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Economic Growth and Credit Balances: Vector Error-Correction Modeling Approach

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Abstract

This study investigates the short run and long-run dependence patterns among real GDP (Y_1), consumer price index (Y_2), and real credit balances (Y_3) using Jordanian data for the period 1969-2001. In this study, unit root is tested using Augmented Dickey-Fuller (1981) (ADF) test. The tests yield evidence showing that nonstationarity cannot be rejected for the levels of all variables. Since a unit root has been confirmed for the data series, cointegration tests are carried out using the method proposed by Johansen (1988). The cointegration relations along with the nonstationarity of the three data series used in the study, allow us to employ the vector error-correction (VEC) technique to examine the short-run and long-run relationships among these variables. The estimation results provide empirical evidence showing that an increase in real GDP will result in an increase in credit balances.

1. Introduction

For the last ten years, there have been voluminous studies on the so-called 'bank credit channels' of monetary transmission mechanism (Bernanke and Blinder, 1992; Meltzer, 1995; Panagopoulos and Spiliotis, 1998). This kind of studies has been partially motivated by the growing literature on asymmetric information theory in financial markets. From the viewpoint of asymmetric information theory, large fluctuations in aggregate economy can be generated from an apparently small shock.

The bank-lending channel is based on the view that banks play a special role in the financial system because they are almost well suited to solve asymmetric information problems in credit markets. The importance of this role arises from the fact that most of borrowers will not have access to the credit market unless they borrow from banks. As long as there is no perfect substitutability of retail bank deposits with other sources of funds, an expansionary monetary policy, which increases bank reserves and bank deposits, increases the quantity of loans available, which will cause investment spending to rise (Mishkin, 1998).

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Bernanke and *et al.* (1996) argue that the role of banks in the transmission of monetary policy relates to both their liability and assets. In a monetary contraction, for example, bank reserves decrease and, owing to reserve requirements, bank deposits fall. If the decrease in bank deposits is not offset by other funds, which are not subject to reserve requirements, or decrease in securities, this will result in a decrease in bank loans. If bank loans fall, the monetary policy reduces both investment and consequently economic activity. The existence and effectiveness of this channel depend on two factors: (1) bank loans and securities must be an imperfect substitute for some borrowers, or some borrowers are bank-dependent; (2) central bank must constrain the supply of bank loans.

An important implication of the credit view is that monetary policy will have a greater effect on expenditure by smaller firms, which are more dependent on bank loans, than it will on large firms, which can access the credit markets directly through stock and bond markets. This implication can also be generalized to less developing countries such Jordan, in which firms are more dependent on bank loans than firms in developed countries.

The purpose of this paper is to investigate the short run and long-run dependence patterns among real GDP (Y_1), consumer price index (Y_2), and real credit balances (Y_3) using Jordanian data for the period 1969-2001. As shown in figure 1, these variables (figures are in logarithmic form) turn out to show remarkable regularities about their long-run trends, and appear to contain stochastic trends. One important source of non-stationarity can be attributed to the economic growth that resulting from technological progress. This change definitely ensures secular trends in real GDP series.

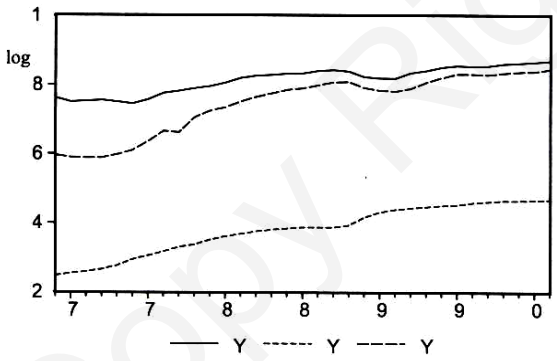

Figure 1. Real GDP (Y_1), consumer price index (Y_2), and real credit balances (Y_3)

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The five-year average growth rates of these variables are also reported in table 1. From these data we can derive three observations. First, the average rate of economic growth for the entire period exceeded 3.5 percent per year, while consumer price index and bank credit grew at 7.27 and 8.84 percents, respectively. Second, the economy shows a sharp drop in productive capacity as a result of the 1973 and 1990-1991 wars and the foreign exchange crisis at the end of 1980's. Third, the most remarkable aspect of the Jordanian economy is its rapid growth over the sub-periods 1976-1980 and 1991-1995, resulted from acceleration in aggregate demand.

Table 1

Average growth rates of real GDP (Y_1), consumer price index (Y_2), and real credit balances (Y_3)

Period	Y_1	Y_2	Y_3
1971-1975	1.57	10.97	10.30
1976-1980	10.35	11.69	23.44
1981-1985	5.71	5.46	11.74
1986-1990	-2.38	9.62	-0.69
1991-1995	7.64	4.29	10.29
1996-2000	2.15	2.76	1.42
1970-2001	3.68	7.27	8.84

Source: Original time series data come from International Financial Statistics Yearbook, Vol. LV, 2002.

2. Literature Review

In his extensive survey study, Levine (1997) refers to Goldsmith (1969) as the first cross-country study of growth and financial development. In this study, Goldsmith used the value of intermediary assets to GDP as an indicator of the provision of intermediary services. Looking at decade averages for 35 countries over 100 years, he found broad indications of a positive relationship between finance and growth. Goldsmith's work was econometrically unsophisticated and did not seem to spur much research

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interest at that time. Later on, more specific econometric works were conducted to identify the determinants of growth and the direction of causality.

Econometric studies conducted on the relationship between financial sector development and economic growth began to appear in the 1990s with King and Levine's (1993a, b) and Wachtel and Rousseau's (1995). Overall, these studies provide empirical evidence showing that the depth of financial sector development is likely to be associated with economic growth across countries. In the decade since those studies appeared, there has been a veritable explosion of empirical interest in the finance–growth relationship.

Stiglitz (1994) argues that due to poor accounting disclosures and credit rating agencies, asymmetric information seems to be prevalent in the financial markets of most developing countries. With this situation, most classes of borrowers find it difficult to issue securities. As a result of the predominant role of banks in developing countries and their ability to overcome the asymmetric information problem in credit markets, many borrowers are substantially bank-dependent.

In an extensive empirical study, Peek and *et al.* (1999) present evidence supporting the notion that confidential bank supervisory information could be used to improve macroeconomic forecasts. While this finding provided support for the hypothesis that the central bank should retain bank supervisory authority to facilitate the conduct of monetary policy, it failed to address an important underlying question: Why does the information on the health of the banking system improve macroeconomic forecasts? They argue that one possible explanation is that confidential supervisory information might be serving as a leading indicator of future macroeconomic activity. Alternatively, the financial health of the banking system might causally affect macroeconomic activity if the banking sector plays a central role in the transmission mechanism of monetary policy. Peek and *et al.* (2003) addressed the same issue. Their study confirms compelling new support for the hypothesis that the credit channel is operative in the U.S. economy, and that loan supply shocks have had a significant impact on real macroeconomic variables over the past two decades.

To identify the channel of monetary policy, empirical studies (Kashyap and Stein, 1995 and 1997; Dale and Haldane, 1994) are carried out using cross sectional data to determine whether there are distributional effects of monetary policy across lenders and borrowers, as predicted by the bank-lending channel argument. On the lenders' side, these studies yield findings that are compatible with the lending view, which simply suggests that monetary policy shock should constrain bank loan supply since banks cannot raise non-deposit funds to make up for a shortfall in their deposits. But this will depend on the ability of banks to be insulated from the shock. Furthermore, small banks which have relatively limited access to non-deposit

funds such as securities issues or foreign borrowing are found to be more affected by the monetary shock and tend to cut their loan supplies immediately following the shock. On the borrower side, small firms, which have limited access to external finance, are found to be very sensitive to monetary shock (Gertler and Gilchrist, 1994).

In a study conducted on the corporate structure of large firms listed in the Indonesian stock market, Ang *et al.* (1997) found that there was no evidence of asymmetric information problems in the firm-bank relationship. This result is fully compatible with the hypothesis that larger firms that have long operating histories and thus built up relationships and reputations with their banks and the credit market, asymmetric information diminishes. They argue that in countries, such as Japan and Korea, business groups are organically linked with banks. As a result, firms which are members of business groups tend to be less sensitive to cash flow than those unconnected with the business groups.

Agung (1998) examined the behavior of the bank balance sheet across bank categories using Indonesian data for the period 1983:01–1995:12. He found significant different responses across the bank-size classes to a change in monetary policy. In particular, a monetary contraction does not significantly influence lending by state banks, but it leads to a decline in lending by smaller banks. A similar pattern emerges when loans are disaggregated into investment and working capital loans. However, consumer loans of all bank categories drop substantially after the monetary tightening. Thus, the empirical results provide evidence of capital market imperfections for banks. The results indicate that monetary policy has been unable to restrain lending by the state banks because of their ability to liquidate their securities following a monetary tightening, and possibly also because of their access to non-deposit funds including foreign borrowing.

Despite the alarming econometric problems, a wide body of literature has firmly provided empirical evidence supporting the relationship between financial sector development and economic growth. For example, Rousseau and Wachtel (2000, 2001) examine the ratio of the broad money supply to GDP with panel data for large number of countries. Their results indicate that increasing that ratio by 10 percentage points will, particularly in countries without high inflation, increase the rate of growth by between 0.6 and 1.0 percentage points a year. To address the issue of causality more directly, they estimate VAR systems approach suggested by Arrellano and Bond (1991). They find evidence of significant causality from financial measures to real GDP and no evidence of feedback from GDP to the financial variables.

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3. Methodology

The recent development in time-series analysis and the empirical evidence have shown that most time series data are not stationary in their levels in the sense that the mean and variance tend to explode as time goes on (Burney, 2002). This implies that nonstationary time series, when subjected to exogenous shocks do not return to their long-run equilibrium path. Under these circumstances, most of the statistics calculated from the regression involving nonstationary time-series data as well as tests of significance are invalid (Hendry and Juselius, 2000; Chang, 2002). In more specific words, the regression models containing nonstationary variables show spurious relationships and yield inconsistent and less efficient OLS parameters.

Tests for nonstationarity of a time series (y_t) involve testing for the presence of unit root. In this study, unit root is tested using Augmented Dickey-Fuller (1981) (ADF) test. The test is the t- statistics on θ in the following regression:

$$\Delta y_t = \delta_0 + \delta_1 T + \theta y_{t-1} + \sum_j^n \varphi_j \Delta y_{t-j} + \eta_t \qquad (1)$$

where Δ is the first-difference operator, y_t is the series under consideration (small letters denote logarithmic values), η_t is a stationary random error, T is the time trend, δ_0, δ_1, θ, and φ_j are parameters to be estimated. The hypothesis of non-stationarity is rejected when θ is significantly negative. Here n must be selected large enough to ensure that η_t is a white noise. In this study, the Akaike (1974) information criterion (AIC) is used to determine the appropriate lag length n that will be enough to ensure the stationarity of the error term η_t. The AIC is defined as

$$AIC = N*Ln(ESS/N) + 2k \qquad (2)$$

where N is the sample size, ESS is the sum of squared error of the regression in equation 1, and k is the number of parameters, $k = n + 3$. The appropriate lag length selected by estimating equation 1 over a selected grid of values of n and finding that value of n at which AIC attains its minimum (Engle and Yoo, 1987).

Once a unit root has been confirmed for a data series, the question is whether there exists some long-run equilibrium relationship among the variables (y_{1t}, y_{2t}, and y_{3t}).While the theory of cointegration reveals a long-run equilibrium relationship among the dependent and independent variables, analysis of the short-run dynamics of the system is equally important. An important issue in econometrics has been the need to integrate short-run with long-run equilibrium.

Cointegration tests in this study are carried out using the method proposed by Johansen (1988). The Johansen method applies the maximum likelihood procedure to examine the presence of cointegrating vectors in

nonstationary time series. Following Hendry and Juselius (2000b), a three-dimensional (3×1) vector autoregressive model with Gaussian errors can be expressed by

$$y_t = \phi_1 y_{t-1} + \phi_2 y_{t-2} + \ldots + \phi_k y_{t-k} + \mu + \varepsilon_t \qquad (3)$$

$$t = 1,2, \ldots, T$$

where $y_t = (y_{1t}, y_{2t}, y_{3t})$, and $\varepsilon_t \sim$ i.i.d. N $(\mathbf{0},\Lambda)$. The covariance matrix of the error process, Λ, and the parameters ϕ_1, ϕ_2, ϕ_k, and μ are to be estimated. By taking first differencing on the vector level, the model in error correction form is

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \Gamma_2 \Delta y_{t-2} \ldots\ldots + \Gamma_{k-1} \Delta y_{t-k+1} - \Pi y_{t-1} + \mu + \varepsilon_t \qquad (4)$$

where $\Gamma_i = -(I - \phi_1 - \phi_2 - \ldots\ldots \phi_i)$ are short-run parameter matrices, $\Pi = (I - \phi_1 - \phi_2 - \ldots\ldots - \phi_k)$, sub-index k is the lag-length. The matrix Π conveys information about the long-run relationship among y_{1t}, y_{2t}, and y_{3t}. Testing for cointegration involves testing for the rank of Π matrix r by examining whether the eigenvalues of Π are significantly different from zero. Three possible conditions exist: (a) the Π matrix has full column rank, implying that y_t was stationary in level to begin with; (b) the Π matrix has zero rank, in which case the system is a traditional first-differenced VAR; and (c) the Π matrix has rank r such that $0 < r < 3$, implying that there exist r linear combinations of Y_t that are cointegrated. If the condition (c) prevails, then Π matrix can be decomposed into two 3×r matrices, α and β, such that $\alpha\beta' = \Pi$. The loading matrix α represents the error correction parameters, which can be interpreted as speed of adjustment, while the vectors of β represent the r linear cointegrating relationships such that $\beta' Y_t$ is stationary.

Following Johansen (1988) and Johansen and Juselius (1990), the likelihood ratio will be used for testing the number of cointegrating vectors (or the rank of Π). The likelihood ratio statistic for the trace test is

$$LHR = -T \sum_{i=r+1}^{p-2} \ln (1 - \hat{g}_i) \qquad (5)$$

where $\hat{g}_{r+1}, \ldots, \hat{g}_p$ are the estimated p-r smallest eigenvalues. The null hypothesis to be tested is that there are at most r cointegrating vectors. That is, the number of cointegrating vectors is less than or equal to r, where r is 0, 1or 2. In each case, the null hypothesis is tested against the general alternative of r + 1 cointegrating vectors. Thus, the null hypothesis r = 0 is treated against the alternative that r =1, r =1 against the alternative that r =2, and so forth.

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Since cointegration tests are very sensitive to the choice of lag length used in carrying out such tests, the Schwarz (1978) criterion (SC) will be used to select to select the optimal number of lags required in estimating the cointegration test. The SC is defined as follows:

$$SC = \text{Ln}\ \Omega_n^2 + n\ \text{Ln}\ (N)/N \qquad (6)$$

where Ω_n^2 is the maximum likelihood estimator of the residual variance obtained from a model with lag length n, that is $\Omega_n^2 = SSE_n/N$, N is the sample size, and n is the number of lags selected to numerically minimize SC in equation (6).

4. Empirical Findings

This section begins by reporting the results of nonstationary test for real GDP (y_1), price level (y_2), and real credit balances (y_3) using ADF test. The regression models are estimated using 1, 2, and 3, lagged differences. As shown in table 2, nonstationarity cannot be rejected for the levels of all variables at the 5% significance levels. In contrast, when the data are differenced, nonstationarity can be rejected for real GDP at 5%, while for consumer price index and real credit balances at 10% significance level. This finding indicates that all the data series are integrated of order one (or I(1)). Table 2 also reports the minimum AIC(n) which provides the appropriate order of the autoregressive process, n, in equation 1.

Table 2

Augmented Dickey-Fuller unit root tests

Variable	Order(n)	Level: Intercept and Trend		First Difference: Intercept	
		ADF	AIC(n)	ADF	AIC(n)
Y_1	1	-2.25	-5.49	-3.40**	-5.38
	2	-1.98	-5.40	-3.03**	-5.28
	3	-1.87	-5.29	-2.66*	-5.31
Y_2	1	-1.56	-6.18	-2.71*	-6.05
	2	-1.45	-6.18	-2.42	-5.95
	3	-1.40	-6.12	-2.06	-5.85
Y_3	1	-1.25	-4.32	-2.66*	-4.26
	2	-1.75	-4.35	-2.15	-4.16
	3	-2.31	-4.38	-2.21	-4.01

Note: (*)** denotes significance at the 10% and 5% level, respectively.

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Since a unit root has been confirmed for the data series, we now turn to examine for the existence of some long-run steady-state relationships among the variables using Johansen's cointegration test. Since cointegration tests are quite sensitive to the number of lags used in implementing the test, two vector error-correction (VEC) models with r =1 and r =2 are first fitted to the data to select an appropriate lag structure (see appendix A). The AIC and SC statistics computed on the basis of these models suggest a one lag length. On the basis of this lag structure, the cointegration test of Johansen is then carried out. Four vectors of dummy variables are introduced in the model as exogenous variables to capture the mean-shift in the variables during the sample period studied. The first dummy variable, D1 (= 1 for 1970, 1973-1974, 1988-1991, 0 otherwise), to capture supply shocks of the 1973, 1990-1991 wars, and the deterioration in the foreign exchange rate of the Jordanian dinar in 1988-1989. The second dummy variable, D2 (= 1 for 1975-1981 and 1992, 0 otherwise), is assumed to capture rapid increases in aggregate demand during these years. The third dummy, D3 (= 1 for 1977, 0 otherwise) to capture the sharp deterioration in credit market in 1977. The fourth dummy variable D4 (=1 for 1979, 0 otherwise) to capture the dramatic increase in credit balances in 1979. The test is a likelihood ratio test, namely the trace test derived in Johansen (1988, 1991).

To examine the robustness of the results, Johansen's test was estimated under three different assumptions. The first assumption assumes no deterministic trend in data (no intercept or trend in CE or test VAR). The second assumption assumes no deterministic trend in data (intercept (no trend) in CE-no intercept in VAR). The third assumption test allows for linear deterministic trend in data (intercept (no trend) in CE- and test VAR) (see table 3). These tests confirm the existence of one to two cointegrating equations at the 5% significance level.

This cointegration relations along with the nonstationarity of the three data series used in the study, allow us to employ the vector error-correction (VEC) technique to examine the short-run and long-run relationships among real GDP, consumer price index, and banks lending. To avoid misspecification error, two error-correcting terms are included in the VEC model. Table 4 reports the estimation of VEC model (6), i.e. estimation results of long-run structure (EC), the short-run structure, the deterministic structure.

We first turn to the long-run equilibrium relationships between the variables and the error-correcting terms. The estimation results indicate that real GDP is negatively and positively equilibrium-correcting to EC_{1t-1} and EC_{2t-1}, respectively. This error correction mechanism supports the existence of the co-movement between real GDP, consumer price

Table 3

Johansen's Test for Cointegrating Ranks

Test assumption: assumes no deterministic trend in data (no intercept or trend in CE or test VAR)

		Null hypothesis		
		r = 0	r ≤ 1	r ≤ 2
Eigenvalue		0.49	0.33	0.04
Trace		34.78**	13.73*	1.25
Critical	1%	29.75	16.31	6.51
Value	5%	24.31	12.53	3.84

Test assumption: assumes no deterministic trend in data (intercept (no trend) in CE-no intercept in VAR)

		Null hypothesis		
		r = 0	r ≤ 1	r ≤ 2
Eigenvalue		0.52	0.45	0.07
Trace		43.76**	21.14*	2.34
Critical	1%	41.07	24.60	12.97
Value	5%	34.91	19.96	9.24

Test assumption: test allows for linear deterministic trend in data (intercept (no trend) in CE- and test VAR)

		Null hypothesis		
		r = 0	r ≤ 1	r ≤ 2
Eigenvalue		0.48	0.27	0.01
Trace		30.41*	10.29	0.45
Critical	1%	35.65	20.04	6.65
Value	5%	29.68	15.41	3.76

*(**) denotes rejection of the hypothesis at 5% (1%) significance level.

Index, and real credit balances. In contrast, consumer price index does not seem to be equilibrium-correcting in the VEC model. Real credit appears to be positively equilibrium-correcting to EC_{2t-1} only.

In the short-run, the estimation results indicate that that real GDP is positively influenced by its lag, and negatively by real credit balances. On the other hand, consumer price index appears to be positively affected with real credit balances. Finally, real credit balances turn out to be positively affected by real GDP. Supply shocks turn out to have an impact which is statistically negative on real GDP and positive on prices at 1% significance level, while

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demand shocks turn out to have no impact which is statistically significant on both real GDP and prices.

Table 4
Estimation Results of VEC Model

	r = 2		
	Δy_{1t}	Δy_{2t}	Δy_{3t}
Long-term			
EC_{1t-1}	-0.42*	0.16	-0.16
	(2.26)	(0.82)	(0.50)
EC_{2t-1}	0.05*	-0.00	0.09*
	(2.45)	(0.09)	(2.36)
Short-run			
Δy_{1t-1}	0.37*	-0.18	0.57*
	(2.81)	(1.34)	(2.52)
Δy_{2t-1}	-0.10	0.37	0.21
	(0.45)	(1.62)	(0.56)
Δy_{3t-1}	-0.17*	0.23*	0.26
	(2.13)	(2.78)	(1.83)
Determination			
μ	0.06**	0.01	0.04
	(3.13)	(0.31)	(1.11)
D1	-0.10**	0.09**	-0.05
	(5.55)	(4.69)	(1.43)
D2	0.05	0.03	0.01
	(1.76)	(0.99)	(0.25)
D3	-0.08*	0.03	-0.34**
	(2.14)	(0.68)	(5.08)
D4	-0.05	-0.00	0.30**
	(1.01)	(0.07)	(3.73)
Summary Statistics			
R^2	0.85	0.76	0.84
Adj-R^2	0.79	0.65	0.78
AIC	-6.63	-6.59	-5.54
SC	-6.17	-6.12	-5.08

*(**) denotes rejection of the hypothesis at 5% (1%) significance level

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These findings suggest that real GDP, consumer price index, and real credit would not move too far away from each other, displaying a co-movement phenomenon. An important finding is that real GDP turns out to have a positive impact on real credit in the short run as well as in the long run, suggesting that an increase in real output will in a way result in an increase for credit to finance investment. This implication is very consistent with the view points that firms in less developed countries are likely to rely on banks in financing investment projects.

5. Concluding Remarks

In this study the relationship among real GDP, consumer price index, and real credit is analyzed using Jordanian data for the period 1969-2001. Because cointegration equations require the use of nonstationary variables and error-correction equations require the use of stationary variables, each data series is first examined for the probable order of difference stationary. The estimation results indicate that nonstationarity cannot be rejected for the first difference of variables.

The long-run relationship among the variables is tested using multivariate cointegration system and employing the methodology of cointegration analysis as suggested by Johansen (1988) and Johansen and Juselius (1990). Two cointegration relations are found in the empirical analysis. This finding enables one to set up a vector error-correction model.

A VEC model is estimated including error-correction mechanisms, short-run relationships, and a set of dummy variables. The estimation results indicate that real GDP is negatively and positively equilibrium-correcting to EC_{1t-1} and EC_{2t-1}, respectively. This finding provides empirical evidence supporting the existence of the co-movement between real GDP, consumer price index, and real credit. Consumer price index is not equilibrium-correcting in the VEC model. Real credit is positively equilibrium-correcting to EC_{t-2} only.

In the short-run, the estimation results indicate that that real GDP is positively and negatively influenced by its lag and the lag of with real credit, respectively. On the other hand, consumer price index appears to be positively affected by real credit balances, while real credit turns out to be positively affected by real GDP. Supply shocks turn out to have an impact which is statistically negative on real GDP and positive on prices at 1% significance level, while demand shocks turn out to have no impact which is statistically significant on both real GDP and prices.

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النمو الاقتصادي والتسهيلات الائتمانية
نموذج متجه الخطأ المعدل

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ملخص

قامت الدراسة بفحص طبيعة العلاقة بين الناتج المحلى الإجمالي الحقيقي، والرقم القياسي للأسعار، ورصيد التسهيلات الانتمانية في الأردن خلال الفترة 1969-2001.أظهر اختبار ديكي-فولر عدم استقرار السلاسل الزمنية المستخدمة في الدراسة. كما أظهرت نتائج اختبار جوهانسن وجود علاقة تكامل مشترك بين المتغيرات في الأجل الطويل. إن عدم استقرار السلاسل الزمنية التي تتصف بوجود التكامل المشترك بينها أتاح استخدام نموذج متجه الخطأ المعدل التي اظهر أن الزيادة في الناتج المحلي الإجمالي تودي إلى زيادة في التسهيلات الانتمانية.

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Appendix A

Lag order determination tests results (multivariate)

Cointegration Type [a]	Number of Cointegration relations(r)	Lags	AIC	SC
1	1	1	-19.53	-18.28
		2	-19.25	-17.57
2	1	1	-19.51	-18.22
		2	-19.26	-17.53
3	1	1	-19.73	-18.34
		2	-19.31	-17.49
4	1	1	-19.68	-18.25
		2	-19.52	-17.66
5	1	1	-19.58	-18.05
		1	-19.63	-17.67
1	2	1	-19.54	-18.01
		2	-19.09	-17.12
2	2	1	-19.67	-18.05
		2	-19.20	-17.14
3	2	1	-19.66	-18.00
		2	-19.18	-17.08
4	2	1	-19.55	-17.79
		2	-19.55	-17.35
5	2	1	-19.49	-17.68
		2	-19.50	-17.25

a refers to cointegrating equation (CE) and VAR specification: 1 = no intercept or trend in CE or test VAR; 2 = intercept (no trend) in CE-no intercept in test VAR; 3 = intercept (no trend in CE and test VAR; 4 = intercept and trend in CE –no trend in test VAR; 5 = intercept and trend in CE, linear trend in VAR.

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