INTERNATIONAL JOURNAL OF BUSINESS, 19(3), 2014 ISSN: 1083-4346

Does An Emerging Equity Market Stimulate Long-Term Economic Growth? Evidence from Jordan

Mohammad I. Elian[a*] and Adil H. Suliman[b]

[a] *Department of Economics and Finance,*
Gulf University for Science and Technology (GUST), Kuwait
elian.m@gust.edu.kw

[b] *Department of Economics and Finance,*
Gulf University for Science and Technology (GUST), Kuwait
ElSharif.A@gust.edu.kw

ABSTRACT

After considering the time series characteristics of our dataset, this paper examines the causal relationship between financial development and economic growth in Jordan during 1980-2009. Financial development is proxied by both equity market capitalization and bank credit supplied to the private sector. Equity market liquidity (traded values) is used as a robustness test. After applying Granger causality tests, by using the cointegration and vector error-correction (VEC) framework, our results support a limited bi-directional causality relationship between equity market and economic growth. A bank has dual impact on the equity market and economic growth, though it has greater influence on the equity market, reflecting its vital role in both financial development and economic growth. Overall, these findings suggest the need to improve the efficiency of the financial system, especially the banking system, to channel more financial resources into its most productive areas and to thereby promote long-term economic growth.

JEL Classifications: *C32, F43, O4, O53*

Keywords: *financial development; economic growth; cointegration; granger causality; vector error correction models*

I. INTRODUCTION

The theoretical and empirical literature shows that extensive work has been done to examine the causal relationship between financial development and economic growth. The first part of the literature supports the supply-leading hypothesis, or the "finance-led growth" hypothesis, which considers the development of financial sectors as the precondition for economic growth. This hypothesis asserts that development in equity markets and financial intermediaries will increase the supply of financial services, thus causing an expansion of economic growth (see Bencivenga and Smith, 1991; Pagano, 1993; King and Levine, 1993; Demetriades and Maksimovic, 1996; Levine, 1997; Levine and Zervos, 1998; Yang and Yi, 2008). The second part of the literature emphasizes demand-following theories, or the "growth-led finance" hypothesis, which consider financial development as a result of economic growth and claim that economic growth normally causes an increase in demand for financial services, leading in turn to financial market development (see Romer, 1986; Shahnoushi et al., 2008).[1] The third part of the literature focuses on the "feedback" hypothesis, which suggests a bi-directional relationship between finance and economic growth (see Stiglitz, 1985; Boyd and Prescott, 1986; Mayer, 1988; Bencivenga and Smith, 1991 and 1993; Demetriades and Hussein, 1996; Gurspu and Muslomov, 1998; Arestis et al., 2001; Suleiman and Aamer, 2008).

Empirical analyses of the causality relationship between financial development and economic growth have been extensive over the years, in order to cover different developed and developing financial markets, as well as newly emerging markets. King and Levine (1993) use a cross-sectional and panel regression of 80 developed and developing countries, finding a strong and robust correlation between financial development and economic growth. Jung (1986) investigates evidence on the causality between financial and real development, using annual data on 56 countries, 19 of which are industrial countries. He reaches a slightly different conclusion when he considers developing countries and less-developing countries separately, with the result indicating moderate support for the supply-leading phenomenon in less well-developing countries. For less developed countries, Asli and Levine (1996) and Filler et al. (1999) both suggest that policymakers should remove impediments to the stock markets in order to open up the role of financial development in economic growth. Mohtadi and Agarwal (2004) examine the relationship between stock market development and economic growth for 21 emerging markets, using a dynamic panel regression, and find a positive relationship between stock market development and economic growth. For the less well-developing countries of the Middle East and North Africa (MENA), Ben Naceur and Ghazouani (2005) find no significant relationship between banking and stock market development and growth, in a dynamic panel model with GMM estimators; in fact, they indicate that the association between bank development and economic growth is actually negative after controlling for stock market development. They link this result to underdeveloped financial systems in the MENA region which hamper economic growth. Spears (1991) examines the causal relationship between financial intermediations and economic growth for five Sub-Saharan African countries, and he finds strong support only for the supply-leading hypothesis, or the "finance-led growth" hypothesis, that considers the development of financial sectors as the precondition for economic growth. For single, small open economies, for example Tunisia, Ghali (1999) concludes that financial development causes economic growth. For the Egyptian experience, Suleiman and Aamer (2008) indicate that there is bi-directional causality between financial development and economic growth, with financial development causing economic growth through increasing resources set aside for investment and improving efficiency. Shahnoushi et al. (2008) establish that only economic growth leads to financial development in Iran, concluding that financial development will not be an effective factor in economic growth. For more literature on this issue, see the footnote below.[2]

This paper tests the term and the direction of causality between financial development and economic growth for Jordan between 1980 and 2009, using a cointegration and vector error-correction (VEC) framework. Testing the relationship between financial development and economic growth, by using Granger causality, offers a good opportunity to capture the dynamics of the relationship between these variables, in both the short and the long term. The use of time series data analysis, which relies on cross-sections and ignores country-specific dynamics, has provided only a pooled estimation of the impact between these variables which does not consider the country-specific effect. Suleiman and Aamer (2008) indicate that a significant coefficient of the measure of financial development in growth regressions does not necessarily imply causality running from finance to growth, or vice versa. We believe that time series studies for a specific country, using Granger causality tests and a cointegration and vector error-correction (VEC) framework, in conjunction with an appropriate selection of financial proxies, will provide more precise estimations of the causal relationship between financial development and economic growth. Our trivariate vector error-correction system includes both equity market capitalization and bank credit, supplied to the private sector as proxies for financial development, and GDP per capita as a proxy for economic growth, while equity market liquidity is used as a robustness test. To validate our results further, we use a variance decomposition test to examine the relationship between variables beyond the sample period, in order to gauge the relative importance of financial development in explaining economic growth beyond this time frame. The variance decomposition test will also enrich our understanding of the relationship between financial development and short- and long-term economic growth.

The Jordanian economy is considered a small economy with limited natural resources and is supported heavily by dwindling external resources, while Jordanian financial sectors are seen as being relatively well-developed and regulated compared with other small economies. Since 1980, the Jordanian economy has passed through many stages of economic, financial, and banking reform, particularly on the institutional and supervision fronts, in a move to improve the entire financial system and make it capable of withstanding long-term economic growth.[3] To maintain Jordan's current prospects for economic growth, there is significant public and private enthusiasm for maintaining the current direction of reforms, including furthering financial and capital market reform. As part of this effort, the private sector continues to grow as a leading player in transforming local and international investments into productive services and industries, in order to promote sustainable economic growth. However, the question which still remains unanswered is: To what extent does all this effort accomplish its objective by increasing economic opportunities and facilitating the fairer and more efficient distribution of financial resources to its most effective uses, in order to promote long-term economic growth? This paper, to our knowledge, is the first to test the financial development economic growth nexus in Jordan in an attempt to establish its validity in the literature. The rest of this article is organized as follows: Section II reviews economic and financial developments in Jordan over the past three decades. Section III reviews the methodology specification, by presenting the proxy measurement of and economic rationale behind the variables, designating the specification of the model, developing the Johansen cointegration test, and describing the vector error correction models. Section IV presents the empirical results, and concluding remarks and policy implications are given in Section V.

II. THE JORDANIAN ECONOMY—AN OVERVIEW

The Jordanian economy is considered a relatively small economy with limited natural resources. In fact, such a small economy is supported heavily by external resources, including foreign and Gulf regional aid and remittances. However, mainly during the second half of the 1980s, the availability of external resources became limited, a reflection of major changes in the region. These included, for example, the economic slowdown in oil prices after 1983 and the reduction in Gulf aid to Jordan as a result of the first Gulf War (Alissa, 2007). These pressurized conditions during the 1980s, along with currency flotation in October 1988 and the liberalization of the interest rate regime in 1989, all led to the following: (1) A reduction in the value of the Jordanian currency, which had depreciated by more than 33% as of October 1988, and (2) an appreciation in external and internal debt, which increased from 97% of GDP to 200% of GDP for 1988 and 1989, respectively. As for 1989, there was a reduction in real GDP by 10.7%, an increase in the inflation rate of 25.7%, and a budget deficit, excluding grants, of up to 24% of GDP.

Against such unstable economic conditions, by 1989, the intervention of the International Monetary Fund (IMF) and the World Bank was given priority, and three phases of reforms were applied (Alissa, 2007).[4] The first reform phase, 1989-1991, aimed at stabilizing the economy and helping it recover from negative economic growth,[5] while the second phase, in 1992-1999, aimed at implementing limited structural adjustment measures, trade and financial liberalization, and privatization, in order to enhance exports and place more emphasis on expanding the contribution of the private sector to economic development. The third phase, during 1999-2007, aimed at promoting economic development through more trade and financial liberalization, establishing special economic zones, placing more weight on accelerating the privatization process, and enhancing pension reforms, thereby leading to more integration with the global economy, i.e. entering into free trade agreements with the USA (October 2000), the European Union (May 2002), and the World Trade organization (WTO) (April 2000). In order to ensure that these structural adjustment measures would lead the economy toward steady economic growth, Jordanian policymakers imposed more financial reforms to deepen equity and debt markets and to enhance both the volume and efficiency of overall economic sectors.[6] As a result, from 2003 to the first half of 2008, the Amman Stock Exchange (ASE) displayed an outstanding performance record in regard to a number of targeted indicators, including a number of listed firms, market caps, value traded, average daily trading, and traded shares.[7] This success was attributed to the rise in stock prices of heavyweight blue-chip firms, namely Jordan Phosphate Mines and the Arab Potash Company. The number of listed firms rose from 161 in 2003 to 261 by November 2008, while market capitalization surged from $11 billion, or 117% of GDP, in 2003 to $57 billion in June 2008, before sliding back to $35 billion in November, or 230% of the 2007 GDP. Monthly trading rose from $218 million in 2003 to nearly $5 billion in June 2008, and then it eased back to $1.2 billion in November 2008.[8] For the banking sector in Jordan, which is considered relatively developed and well-regulated, with a limited range of banking products, credit supplied to the private sector recorded double-digit growth for the period 2000-2008, with an increase of 16.5% up to September 2008 after annual growth of 22% over several years. Nevertheless, the banking industry remains burdened by non-performing loans that have been estimated to account for 20% of total outstanding loans. In fact, since 2002, the Jordanian banking industry has become engulfed in scandal, suggesting that stricter credit controls are needed, together with closer supervision of banks and exposure to the equity market. It is clear that Jordan's growth and globalization are linked inextricably to its new economic and financial structure. The Jordanian financial market, emerging as a competitive player in the global marketplace, is the centerpiece of the country's economic growth.

III. METHODOLOGY SPECIFICATION

A. Proxy Measurement and Economic Rationale

As a proxy for measuring indicators, by following standard practice, our indicator of economic growth (Y1) was GDP per capita, which was measured as the ratio of real GDP to total population. Financial development was proxied by two variables: Stock market size (M1) and stock market liquidity (M2). For M1, it was expected that overall market size or capitalization would be correlated positively with the ability to utilize capital and diversify risk on an economy-wide basis. The stock market size proxy was calculated by dividing the value of listed local stocks by GDP. For M2, we used the value traded ratio as a market liquidity indicator. M2 was used to indicate the ability of investors to buy, sell, and modify their investment portfolios in a short time and by incurring low transaction costs. This provided the chance to reduce investment risk and facilitate investment for profitable and growth-sustainable projects. We considered liquidity as an indication of stock market development by signaling the market's ability to enhance the allocation of capital, and, in turn, reflect the possibility of long-term economic growth. Furthermore, market liquidity, as indicated by the value of traded shares, which was also used as an activity proxy, was used to capture trading relative to the size of the economy. Market liquidity was measured by dividing the value of local stock trades on the local exchange by GDP. This ratio measured the organized trading of firm equity as a share of national output, and hence it was expected to reflect positively higher liquidity positions on an economy-wide basis. Yet, theoretical models of stock market liquidity and economic growth directly motivated the value traded, indicating positive liquidity for the overall economy.

For banking development, following Levine and Zervos (1998), Rousseau and Wachtel (2000), Beck and Levine (2002), and Ben Naceur and Ghazouani (2005), we used the value of commercial banks' loan portfolio (credit provided to the private sector) to GDP as a proxy for banking development (BANK). This measure has the advantage of excluding credit provided to the public sector, thereby measuring more precisely the contribution of financial intermediaries in funding the private sector. In fact, banks can select entrepreneurs who are capable of adopting money-spinning ventures, guiding savings to higher returns and more productive projects, lowering monitoring costs, and reducing credit controls, all of which provide the chance for more economic growth. Data related to economic growth and financial development, including stock market and banking indicators, were extracted from the World Bank database. Each group of data was imported from one single source, and as a result we overcame any potential consistency and measurement problems. As a small economy within emerging stock markets, the sample included all 31 years (1979-2009) on an annual basis.[9]

B. Model Specification and the VAR Framework

In this article, we applied the vector autoregressive model (VAR) methodology to test for the short- and long-term relationship between economic growth (Y1), equity market capitalization (M1), and bank credit provided to the private sector (BANK). We built a general VAR model to be estimated separately for each targeted variable's time series, and we regressed each variable against the other variables. Following the standard literature, we specified the bi-directional relationship between the Y1, M1, and BANK variables by applying Equation (1) below. In addition to M1, in the robustness test, we also used equity market liquidity (M2) as a traditional measure of financial development:

$$\gamma_t = f\,(\text{M1, BANK}) \tag{1}$$

We assume that the level of γ_t can be modeled as a nonstationary i^th^ order vector autoregression equation:[10]

$$\gamma_t = \Gamma + \phi_1\gamma_{t-1} + \phi_2\gamma_{t-2} + \phi_3\gamma_{t-3} + \ldots\ldots\ldots\ldots + \phi_{i-1}\gamma_{t-i+1} + \phi_i\gamma_{t-i} + \nu_t \tag{2}$$

This VAR (*i*) model can be remodeled as:

$$\Delta\gamma_t = \Gamma + \Phi_1\Delta\gamma_{t-1} + \Phi_2\Delta\gamma_{t-2} + \Phi_3\Delta\gamma_{t-3} + \ldots\ldots\ldots\ldots + \Phi_{i-1}\Delta\gamma_{t-i+1} + i\gamma_{t-1} + \nu_t \tag{3}$$

where γ_t is a 3×1 vector of the first-order integrated [I(1)] variables [y1, m1, BANK], and the same for γ_t = [y1, m2, BANK] in the robustness test.

$$\Phi_k = -\,[\phi_{k+1} + \phi_{k+2} + \ldots\ldots\ldots\ldots + \phi_i] \text{ for } k = 1, 2, 3, \ldots\ldots, {}_{i\text{-}1} \text{ and } i = \phi(1)$$

where Φ_i is a (3×3) coefficient matrix of parameters, while ν_t is a vector of normally and independently distributed error terms. Johansen (1988) and Johansen and Juselius (1992) proposed two test statistics, the trace (λtrace) and the maximum Eigenvalue (λmax), in order to test for the number of cointegrated vectors in the VAR framework.[11]

In our case, if there exist $r(0 < r < 3)$ cointegrating vectors, this implies that i is rank-deficient, in which case i can be decomposed as $i = \psi\lambda$, where $\psi_{(3*r)}$ and $\lambda_{(r*3)}$. Therefore, Equation (3) can be reformulated into a vector error correction model (VECM) as follows:

$$\Delta\gamma_t = \Gamma + \Phi_1\Delta\gamma_{t-1} + \Phi_2\Delta\gamma_{t-2} + \Phi_3\Delta\gamma_{t-3} + ... + \Phi_{i-1}\Delta\gamma_{t-i+1} + \psi\lambda\gamma_{t-1} + \nu_t \quad (4)$$

where Δ denotes the first difference and ψ is the error correction term (the cointegration term) lagged by one period, indicating the speed of adjustment toward the long-run equilibrium. λ is interpreted as distinct cointegration vectors, while ν_t is the white noise term. The linear combinations $\lambda\gamma_{t-1}$ are stationary; thus, variables in Equation (4) are stationary. In our VAR model, considering $\Psi_{i,t-1}$ is the i^{th} error correction term lagged by one period, Equation (4) can be remodeled as:

$$\Delta Y_t = \Gamma_1 + \sum_{i=1}^{m}\alpha_1, i\Psi_{i,t-1} + \sum_{k=1}^{p-1}\beta_{11},_k\Delta Y_{1,t-k} + \sum_{k=1}^{p-1}\beta_{12},_k\Delta EM_{t-k} + \sum_{k=1}^{p-1}\beta_{13},_k\Delta BANK_{,t-k} + \varepsilon_i \quad (5)$$

$$\Delta EM_t = \Gamma_2 + \sum_{i=1}^{m}\alpha_2, i\Psi_{i,t-1} + \sum_{k=1}^{p-1}\beta_{21},_k\Delta Y_{1,t-k} + \sum_{k=1}^{p-1}\beta_{22},_k\Delta EM_{t-k} + \sum_{k=1}^{p-1}\beta_{23},_k\Delta BANK_{,t-k} + \varepsilon_i \quad (6)$$

$$\Delta BANK_t = \Gamma_3 + \sum_{i=1}^{m}\alpha_3, i\Psi_{i,t-1} + \sum_{k=1}^{p-1}\beta_{31},_k\Delta Y_{1,t-k} + \sum_{k=1}^{p-1}\beta_{32},_k\Delta EM_{t-k} + \sum_{k=1}^{p-1}\beta_{33},_k\Delta BANK_{,t-k} + \varepsilon_i \quad (7)$$

where $\Psi_{i,t-1}$ is the i^{th} error correction term lagged by one period, and $\beta_{hj},_k$ typifies the impact of the k^{th} lagged value of variable j on the current value of variable h: hj = Y1, EM (described as M1), and BANK.

IV. EMPIRICAL RESULTS

In order to examine the causal relationship between the Y1, M1, and BANK series, we tested initially for the order of integration of each series, using the standard unit root test, namely the ADF test (Dickey & Fuller, 1979). Both the Kwiatkowski-Phillips-Schmidt-Shin (KPS) and Phillips-Perron (PP) (Phillips and Perron, 1988) unit root tests[12] were used as confirmation tests. Secondly, we used the trivariate framework of Johansen and Juselius (1990) to determine the number of cointegrating vectors between the series. Finally, a vector error correction model (VECM) was fitted to the data, to uncover the relationship dynamic between the series. The results for the unit root tests, which are reported in Table 1 and Table 2, confirmed that all the series were non-stationary processes on their own levels, but they were stationary in the first difference, thus confirming that these series were integrated to the order of one I(1).

Table 1
ADF unit root test results—Level

Series	ί	τ	ή
Y1	2.4032(0)	4.1248(0)	1.4081(2)
M1	-2.9873(4)	-1.4725(0)	1.7074(5)
M2	-2.8467(0)	-2.8127*(0)	-1.8917*(0)
BANK	-2.5619(1)	-0.9244(1)	-1.5742(1)

Table 2
ADF unit root test results—1st difference

Series	ί	τ	ή
Y1	-1.0178(1)	0.0438(1)	0.6181(1)
M1	$-6.3591^{***}(0)$	$-6.4857^{***}(0)$	$-6.4995^{***}(0)$
M2	$-6.8168^{***}(0)$	$-6.9549^{***}(0)$	$-7.0783^{***}(0)$
BANK	$-6.3432^{***}(0)$	$-6.1453^{***}(0)$	$-6.2515^{***}(0)$

Notes: (1) The *t* statistic refers to the ADF tests. Subscriptions ί, τ, ή indicate the models that allow for an intercept and trend (constant, linear trend), intercept (constant), and none, respectively. (2) Asterisks ***, **, and * show significance at 1%, 5%, and 10%, respectively. Figures in parentheses indicate lag length (SBC). Y1 shows stationary under the 2nd difference ($-9.8187^{***}(0)$, $9.0672^{***}(0)$, and $-9.1175^{***}(0)$ for ί, τ, ή, respectively).

Finding that all the series were integrated to the order of one I(1), the second step involved testing for the number of cointegration relationships between these series, by using the Johansen and Juselius (JJ) cointegration framework. The JJ test is very sensitive to the choice of lag length and to the specification of the deterministic components that need to be included in the test. Selecting the wrong deterministic component specification can lead to wrong conclusions from the data. To determine the appropriate lag structure to be used in the JJ cointegration test, a VAR model was fitted to the data. Both the Akaike information criterion (AIC) and the Schwarz information criterion (SC) were used to select the appropriate lag, from which lag three was selected as being the optimal lag for testing by both criteria. To determine the most appropriate deterministic components for the JJ test, Hansen and Juselius (1995) suggest a procedure called the Pantula principle, which simultaneously determines rank and deterministic components in the JJ cointegration framework test, based on five cases suggested by Johansen (1995). Case 1 suggests that level data have no deterministic trends and the cointegrating equations have no intercepts. Case 2 suggests that level data have no deterministic trends but the cointegrating equations have intercepts. Case 3 suggests that level data have linear trends but the cointegrating equations have only intercepts. Case 4 suggests that level data and the cointegrating equations have linear trends. Case 5 suggests that level data have quadratic trends and the cointegrating equations have linear trends. The five cases are nested so that Case 1 is contained in Case 2, which is then contained in Case 3, and so on. In order to choose one of these cases, firstly we tested the null hypothesis of zero cointegrating vectors for Case 1. The test should move from the most restrictive case, Case 1, to the least restrictive case,

Case 5, and at each stage compare the trace (λtrace) or the maximum Eigenvalue (λmax) statistics to a critical value and only stop the first time the null zero cointegrating vector hypothesis is not rejected. Hjelm and Johansson (2005) indicated that the Pantula principle is partially likely to choose a wrong deterministic component, so they modified it to improve its ability in this respect. They suggested that cases that are not well-suited to economic theory should be excluded before the procedure is applied to the data.[13] The ability of the modified Pantula principle to find the correct component depends on the properties of the dataset to hand. If the trend is weak and/or the sample size is small, this will have an adverse effect on the power of the result.

We used the JJ cointegration framework test, based on the so-called Pantula principle, to determine simultaneously the rank and the deterministic components, the results for which are displayed in Table 3. The first time the null hypothesis of zero cointegrating vectors is not rejected at the 1% significance level is indicated by *, while the Eigenvalue statistics of 25.7788 are less than the critical value of 27.0678. One cointegration vector is accepted, whereas level data have no deterministic trends and the cointegrating equations have an intercept. Cheung and Lai (1993) indicate that Johansen's cointegration tests are more sensitive to under-parameterizations in lag length, while they are not so sensitive to over-parameterizations. AIC and SC unit root tests are useful for choosing the optimal lag length for Johansen's tests, but these tests perform poorly in the presence of moving-average dependence. Given that our sample size was small, it was desirable to investigate the sensitivity of the tests for under-parameterization. As presented in Table 4, an additional test with a lower lag, lag 2, was undertaken. Moving through Table 5, the first time the null hypothesis of zero cointegrating vectors is not rejected, at the 1% significance level, is indicated by *. The Eigenvalue statistics are 16.1863, greater than the critical value of 22.2517. One cointegration vector is accepted, but the result suggests that the level data have no deterministic trends and the cointegrating equations do not have an intercept. It is clear that when a relatively low-order autoregressive model is used, maximal Eigenvalue tests are biased seriously toward finding a more restrictive deterministic component.

Table 3
Cointegration rank and deterministic components within the JJ framework

		No Deterministic Trend	Restricted Constant
Lag 3		A (case 1)	B (case 2)
		λmax	λmax
Null	Alternative	Max-Eigen	Max-Eigen
$r = 0$	$r = 1$	22.8510	25.7788
		(22.2517)	(27.0678)*
$r \leq 1$	$r = 2$	3.8448	18.8414
		(15.0913)	(20.1612)
$r \leq 2$	$r = 3$	0.2386	3.6603
		(6.9406)	(12.7608)

* denotes the selected model.
MacKinnon-Haug-Michelis (1999); critical values are reported in parentheses.

Table 4
Cointegration rank and deterministic components within the JJ framework

		No Deterministic Trend	Restricted Constant
Lag 2		A	B
		λmax	λmax
Null	Alternative	Max-Eigen	Max-Eigen
r = 0	r = 1	16.1863	40.3749
		(22.2517)*	(27.0678)
r ≤ 1	r = 2	1.8228	13.3159
		(15.0913)	(20.1612)
r ≤ 2	r = 3	0.2268	1.7805
		(6.9406)	(12.7608)

* denotes the selected model.
MacKinnon-Haug-Michelis (1999); critical values are reported in parentheses.

Table 5
Johansen cointegration rank test (maximum eigenvalue)

Hypothesized		Max-Eigen	0.01
No. of CE(s)	Eigenvalue	Statistic	Critical Value
None *	0.7146	32.5919	27.0678
At most 1	0.5213	19.1538	20.1612
At most 2	0.3579	11.5167	12.7608

Max-Eigenvalue test indicates 1 cointegrating eqn(s) at the 0.01 level.
* denotes rejection of the hypothesis at the 0.01 level.

We tested for the number of cointegration vectors within the JJ framework among the variables, using lag 3 as an optimal lag, in which level data had no deterministic trends and the cointegrating equations had an intercept. The result, which is displayed in Table 5, supports the existence of a unique long-term relationship between Y1, M1, and BANK at the 1% significance level.

Given that the variables included in the VAR model were found to be cointegrated, a vector error correction model (VECM) including the error correction term was estimated, in order to investigate the dynamic behavior of the model. The VECM showed how the examined model adjusted in each time period toward its long-term equilibrium. Both β (lag variable coefficients) and α (adjustment coefficients) were estimated for the cointegrating vector by applying the JJ technique. The estimated coefficients and the associated *t*-values were obtained and are reported in Table 6. As appears from β coefficients, the vector possibly contains information about the relationship between Y1, M1, and BANK. Most of the lags' coefficients are highly significant, signifying the existence of short-term relationships between economic growth and financial development indicators. Adjustment coefficients are significant and have negative signs, confirming that there is an adjustment to the steady-state equilibrium in the long run between Y1, M1, and BANK. Residual Portmanteau tests for autocorrelations, and VEC residual serial correlation LM tests, indicate that there are no autocorrelations or residual serial correlation up to lag 6, therefore confirming that there is nothing to suggest that the model is mis-specified.

Table 6
Estimated vector error correction

Normalized	Y1	M1	BANK
	1.000	0.680	1.091
		(2.651)**	(2.684)**
α	0.075	-0.638	-0.165
	(1.719)*	(-4.413)***	(-2.236)**
Π (β)			
Y(-1)	-0.440	2.252	-0165
	(-1.799)*	(2.286)**	(2.236)**
Y(-2)	-0.156	0.185	0.411
	(0.693)	(0.246)	(0.911)
Y(-3)	0.407	-1.491	-0.119
	(1.986)*	(-2.184)**	(-0.344)
M(-1)	0.043	-0.414	0.0451
	(0.926)	(-2.639)**	(0.569)
M(-2)	-0.081	-0.347	0.041
	(-1.895)*	(-2.427)**	(0.551)
M(-3)	0.035	-0.346	-0.089
	(0.758)	(-2.235)**	(-1.136)
BANK(-1)	-0.252	2.611	0.583
	(-1.817)*	(5.641)***	(2.473)**
BANK(-2)	-0.295	1.327	0.122
	(-1.505)*	(2.036)**	(0.368)
BANK(-3)	0.307	-0.803	0.293
	(2.189)**	(-1.717)*	(1.005)
Break	0.064	0.283	0.088
	(1.958)*	(2.610)**	(1.595)
Adj. R-squared	0.45	0.70	0.25

Both residual Portmanteau tests for autocorrelations and VEC residual serial correlation LM tests indicate that there is no Autocorrelation or Residual Serial Correlation up to lag 6.
Asterisks ***, **, and * show significance at 1%, 5% and 10%, respectively.

Table 7
VEC pairwise Granger causality tests

	F-statistic		t-statistic	
	A	B	C	D
Lead (Dependent) variable	Y1	M1	BANK	
Excluded				
Y1	~	$(8.563)^{**}$	(1.779)	$(3.531)^{*}$
M1	$(6.558)^{*}$	~	(3.873)	$(12.111)^{***}$
BANK	$(9.921)^{**}$	$(40.774)^{***}$	~	$(3.955)^{**}$

Notes: All variables except for the lagged error correction terms are in the first difference. The reported *F*-statistics are block exogeneity Wald-type causality tests from the estimated VECM. Block exogeneity refers to the exclusion of all the endogenous variables from the VECM other than the lags of the dependent variable. The *t*-statistics test each lagged error correction term where the null is equal to zero. The figures in the brackets are *t*-statistics. ***, **, and * indicate that the null hypothesis ("no Granger causality") can be rejected at the 1%, 5%, and 10% significance levels, respectively.

In a VECM, a short-term causal effect is indicated by the significance of lagged dynamic coefficients, while the long-term causal relationship is indicated by the level and significance of lagged error correction coefficients. The use of a block exogeneity Wald test allowed us to test the joint significance of lagged endogenous variables in each equation in the vector. The statistic for block exogeneity is the χ^2 (Wald) statistic for the joint significance of all other lagged endogenous variables in the equation. On the other hand, long-term causality was tested by checking the significance of the *t*-test on each lagged error correction term. The results of these tests are displayed in Table 7.

Looking at Table 7, column A, when Y1 is used as a lead variable, M1 lags are statistically significant at the 5% level. For column B, when M1 is used as a lead variable, Y1 lags are statistically significant at the 1% level. These results suggest that both series are strongly endogenous to the system in the short term—there is bi-directional causality between Y1 and M1 in the short-term period, confirming the previous results for the vector error correction in Table 6. Looking at the test results for lagged error correction terms in column D, the significance of the *t*-test on each lagged error correction term for both Y1 and M1 implies that when deviations from the equilibrium cointegrating relationship occur, M1 will adjust to clear this disequilibrium. These results again confirm our previous findings from the Johansen cointegration test in Table 5, and as such we can conclude that the Jordanian economy is influenced by stock markets in the long term.

For BANK, the results seem quite different—looking at Table 7, column A, BANK lags are statistically significant at the 5% level when Y1 is used as a lead variable. In column C, Y1 lags are not significant when BANK is used as a lead variable, which suggests there is one-way causality from BANK to Y1 in the short-term period. The significance of the *t*-test on the lagged error correction term for BANK implies that when deviations from the equilibrium cointegrating relationship occur (as measured by the error correction term), BANK will also adjust to clear this disequilibrium in the long term. Similar to the relationship between Y1 and BANK, the relationship between M1 and BANK indicates that there is one-way causality from BANK to M1. Looking at the M1 equation in column B, when M1 is used as a lead variable, BANK lags are highly statistically significant at the 1% level. On the other hand, for the BANK equation in column C, when BANK is used as a lead variable, the M1 lags are not statistically significant. In a nutshell, there is bi-directional causality between Y1 and M1, one-way causality from BANK to M1, and one-way causality from BANK to Y1 in the short term. In the long term, both M1 and BANK adjust to clear any deviations from the equilibrium or the cointegrating relationship between Y1, M1, and BANK. These results also indicate that BANK influences both Y1 and M1.

To summarize, our cointegration results support the existence of a bi-directional relationship between M1 and Y1, one-way causality from the BANK to M1, and one-way causality from BANK to Y1. In the long term, both M1 and BANK adjust to clear any deviations from the equilibrium, thereby confirming a long-term cointegrating relationship between Y1, M1, and BANK. The results are confirmed by using M2 as a robustness test,[14] reflecting clear bi-directional causality between Y1 and M2, one-way causality from BANK to M2, and one-way causality from BANK to Y1. In the long run, only M2 adjusts to clear any deviations from the equilibrium, or to support the cointegrating relationship between Y1, M2, and BANK.

To gauge the result further, we used variance decomposition to acquire information about the relative importance of each random innovation in affecting the variables in the VAR. The breakdown of the forecast variance error for each variable, into components attributed to each endogenous variable, enabled us to determine the relative importance of each variable in generating fluctuations in other variables in the model. Table 8 shows the decomposition of the forecast errors for 25 years. In order to assign variance shares to different variables, residuals in the equation were orthogonalized by using Choleski's decomposition method under the following order of variables: Y1, M1, and BANK. In Table 8, Panel A, the first row shows the percentage of the forecasting variance of Y1 that contributes to shock in M1. This forecasting variance is very small in value, at about .0523, and it reaches its maximum value in the short term, at five years, and then starts to die out. The second row shows the percentage of the forecasting variance Y1 that contributes to shock in BANK, at about 17.43—this variance has a greater percentage, and it increases gradually throughout the long term, 25 years, to reach its maximum value. In Panel B, the first row shows the percentage of the forecasting variance of M1 that contributes to shock in Y1; this variance reaches 18.66 in the short term and then continues to increase throughout the long term to reach 25.26 at the end of 25 years. The second row shows the percentage of the forecasting variance M1 that contributes to shock in BANK; this variance has the greatest percentage, reaching 52.27 in the short term, and then continues to increase throughout the long term to reach 65.83 at the end of 25 years. It is clear that while M1 explains a trivial and limited portion of the percentage of the forecasting variance of Y1, Y1 goes a little way to explaining the larger percentage of the forecasting variance of M1. The result for BANK is different, in that BANK explains not only a slightly larger portion of the percentage of the forecasting variance of Y1, but also a tremendous portion of the percentage of the forecasting variance of M1. To sum up the result from a variance decomposition point of view, although it supports the empirical results of the VECM, it nonetheless indicates that bi-causality between M1 and Y1 is short-lived and terminates quickly, with BANK playing dual roles, influencing both Y1 and M1, but having a greater influence on M1.

Table 8
Variance decomposition

Variance Decomposition of Y1					
	At 5 Years	At 10 Years	At 15 Years	At 20 Years	At 25 Years
Panel A					
M1	5.23	3.50	2.26	1.79	1.55
BANK	3.16	10.82	15.25	16.72	17.43
Variance Decomposition of M1					
	At 5 Years	At 10 Years	At 15 Years	At 20 Years	At 25 Years
Panel B					
Y1	18.66	22.24	23.48	24.57	25.28
BANK	52.27	59.78	63.79	65.09	65.83

V. CONCLUDING REMARKS AND POLICY IMPLICATIONS

The aim of this article was to test for the short- and long-term causal relationship between financial development and economic growth in Jordan between 1980 and 2009, by using a cointegration and vector error-correction (VEC) framework. Financial development is proxied by both equity market capitalization and bank credit provided to the private sector, and equity market liquidity is used as an additional variable for the robustness test. After applying Granger causality tests using the cointegration and vector error-correction (VEC) framework, the finding supports the bi-directional causality relationship between the equity market and economic growth. For BANK, it causes both equity market and economic growth. The finding from the variance decomposition tests indicates that the influence of the equity market on economic growth is small and limited, but BANK nevertheless does play a dual role in influencing both the equity market and economic growth.

In order to maintain sustainable economic growth, Jordanian policymakers have to undertake the necessary actions to excavate the long-term interrelationship between the equity market and real economic entities. These actions include placing greater emphasis on the functioning of the equity market, by focusing on its capability and creditability in decreasing savings' mobilization costs, as well as attracting more and productive foreign investment portfolios, particularly from the Gulf Cooperation Council (GCC) region. In order to attract more productive foreign investment there is a need to reduce all barriers to regional and/or international capital inflow, in order to have more capital liberalization. In this regard, policymakers have to drop impairments to the equity market, such as tax, legal, and regulatory barriers, in order to gain more and rapid capital accumulation and technological enhancements. In order to decrease savings' mobilization costs there is a need to boost equity market size and activity by adopting other operational actions, including less government intervention in the financial system, the fair privatization of major state-owned firms, more deregulation of facilitating partnership-based firms that unite domestic and foreign investors, and enhancing transparency to improve legislation and the business climate.

While the influence of the equity market on economic growth is limited, as it is small and terminates quickly, policymakers need to take action that will increase both the volume and efficiency of capital accumulation, increase investment and savings rates, finance productive projects, and increase the contributions of small and medium enterprises (SMEs), in order to increase greenfield investment in the economy. It must be kept in mind that the abovementioned actions may create more volatility in stock returns in the short term, but countries that are more open to capital inflow will have less volatile equity markets in the long term, compared to those countries with tighter capital controls (see Demirguc-Kunt and Levine, 1996, for more details).

For the dual role of the banking sector, influencing economic growth and the equity market, where the banking sector has a greater influence on the equity market and the equity market itself makes a minor contribution to economic growth, it is clear that there is a need to place more emphasis on credit allocation by strengthening credit controls and reinforcing competition within the banking industry. Although the Jordanian banking sector is considered well-regulated and performs adequately, it is still limited in terms of the range of products and services provided to the private sector, and it is still burdened by non-performing loans. Based on this finding, policymakers are advised to monitor closely banks' exposure to the equity market, and to direct the banking industry to expand its credit to non-service sectors, in order to promote sustainable and real investment projects that will reflect positively on the long-term economic growth of Jordan.

ENDNOTES

1. Finance-led growth supporting studies concluded that countries with better developed financial systems, including more efficient banks and well-organized stock markets, tend to grow faster by providing access to more funds needed for enterprises, thus leading to more economic growth. This development will be greater for industrialized economies than for agricultural economies (Beck 2002). Growth-led finance studies contend that stock markets do not necessarily cause high levels of economic growth, and stock markets may negatively impact economic growth due to the possibility of failure, e.g. Singh (1997), Singh and Weiss (1998).
2. Mayer (1988), Levine and Zervos (1998), Filler et al. (1999), Shahbaz et al. (2008), Yang and Yi (2008), Nowbutsing (2009), and Antonios (2010). Also, for banks, the evidence reinforces the importance of bank credit to economic growth, supporting earlier evidence that is concluded in Stiglitz (1985), Boyd and Prescott (1986), Bencivenga and Smith (1991), Greenwood and Javonovic (1990), Atje and Jovanovich (1993), Bencivenga and Smith (1993), and Arestis et al. (2001).
3. More details are available through the related section (Section II).
4. The goals were to stabilize the economy and control the transition from a state-dominated model to a more liberalized and privatized model in which the private sector could have a dominant role, in addition to transforming the economic structure of Jordan to a model that would generate self-sustaining international

activity, and to be more productive, more competitive, and integrated at regional and international levels.

5. This includes reducing the current account and budget deficit, controlling inflation rates, and foreign reserves.
6. MENA-OECD Investment Program, 2007, and Amman Stock Exchange, Monthly Statistical Bulletin, February 2006.
7. This performance was led by the banking industry, which increased by 134%, whilst the index for the insurance industry almost doubled; services increased by 78% and industry by 15%.
8. Economic and Financial Review, Jordan, November 2008.
9. Masih and Masih (1996) recommended using annual data for such types of research.
10. Similar models were used previously by Liang and Teng (2006) and Suleiman and Aamer (2008).
11. Under the λtrace, the H_o that there are r cointegrating vectors against the alternative of more than r cointegrating vectors is tested. When we do not reject H_o, this means that the trace statistic is lower than the critical value, as suggested by Johansen and Juselius (JJ). The H_o under the λmax statistic test is that the number of r cointegrating vectors is r against the alternative of (r+1) cointegrating vectors. Therefore, the H_o applies where r=0 is tested against the alternative that r=1, r=1 against the alternative r=2, r=2 against the alternative r=3, r=3 against the alternative r=4, and so on. The existence of r cointegrating vectors is accepted when the maximum Eigenvalue (λmax) statistic is lower than the critical value. In other words, the hypothesis of the existence of zero cointegration vectors (r=0) is tested by likelihood ratio (LR) statistics: λtrace and λmax. Then, it is proposed that the likelihood ratio test statistic (LR) of λtrace and λmax deterministic presence or no presence of the long-run relationship between variables in the model. As suggested by JJ, the likelihood ratio statistic (LR) for the trace test (λtrace) is computed as follows:
$$\lambda\text{trace}(r) = -n \sum_{i=r+1}^{p} \text{In}(1-\lambda i)$$
where, r = 0, 1, 2, 3, ……, p-1, n is the number of usable observations. λi is the largest estimated value of the i^{th} characteristic root (the estimated Eigenvalue) from the matrix. Alternatively, the likelihood ratio statistic (LR) for the maximum Eigenvalue (λmax) is computed as follows:
$$\lambda\max(r, r+1) = -n\text{In}(1-\lambda r+1) \text{ where, } r = 0, 1, 2, 3, \ldots\ldots, p\text{-}1.$$
12. The tests for the unit root are supported by using the Phillips-Perron (PP) test and the Kwiatkowski-Phillips-Schmidt-Shin (KPS) test; the results of which show no discrepancies. The Phillips-Perron (PP) test differs from the ADF test in how it deals with serial correlation and heteroskedasticity in errors; the test is robust to general forms of heteroskedasticity in the error term and in selecting its own lag specification. The Kwiatkowski-Phillips-Schmidt-Shin (KPS) test is run under the null hypothesis of stationarity.
13. See Hjelm and Johansson (2005) for a detailed discussion of the modified Pantula principle.
14. To save space, results are not reported, but they are available upon request.

REFERENCES

Alissa, S., 2007, "Rethinking Economic Reform in Jordan: Confronting Socioeconomic Realities." *Carnegie Middle East Centre* 1-23.

Antonios, A., 2010, "Stock Market and Economic Growth: An Empirical Analysis for Germany." *Business and Economic Journal,* 2010 BEJ-1: 1-12.

Arestis, P., P. Demetriades, and K. Luintel, 2001, "Financial Development and Economic Growth: The Role of Stock Markets." *Journal of Money, Credit and Banking* 33:16-41.

Asli, D., and R. Levine, 1996, "Stock Market, Corporate Finance, and Economic Growth: An Overview." *The World Bank Economic Review* 10(2): 223-239.

Atje, R., and B. Jovanovic, 1993, "Stock Market Development." *European Economic Review* 37: 632-640.

Beck, T., 2002, "Financial Development and International Trade: Is There a Link?" *Journal of International Economics* 57: 107-131.

Beck, T., and R. Levine, 2002, "Stock Markets, Banks and Growth: Panel Evidence." *NBER Working Paper* No. 9081, Cambridge, USA.

Bencivenga, V., and B. Smith, 1991, "Financial Intermediation and Endogenous Growth." *The Review of Economic Studies* 58: 195-209.

Bencivenga, V., and B. Smith, 1993, "Some Consequences of Credit Rationing in an Endogenous Growth." *Journal of Economics Dynamic and Control* 17: 97-122.

Ben Naceur, S., and S. Ghazouani, 2005, "Stock Markets, Banks, and Economic Growth: Empirical Evidence from the MENA Region." *Research in International Business and Finance* 21: 297-315.

Boyd, J., and E. Prescott, 1986, "Financial Intermediary-coalitions." *Journal of Economic Theory* 38: 211-232.

Cheung, Y., and K. Lai, 1993, "Finite-sample Sizes of Johansen's Likelihood Ratio Tests for Co-integration." *Oxford Bulletin of Economics and Statistics* 55: 313-328.

Demetriades, P., and K. Hussein, 1996, "Does Financial Development Cause Economic Growth? Time Series Evidence from Sixteen Countries." *Journal of Development Economics* 51(2): 387-411.

Demetriades, P., and V. Maksimovic, 1996, "Financial Constraints, Use of Funds and Firm Growth: An International Comparison." World Bank Working Paper No. 1671, USA.

Demirguc-Kunt, A., and R. Levine, 1996, "Stock Market Development and Financial Intermediaries: Stylized Facts." *The World Bank Economic Review* 10(2): 291-321.

Dickey, D., and W. Fuller, 1979, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root." *Journal of the American Statistical Association* 74(366): 427-431.

Filler, R., J. Hanousek, and N. Campos, 1999, "Do Stock Markets Promote Economic Growth?" The William Davidson Institute, Business School, University of Michigan, *Working Paper* No. 267, September, USA.

Ghali, K., 1999, "Financial Development and Economic Growth: The Tunisian Experience." *Review of Development Economics* 3(3): 310-322.

Greenwood, J., and B. Javonovic, 1990, "Financial Development, Growth and the Distributions of Income." *Journal of Political Economy* 98 (5): 1076-1107.

Gurspu, T., and A. Muslomov, 1998, “Stock Markets and Economic Growth: A Causality Test.” MBA: Istanbul Technical University, Turkey.

Hansen, H., K. Juselius, 1995, *Cats in Rats: Co-integration Analysis of Time Series*. Distributed by Estima, Evanston, IL, USA.

Hjelm, G., and M. Johansson, 2005, “A Monte Carlo Study on the Pitfalls in Determining Deterministic Components in Cointegrating Models.” *Journal of Macroeconomics* 27: 691-703.

Johansen, S., and K. Juselius, 1990, “Maximum Likelihood Estimation and Inference on Cointegration with Applications to the Demand for Money.” *Oxford Bulletin of Econometrics and Statistics* 52: 169-210.

Johansen, S., 1995, *Likelihood-Based Inference in Cointegrated Vector Auto-regressive Models*. Oxford University Press, Inc., New York.

Jung, W., 1986, “Financial Development and Economic Growth: International Evidence.” *Economic Development and Cultural Change* 34: 334-346.

King, R., and R. Levine, 1993, “Finance and Growth: Schumpeter Might Be Right.” *Quarterly Journal of Economics* 108(3): 717-737.

Levine, R., 1997, “Financial Development and Economic Growth: Views and Agenda.” *Journal of Economic Literature* 35: 688-726.

Levine, R., and S. Zervos, 1998, “Stock Markets, Banks and Economic Growth.” *American Economic Review* 88(3): 537-558.

Liang, Q., and J. Teng, 2006, “Financial Development and Economic Growth: Evidence from China.” *China Economic Review* 17: 395-411.

MacKinnon, J., A. Haug, and L. Michelis, 1999, “Numerical Distribution Functions of Likelihood Ratio Tests for Cointegration.” *Journal of Applied Econometrics*, 14 (5): 563-577.

Masih, R., and A. Masih, 1996, “Macroeconomic Activity Dynamics and Granger Causality: New Evidence from a Small Developing Economy Based on a Vector Error Correction Modelling Analysis.” *Economic Modelling* 13: 407-426.

Mayer, C., 1988, “New Issues in Corporate Finance.” *European Economic Review* 32: 1167-1188.

Mohtadi, H., and S. Agarwal, 2004, “Stock Market Development and Economic Growth: Evidence from Developing Countries.” http://faculty.apec.umn.edu/mohta 001/PA1-4-01.pdf, 1-18.

Nowbutsing, B., 2009, “Stock Market Development and Economic Growth: The Case of Mauritius.” *International Business & Economic Research Journal* 8: 77-88.

Pagano, M., 1993, “Financial Markets and Growth: An Overview.” *European Economic Review* 37: 613-622.

Phillips, P., and P. Perron, 1988, “Testing for a Unit Root in Time Series Regression.” *Biometrika* 75(2): 335-346.

Romer, P., 1986, “Increasing Returns and Long-run Growth.” *Journal of Political Economics* 94: 1002-1037.

Rousseau, P., and P. Wachtel, 2000, “Equity Markets and Growth: Cross Country Evidence on Timing and Outcomes.” *Journal of Banking and Finance* 24: 1933-1957.

Shahbaz, M., N. Ahmad, and L. Ali, 2008, “Stock Market Development and Economic Growth: Ardl Causality in Pakistan.” *International Research Journal of Finance and Economics* 14: 182-195.

Shahnoushi, A., A. Ebadi, M. Daneshvar, E. Shokri, and M. Motallebi, 2008, "Causality between Financial Development and Economic Growth in Iran." *World Applied Sciences Journal* 4 (5): 736-740.

Singh, A., 1997, "Financial Liberalization, Stock Markets and Economic Growth." *The Economic Journal* 107: 771-782.

Singh, A., and B. Weiss, 1998, "Emerging Stock Markets, Portfolio Capital Flows and Long-term Economic Growth: Micro and Macroeconomic Perspectives." *World Development* 26: 607-622.

Spears, A., 1991, "Financial Development and Economic Growth-Causality Tests." *Atlantic Economic Journal* 19(3): 66-76.

Stiglitz, J., 1985, "Credit Markets and the Control of Capital." *Journal of Money, Credit and Banking* 17(2): 133-152.

Suleiman, B., and S. Aamer, 2008, "Financial Development and Economic Growth: The Egyptian Experience." *Journal of Policy Modelling* 30: 887-898.

Yang, Y., and M. Yi, 2008, "Does Financial Development Cause Economic Growth? Implication for Policy in Korea." *Journal of Policy Modelling* 30: 827-840.