of Turkey shows an upward trend to withstand sudden reversals of capital inflows and financial panic in such a highly dollarized financial system and to restrain the overvaluation of national currency in recent years. Our results show that foreign exchange reserves have significant effect on reducing the volatility of real and nominal exchange rates in the short run, which is very important for market players. Therefore, news and expectations on the reserve accumulation do affect the real and nominal exchange rates. Surprisingly, the level of nominal exchange rate is also a determinant of foreign exchange reserve accumulation in the long run.

References


12 At the end of 2005, the share of foreign exchange deposits in total deposits is 33.7%.


