Testing for Asymmetric Cointegration Relationship between Banking Sector Development and Trade Openness: Evidence from Jordan

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ABSTRACT

The paper is an attempt to test the long-run asymmetric equilibrium relationships between financial development indicators, especially in banking sector, and trade openness. Using Jordanian data, over the period 1967 to 2011, the study has found evidence of asymmetry in the relationship between financial assets ratio and trade openness. However, there is no evidence of presence of asymmetric pattern in the relationship between trade openness and the credit to private sector ratio, total banks deposits ratio, and the financial depth indicator. We have failed to reject the null hypothesis of no cointegration with threshold at a conventional level of significance among two banking sector development, namely, credit facilitated to the private sector and financial development measured by M2 divided by GDP. From TAR estimation, it can be concluded that financial assets ratio responds asymmetrically to the change in the level of trade liberalisation in Jordan, where the speed of adjustment to the new equilibrium is faster when the residual in the past period is positive than when is negative.

Keywords: Asymmetric Cointegration, Financial Development, Trade Openness, Jordan.

INTRODUCTION

Over the last decade a substantial literature has been developed considers alternative specifications of investigating the long-run relationships among macroeconomic and financial variables. Most of these studies assume that this potential relationship may be represented as a symmetric linear combination of non-stationary stochastic regressors. Only a very few papers have paid attention to the case of nonlinear dynamics in the variables, where the series under consideration might be linked in an asymmetric relationship. Hence, asymmetric behaviour in economic variables has attracted attention in time series analysis through employing specific tests to investigate whether series adjust asymmetrically towards its long-run equilibrium value.

A number of studies consider the relationship between trade openness and financial development using a variety of econometric techniques, Including cross section and time series analysis and using samples of groups of countries. Some of these studies employed causality analysis on trade-finance nexus. It is clearly found in the literature that a more efficient financial system provides better financial services, this could be resulting from various variables including trade openness, and this enables an economy to increase its gross domestic product growth rate. Considerable body of the previous literature suggests a strong and positive link between the level of trade openness in a country and its financial development indicators. It has been argued that trade openness might stimulate external finance from financial system, and hence encouraging financial sector to facilitate credit provided to the private sector overcoming the problem of liquidity constraints. The arguments then show the vital role played by trade liberalisation in fostering financial development level for countries with unrestricted trade policies.

This paper makes one main contribution by examining the cointegrating relationship between financial development, especially in banking sector, and trade liberalisation using asymmetric cointegration test adopted by Enders and Siklos (2001). The remainder of this paper proceeds as follows. Section 2 summarises the main literature. Section 3 introduces an overview of the performance of both banking sector and trade openness indicators. The background of the main methodology to be used is discussed in Section 4, and Section 5 shows the
empirical results. Section 6 concludes.

**Review of Literature**

There has been considerable debate over the impact of trade liberalisation on financial development indicators, especially in banking sector. Rajan and Zingales (2003) provides clear evidence on the vital role of trade openness in enhancing financial development. They conclude that trade openness without financial openness is unlikely to deliver financial development. This study is one of the first works to discuss the direct impact of trade openness on financial development indicators, where according to this paper, cross border capital flows alone are unlikely to push for financial development, see (Rajan and Zingales 2003, p.22). Their analysis, therefore, suggests that the simultaneous opening of both trade and capital accounts holds the key to successful financial development. They simply highlight the necessity of a simultaneous Current account and Capital account openness. In addition, the same results are found in the study of Baltagi et al. (2009).

The relationship between trade openness and financial development has received relatively considerable attention in the theoretical and empirical literature. However, the debate about the possible links between financial development and international trade seems to become more crucial for many researchers, some of them concentrate on the effect of trade openness on financial development through external finance, they focused on the ability of the financial sector to channel savings to the private sector and therefore help overcome liquidity constraints. This will play a vital role in enabling the economy to specialize and exploit economies of scale. Economies with a better developed financial system and a higher level of external finance should therefore have a comparative advantage in sectors that exhibit high scale economies, see (Beck, 2002). Countries specialising in financially dependent goods would have a high demand for external finance and thus a high level of financial intermediation. However, financial system would be less developed in countries that specialise in goods less relying on external finance. Others argue that trade openness might affect the level of financial development in a country through affecting price elasticity. This will influence the level of uncertainty and increase income volatility. Financial development can be therefore fostered through increasing the demand for insurance, which is known as the consumption insurance view, see (Do and Levchenko, 2007).

It is also found that trade liberalisation affect financial development positively in the long-run. Using a panel data pooled from 88 countries covering the period 1960-2005, Kim et al. (2010) argue that a positive long-run relationship between trade openness and financial development coexists with a negative short-run relationship. It is clearly found that trade openness is detrimental to financial development in the short-run, it ultimately enhances financial development in the long-run. In contrast, when splitting the data into different income or inflation groups, the finding then can be found only in relatively low-income countries or high-inflation economies.

Trade and finance have been also connected in the literature based on the demand-driven framework, so that economies with higher export shares in sectors with scale economies have better developed financial systems. There might finally be a third factor; distortionary government policies might impede both the development of the financial sector and keep the economy relatively closed. While the theoretical model explores this one channel through which financial development affects the structure of trade, the empirical part controls for possible reverse causality and simultaneity bias, see (Beck, 2002). In addition, Svaleryd and Vlachos (2002) investigate this link within the demand side. They argue that trade and finance can be linked through the role of risk diversification, where trade openness is usually associated with greater risks, such as increased exposure to external demand shocks for foreign competition, it will create new demands for external finance. Firms will need credit in order to overcome short-run cashflow problems and adverse shocks.

Most of the previous work investigating the cointegration relationship between trade openness and banking sector development have focused on the use of the Engle-Granger (1987) framework, assuming that the adjustment mechanism of the ECT is symmetric, which means that the adjustment coefficients to the equilibrium level are the same for both positive and negative values of the residual obtained from the long-run relationship. This indicates that the speed of adjustment of the banking sector indicators, used in our analysis, is the same no matter if the shocks to trade openness indicator are positive or negative. The issue is that both the E-G and Johansen tests might be invalid if the adjustments to equilibrium appeared to be asymmetric. In this paper, we
specifically examine the potential cointegration relationship among banking sector development and trade liberalisation on the basis of existence of asymmetric cointegration.

If, however, there is a potential asymmetric pattern in the residual obtained from the long-run regression, one may use the tests by Enders and Granger (1998) and Enders and Siklos (2001) who develop an alternative specification for testing the stationarity of the residual, based on threshold and momentum threshold autoregressive models. They expand the cointegration test of E-G to incorporate an asymmetric error correction term. Therefore, we extend our analysis by employing the Enders and Siklos (2001) test to examine the presence of asymmetric cointegration in our data and extend the standard Engle-Granger Cointegration test used in previous literature.

An Overview of the Performance of Banking Sector Development and Trade Openness Indicators in Jordan

Jordan is one of the countries that are highly dependent on international trade, therefore, international trade sector in Jordan has been at the forefront of the wave of reforms that has extremely developed in during the past decades. As results of that, both exports and imports have experienced huge increases over time. Figure 1 presents the volume of exports and imports over the period 1964-2011 in Jordan. The volume of both exports and imports are expressed using the trade openness indicator, which is calculated by the sum of imports and exports divided by GDP. Although total imports and total exports have experienced a rapid growth over the last decades, this has had a small effect on the gap between imports and exports, which represents the trade deficit. Deficits in trade balance remain huge, due to the increase in the overall price level of the imported goods, more particularly crude oil prices.

It is also apparent that trade openness indicator has been on an upward trend over the last decades. It increased from about 35% in 1970 to about 80% in 1980. In the beginning of the 1990s, after experiencing the worst economic shock in the Jordanian economic history, trade openness indicator started to fluctuate at around 90% and then trade liberalization policies implemented by the Jordanian government through various trade agreements with foreign world started to affect this indicator and therefore trade openness indicator reached a very high rate at about 125% in 2005.

Regarding the banking sector development in Jordan, we produce an overview of the development of the banking sector in Jordan through looking at the trend of main banking sector indicator presented by total deposits held by the licensed banks in the Jordanian banking sector. Figure 2 initially represents the trend in the recorded total deposits held with the licensed banks, over the period 1964-2011, in million JD. This indicator reflects the ability of banks to create loans through the financial system.

The statistics provided by the central bank of Jordan (CBJ) show that commercial bank deposits have
experienced a remarkable increase over the period 1964-2011, and commercial banks have expanded their financial services and attracted more deposits, especially after the 1990s. It is also worth mentioning here that the statistics published by the CBJ indicate that the deposits in local currency has increased in recent years, where the ratio of deposits held in foreign currencies felt from 36.2% in 2005 to 26.3% in 2011 of total deposits held in the banking system. According to the CBJ report in 2011, this is due to the improvement of the attractiveness of the Jordanian Dinar as a saving base.

In order to summarize the development of the banking sector in the Jordanian economy over the period 1964-2011, we also show the main development in the number of bank branches in Jordan. Figure 3 shows that the number of commercial banks branches has been increasing steadily over the whole period. At the end of 2011, there were about 24 banks, of which three were Islamic banks and nine were branches of foreign banks.
These banks carry out their operations through a network of 695 branches and 71 representative offices all over the country (1).

It clearly appears that the Jordanian banking sector has experienced significant improvement through expanding the role of commercial banks in the economy, and that can be noticed from the constant increase in the latest indicators of banking sector development such as Credit to the private sector (CPY), Financial assets ratio (FA), and Financial depth (M2Y) (2), that are extensively investigated in the literature. The performance of the banking sector in Jordan is relatively better than many other developing countries. According to Creane et al. (2003) about financial development in the MENA region, Jordan has strengthened banking supervision and regulations, since very modern procedures have been established to collect prudential information on a regular basis as well as audit commercial banks operating in the country. To sum up, Jordanian banks, although relatively small, provide sophisticated financial services and operations in terms of technology, financial tools, and products in both retail and corporate banking.

Empirical Methodology

In this study, we extend the cointegration analysis on the linkages between trade openness and financial development used in the previous empirical works by employing the Enders and Siklos (2001) approach to test whether banking development indicators respond asymmetrically to trade openness shocks. We examine whether rises in trade openness indicator retard banking sector development by more or less than a fall in trade openness. In order to distinguish positive and negative effects of the error obtained from the cointegration regression, we use the asymmetric cointegration framework proposed by Enders and Siklos 2001. Asymmetric cointegration comes from the analysis of multivariate combinations arising from the decomposition of the series into positive and negative values of its cumulative sums, see Lardic and Mignon (2008, p. 484).

According to Cook (2006), testing for a potential long-run relationships among the main banking sector indicators and trade openness along with other explanatory variables (3), using the following equation for each banking sector development indicator.

\[
LFD_{it} = \beta_0 + \beta_1 LTO_t + \beta_{12} LXD_{it} + \varepsilon_{it}
\]  
(1)

Where \(LFD\) represents the financial development indicator \(i\) at time \(t\), that is, \(LCPY\): credit to private sector divided by GDP, \(LTDY\): total bank deposits divided by GDP, \(LFA\): financial assets ratio, \(LM2Y\): financial depth ratio expressed by M2 divided by GDP.

\(LTO\): Trade openness Indicator which is measured by (Exports + Imports) divided by GDP.

\(LX\): represents other explanatory variables used in the literature that might affect financial development, and may include GDP per capita (\(Y\)), inflation rate (P), and interest rate expressed by the re-discount rate (R). All variables are expressed in logarithms (4).

\(\varepsilon\): is the error term.

Testing for standard cointegration using the Engle-Granger two-step procedure (the E-G) requires that residual from equation (1) is obtained and then tested for a unit root, and then the ECM for the impact of trade openness on several banking development indicators can be constructed as follows:

\[
\Delta LFD_{it} = \alpha_0 + \sum_{j=1}^{p} \alpha_{1j} \Delta LFD_{it-j} + \sum_{j=1}^{p} \alpha_{2j} \Delta LTO_{t-j} + \sum_{j=1}^{p} \alpha_{12j} \Delta LX_{i,t-j} + \alpha_{14} \varepsilon_{it-1} + u_{it}
\]

where \(LFD_{it}\): financial development indicator \(i\) at time \(t\).


(2) \(CPY\): Credit to the private sector divided by GDP, \(FA\): The ratio of commercial banks assets to the total assets held by the commercial banks and the central bank, \(M2Y\): Broad money (M2) divided by GDP.

(3) Unit root tests, descriptive statistics, and correlation matrix results are presented in Tables 1A-3A, respectively, in the appendix of this paper.

Following Enders and Siklos (2001) and Enders and Granger (1998), the standard E-G cointegration test can be expanded to test for the presence of asymmetric cointegration. Given a number of series as it is shown in Equation 1, in the first step, a cointegration regression is estimated using OLS to obtain the long-run equilibrium relationship between each banking sector development indicator ($LFD_k$), trade openness ($LTO$), and other variables ($LX_k$), and then we look at the potential cointegration relationship through examining the stationarity properties of the residuals ($\varepsilon_{i,t}$).

Enders and Siklos (2001, p.167) employ the following equation to test for the stationarity of $\varepsilon_{i,t}$,

$$\Delta \hat{\varepsilon}_{i,t} = I_t \rho_1 \hat{\varepsilon}_{i,t-1} + (1-I_t) \rho_2 \hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{k} y_j \varepsilon_{i,j-1} + \zeta_{i,t}$$

(3)

Where $\rho_1$, $\rho_2$, and $y_k$ are coefficients to be estimated; $\zeta_{i,t}$ is a white noise disturbance; $k$ is the lag length; $I_t$ is the Heaviside indicator function, and Enders and Siklos (2001) consider two specifications for this function based on the level ($\hat{\varepsilon}_{i,t}$) and change in the residual ($\Delta \hat{\varepsilon}_{i,t}$) such that:

$$I_t = \begin{cases} 1 & \text{if } \hat{\varepsilon}_{i,t-1} \geq 0 \\ 0 & \text{if } \hat{\varepsilon}_{i,t-1} < 0 \end{cases}$$

(4)

$$I_t = \begin{cases} 1 & \text{if } \Delta \hat{\varepsilon}_{i,t-1} \geq 0 \\ 0 & \text{if } \Delta \hat{\varepsilon}_{i,t-1} < 0 \end{cases}$$

(5)

In Equation 3, the least squares estimates of $\rho_1$ and $\rho_2$ is based on the threshold autoregressive methods proposed by Tong (1983) and have an asymptotic multivariate normal distribution. The system of Equations 3 and 4 is known as the threshold autoregressive (TAR), and the combination of Equations 3 and 5 is known as the momentum threshold autoregressive (M-TAR). These two models are constructed using the original cointegration regression in equation 1.

In Equations 4 and 5, the value of threshold ($\tau$) is set equal to 0 based on assuming that ($\tau$) is unknown, especially in the TAR model, and needs to be estimated along with other coefficients in (3). In order to find the value of $\tau$, Enders and Siklos (2001) suggest to use of a grid search procedure to derive a consistent estimate of the threshold value based on Chan’s (1993) methodology of searching over possible threshold values to obtain the value which minimises the residual sum of squares.

Starting from the TAR model, the residual series ($\hat{\varepsilon}$) is reordered in ascending order as ($\varepsilon_1 < \varepsilon_2 < ... < \varepsilon_T$), where the 70% of the central observations are lying within the range ($\varepsilon_{i}, i = 0.15T, ..., 0.85T$), after discarding the largest and smallest 15% of $\varepsilon$, this central 70% of the residuals sequence is then considered in turn as potential threshold values in the TAR Equations 3 and 4. These threshold values provide the minimum residual sum of squares and are assumed to provide a consistent estimate of $\tau$ resulting from the estimation of (3) and (4). Similarly, using the same approach for M-TAR model in (3) and (5) the central 70% of observations is lying in the sequence ($\Delta \varepsilon_{i} < \Delta \varepsilon_{i+1} < ... < \Delta \varepsilon_{T}$) and considered as potential threshold values for (5) and the selected threshold is that value yielding the lowest residual sum of squares and again considered as the appropriate consistent indicator function of (3) and (5).

The next step is to examine asymmetric cointegration relationship between the variables starting from the long-run regression in (1) for each banking sector development indicators and determining whether these indicators and trade openness are cointegrated in the TAR and M-TAR models. We test the null hypothesis of no cointegration against the alternative of cointegration with asymmetry, where the null of $\rho_1 = \rho_2 = 0$ is examined using an F-test of the joint hypothesis. However, the F-statistic has non-standard distribution and the critical values are tabulated by Enders and Siklos (2001) and called, $\Phi$ and $\Phi^*$, for the TAR and M-TAR models, respectively.

Alternatively, $t$-statistics obtained from the point estimates of both $\rho_1, \rho_2$ are considered as the second method to test the hypothesis of no cointegration. This test is called the t-max test and it is based on reporting the most negative $t$-value associated with either $\rho_1$ or $\rho_2$ in (3). The negative sign of $t$-values associated with both

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*(Enders and Siklos (2001) have conducted two Monte Carlo experiments to test the null of no cointegration against the alternative of cointegration with threshold adjustment. In the first test (TAR), the value of $\tau$ is set equal to 0, and in the second test (M-TAR), $\tau$ is unknown. For more details about the critical values see their paper pp. 169-171.)*
\( \rho_1, \rho_2 \) are considered as the necessary conditions for convergence. Obviously, if the t-value on either \( \rho_1 \) or \( \rho_2 \) is positive, we would not reject the null and the F-test loses its validity (as it is a two-sided test), see Harris and Sollis (2003, p.105).

In this study, we are testing for the presence of uncovered asymmetric cointegration. The E-G test, implemented in most of the previous empirical studies, is considered as a special case of the threshold cointegration test. The adjustment of the error term is asymmetric when variables are threshold cointegrated, where the adjustment is \( \rho_1 \) if the banking sector indicator lagged one period \( LFD_{t-1} \) is above its long-run equilibrium \( \left( \beta_0 + \beta_1 TIO_t + \beta_2 L\theta_{2t} \right) \) in (1), and \( \rho_2 \) if \( LFD_{t-1} \) is below. If \( \rho_1 = \rho_2 \) in (3), adjustment is then symmetric (equal) and the E-G is a special case of (3) and (3), see Shen et al. (2007, p.1437) and Enders and Siklos (2001, p.167).

**Empirical Results of Asymmetric Cointegration Test**

To examine the potential asymmetric cointegration relationship between banking sector development indicators, trade openness, and other variables, we first estimated both the E-G and Johansen cointegration tests\(^6\). We extend our research taking into account the possibility of having an asymmetric cointegration relationship among variables, employing Enders and Siklos’s (2001) TAR model. The threshold value is set to 0.

If we find evidence of asymmetric cointegration, we then continue our analysis by estimating the ECM including two adjustment coefficients to allow for asymmetric adjustment. The results of our estimation of threshold cointegration relationships between various banking sector development indicators and trade openness in each model are illustrated in Table 1. As we mentioned previously, Equation 3 represents the TAR model, along with Equation 4, of the disequilibrium errors for each model. Both F-statistic and t-Max are reported for the four error equations. We use the critical values tabulated by Wane et al. (2004) for the test as they have extended the critical values of \( \Phi \) for the null hypothesis of no cointegration with asymmetric adjustment for up to five variables with different sample sizes, see Wane et al. (2004, p. 2).

Table 1 reports the results from threshold cointegration analysis for all models. The test is conducted by using both F-statistic and t-Max. It is found that there is no asymmetric cointegration between trade openness and credit to private sector as shown in model (1) since the \( \Phi \) statistic for the null hypothesis \( \rho_1 = \rho_2 = 0 \) is 7.67 which in less than the 1% and 5% critical values and we therefore cannot reject the null hypothesis of no asymmetric cointegration at 1% and 5% level of significance, whereas the largest of t-statistics equals -3.764 rejects the null hypothesis at all significance levels.

In addition, we have failed to detect asymmetric cointegration between trade openness indicator and total bank deposits ratio as well as financial depth as reported in model (2) and (4). F-statistics and t-Max calculated statistics are substantially smaller than their corresponding critical values indicating that the null hypothesis cannot be rejected at any level of significance which implies that the cointegration with asymmetric adjustment does not exist among these series. This also confirms the finding using both the E-G and Johansen cointegration tests, (not reported here). Therefore, neither symmetric nor asymmetric adjustments are found between trade liberalisation and banking sector development indicator using TDY and M2Y in the Jordanian case.

For model (4), which contains the other financial depth indicators using the ratio of broad money (M2) to narrow money (M1), there is no asymmetric cointegration relationship among trade openness and financial depth indicator (M2Y) as well as other variables. The calculated F-statistic (10.42) is smaller than the critical values at both 1% and 5% indicating that we cannot reject the null hypothesis of no cointegration with threshold.

\(^{6}\) The full results of these estimates are not reported here due to the limitation of space.
Table 1. Asymmetric Cointegration Test (TAR) Model

<table>
<thead>
<tr>
<th>Models and Variables</th>
<th>Estimates</th>
</tr>
</thead>
<tbody>
<tr>
<td>Model 1 (LCPY, LTO, LY, R)</td>
<td>$\hat{\epsilon}<em>{1,t} = -0.4871\hat{\epsilon}</em>{1,t-1} - 0.5141(1 - l_t)\hat{\epsilon}<em>{1,t-1} + \zeta</em>{1,t}$</td>
</tr>
<tr>
<td></td>
<td>$(-3.764)*** (-2.346)$</td>
</tr>
<tr>
<td>Model 2 (LTDY, LTO, LY, R)</td>
<td>$\hat{\epsilon}<em>{2,t} = -0.5371\hat{\epsilon}</em>{2,t-1} - 0.1771(1 - l_t)\hat{\epsilon}<em>{2,t-1} + \zeta</em>{2,t}$</td>
</tr>
<tr>
<td></td>
<td>$(-2.813) (-0.769)$</td>
</tr>
<tr>
<td>Model 3 (LFA, LTO, LY, RR)</td>
<td>$\hat{\epsilon}<em>{3,t} = -0.6341\hat{\epsilon}</em>{3,t-1} - 0.2791(1 - l_t)\hat{\epsilon}<em>{3,t-1} + \zeta</em>{3,t}$</td>
</tr>
<tr>
<td></td>
<td>$(-4.916) (-1.451)$</td>
</tr>
<tr>
<td>Model 4 (LFDMY, LTO, LY, R,)</td>
<td>$\hat{\epsilon}<em>{4,t} = -0.6171\hat{\epsilon}</em>{4,t-1} - 0.6441(1 - l_t)\hat{\epsilon}<em>{4,t-1} + \zeta</em>{4,t}$</td>
</tr>
<tr>
<td></td>
<td>$(-3.047)*** (-2.787)$</td>
</tr>
</tbody>
</table>

- $\Phi$ represents F-statistics for the null hypothesis of no cointegration with threshold $\rho_1 = \rho_2 = 0$. Critical values are obtained from Wane et al. (2004) for the case of 4 variables (Table 5), and they are as follows: 10.53, and 13.91 at 5% and 1% respectively.
- t-statistics are shown in parenthesis, and the critical values are obtained from Enders and Siklos (2001) (Table 2) as follows. -2.12 and -2.58 at 5% and 1%, respectively.
- ** and *** indicate significant at 5% and 1%, respectively.

Table A1. Results of Unit Root Tests

<table>
<thead>
<tr>
<th>Variables</th>
<th>PP</th>
<th>ADF</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Level</td>
<td>First Difference</td>
</tr>
<tr>
<td></td>
<td>C</td>
<td>C+T</td>
</tr>
<tr>
<td>LTO</td>
<td>-1.87(2)</td>
<td>-1.97(2)</td>
</tr>
<tr>
<td>LCPY</td>
<td>-1.90(2)</td>
<td>-1.14(1)</td>
</tr>
<tr>
<td>LTDY</td>
<td>-1.49(2)</td>
<td>-0.03(2)</td>
</tr>
<tr>
<td>LFA</td>
<td>-1.35(2)</td>
<td>-3.78(2)</td>
</tr>
<tr>
<td>LM2Y</td>
<td>-0.76(2)</td>
<td>-2.51(2)</td>
</tr>
<tr>
<td>LY</td>
<td>-0.15(0)</td>
<td>1.841(1)</td>
</tr>
<tr>
<td>R</td>
<td>-2.17(2)</td>
<td>-2.22(3)</td>
</tr>
<tr>
<td>P</td>
<td>-1.29(2)</td>
<td>-2.03(0)</td>
</tr>
</tbody>
</table>

- Critical values are -2.62, -2.93, and -3.59 at 10%, 5% and 1% respectively for test with constant.
- Critical values are -3.19, -3.52, and -4.18 at 10%, 5% and 1% respectively for test with constant and trend.
- *(**), and *** denote reject at 1%(5%), and 10%.
- C: including constant, C+T: including constant and trend.
- The numbers in brackets are lag lengths which are selected based on the SBC.
However, we find using the $\Phi$ statistic, asymmetric adjustment between trade openness and financial assets ratio since the F-statistic indicates that the null hypothesis is rejected at 1% level as it is shown in model (3). This verifies that there exists a long-run cointegration relationship among trade openness and financial assets ratio with underlying adjustment process being highly asymmetric. It is clearly noticed that there is a difference between the values of the asymmetric adjustments $\rho_1, \rho_2$ with (-0.634) and (-0.279) coefficient, respectively. In such a case, it can therefore be concluded that financial assets ratio responds asymmetrically to the changes in trade openness and other variables, where the adjustment speed to the new equilibrium at period $t$ is faster when $\hat{\epsilon}_{3t-1}$ is positive than when is negative.

Conclusion

This paper empirically investigates the long-run equilibrium relationships between banking sector development indicators and the level of trade openness in Jordan using cointegration tests assuming asymmetric adjustments. We have employed different models reflecting a variety of indicators of banking sector development. The research extends both the standard Engle-Granger and Johansen Cointegration tests used in the previous empirical work by testing for the presence of asymmetric cointegration among our data. The results obtained show that there is no evidence of presence of asymmetric pattern in the relationship between trade liberalisation and credit to private sector ratio, total banks deposits ratio, and the banking sector development indicators using monetary variables. We have failed to reject the null hypothesis of no cointegration with threshold at a conventional level of significance except for financial assets ratio. However, there is a strong evidence of asymmetry in the relationship between financial assets ratio and trade openness along with other explanatory variables. This results emphasise the fact that
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financial assets ratio responds asymmetrically to the change in trade openness indicator, where the adjustment speed to the new equilibrium is faster when the residual in the past period is positive than when is negative.

REFERENCES


الانفتاح المصرفي لتطوير القطاع:
دراسة حالة الأردن

ملخص
تحاول هذه الورقة اختبار مدى إمكانية وجود علاقة تكامل مشترك غير متماثل في الأجل الطويل بين مؤشرات التطور المالي، لا سيما مؤشرات القطاع المصرفي، ودرجة الانفتاح التجاري، من خلال دراسة حالة الأردن خلال الفترة 1967-2011. أظهرت نتائج اختبار Enders and Siklos (2001) للكملاس المشترك غير المتماثل أنه لا يوجد تكامل مشترك غير المتماثل بين مستوى الانفتاح التجاري وكل من الإثاث الممكن للمؤشرات الخاص، إجمالياً الوحدة البنكية، ومؤشر العمق المالي كنسبة من الناتج المحلي الإجمالي، حيث أظهرت نتائج التقدير القياسي لهذا الاختبار أن نمط العلاقة بين مؤشرات تطور القطاع المصرفي هذه ودرجة التحرر الاقتصادي غير متماثل في الأجل الطويل. في حين أظهرت النتائج أن هناك علاقة تكامل مشترك غير متماثل قوية ومعنوية بين درجة الانفتاح التجاري ونسبة الأصول المالية للبنوك التجارية في القطاع المصرفي الأردني. الأمر الذي يشير إلى أن نسبة الأصول المالية للبنوك التجارية تستجيب بشكل غير متماثل للتغير في درجة الانفتاح التجاري في الأردن حيث أن سرعة التعديل نحو الوضع التوازني الجديد تكون أسرع عندما تكون في بوابات الاندماج في الفترات السابقة.

الكلمات الدالة: التطور المصرفي، الانفتاح التجاري، الأردن.

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