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Free trade agreements and equity market integration: the case of the US and Jordan

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This study is aimed mainly to examine the impact of the US–Jordan Free Trade Agreement (UJFTA) on the degree of equity market's linkage. This issue is carried out through an asymmetric version of the Dynamic Conditional Correlation (DCC) model of Engle (2002) and developed by Sheppard (2002), which allows for asymmetries in both volatilities and conditional correlations. The empirical evidence suggests that the UJFTA has indeed increased substantially and significantly the linkages of the Jordanian capital market with the US equity markets. These results strongly support the argument that the direct trade flows is one of the most important determinant of cross-country linkages in equity markets.

I. Introduction

Given the degree of openness to trade and investment, it is well accepted fact that the international financial markets have become increasingly mutually dependent. When making decisions, traders incorporate information pertaining to price movements and volatility in the assets they are trading including information about related assets. The movement of markets in rhythm and chorus could nullify much of the gain out of diversification across borders, besides being vulnerable to the caprices of global capital (Obstfeld, 1992, 1994; Lewis 1996). As a consequence, clear understanding of the nature of stock markets linkages and interactions is now an essential consideration of investors and policy makers. Thus, increased knowledge of how markets influence one another is important in the determination of pricing, hedging and regulatory policies.

In recent years, globalization of capital flows has led to growing relevance of emerging capital markets and Jordan is one of the countries with an expanding stock market that is increasingly attracting funds from abroad.[1] In particular, deregulation and market liberalization measurers, rapid developments in communication technology and computerized trading system, and increasing activities of multinational corporations have accelerated the growth of Jordanian capital market (Amman Stock Exchange, ASE). From year 2000 onwards, Jordanian firms are raising capital from the US market by listing themselves in US exchanges. At present four Jordanian companies have issued ADRs and are cross-listed in US exchanges and many more companies are planning to cross list in the near future.

Furthermore, in October 2000, Jordan and the US signed a historic Free Trade Agreement (UJFTA) eliminating most duties and commercial barriers to

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[1] According to Amman Stock Exchange, the net investment of foreign investors in the calendar year 2000 is JD 15.5 million, in 2001 it is JD 122.6 million, and in 2002 JD 155.8 million (1 US$ = 0.701 JD).

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bilateral trade in goods and services between the two countries. This agreement rests on the base of an earlier programme known as Qualifying Industrial Zones (QIZs) dating to 1997. The QIZs are industrial parks and are exempted from American tariffs. There are now 40 QIZs, whose production enabled Jordan's exports to the US to grow from $16 million in 1998, when two-way trade was $298 million, to $412 million in 2002 with two-way trade at $817 million, making the US the major trading partner of Jordan. Thus it will be interesting to understand the impact of these developments on the dynamic co-movement of the ASE with US markets.

This study intends to investigate whether, and to what extent, the emerging ASE is integrated with the US markets. It also addresses the issue of whether such a relationship, if it exists, is affected by the UJFTA. The investigation into the role of UJFTA in market integration can have important practical and theoretical implications. In particular, if the ASE has indeed become more integrated with the US markets in the post FTA period, then asset diversification involving the two markets would lose much of its appeal. Moreover, a higher degree of market integration promotes faster adjustments of equity prices to information flows from abroad, leading to a more efficient market. From a theoretical perspective, a higher degree of market integration in the post-UJFTA period suggests that a multinational version of the capital asset pricing model may be a more appropriate model of analysis than the (domestic) version commonly used for the ASE.

The study contributes to the literature in several respects. First, equity market integration is examined using an asymmetric version of Dynamic Conditional Correlation (DCC) model of Engle (2002) and developed by Sheppard (2002) and was used recently by Cappiello *et al.* (2003) to investigate the asymmetric dynamics in correlation of global equity and bond returns. This model estimates the DCC parameters and the time-varying conditional correlation among the returns taking into account the asymmetries' dynamics in the correlation in addition to the asymmetries' response in variance. The estimates of correlations allow us to analyse the significance of any event, such as the UJFTA, that occurred during the period covered by the study. Second, this study provides evidence suggesting that the direct trade flows have important influence on stock markets' integration.

The rest of paper is organized as follows. Section II is devoted to present the review of literature, Section III presents the methodology. Section IV contains the details of the source of data. Section V analyses the results. Section VI concludes.

II. Review of the Related Literature

A number of recent papers have shown that potential for international diversification, of the sort outlined by Levy and Sarnat (1970), Lessard (1973) and Solnik (1976) has diminished. More recent research, such as Kaplanis (1988) and Bekaert and Harvey (1995), has shown that world equity markets have responded to institutional and economic evolutions by becoming more integrated.

Eun and Shim (1989) have argued that greater stock market integration is a natural consequence of greater economic integration, which has gradually taken place. Janakiramanan and Lamba (1998) agree, but suggest that the world's dominant economy is likely to be the driving force behind stock markets elsewhere in the world. They suggest that the stronger the ties with the dominant economy, the more integrated the stock market would be. This is supported empirically by Eun and Shim (1989) who found that the US market heavily influenced most world markets, but no single country had a strong influence on US returns. Frankel and Rose (1998), Glick and Rose (1999), Forbes and Chinn (2003) and Imbs (2003) argued that direct trade flows are the most important determinant of cross-country linkages in stock markets.

During the early 1980s, associated with major developed markets were the extensive liberalization of financial markets, improvements to market surveillance and the removal of impediments associated with foreign investment. Taylor and Tonks (1989) and Chelley-Steeley and Steeley (1999) showed that these initiatives led to an increase in capital mobility and market integration. An accompaniment to these institutional changes that took place were, technological improvements, the widespread cross listing of stocks and a move towards electronic trading systems, which has also increased the links between capital markets. For example, Wu and Su (1998) found that there is significant dynamic relation exist among four major stock markets including the USA, Japan, the UK, and Hong Kong, in addition, the correlation among stock markets has been increased considerably in recent years. However, Byers and Peel (1993) examined the interdependence between stock price indices of the USA, the UK, Japan, Germany and the Netherlands and they did not find any cointegration either for the group as a whole or for the pairs of markets. Furthermore, Kanas (1998) explored the linkages between the USA and six

European stock markets (the UK, Germany, France, Switzerland, Italy and the Netherlands). He found that the US stock market was not pairwise cointegrated with any of the six European stock markets.

Hardouvelis *et al*. (1999) examined the speed of integration among the EU equity markets and found that the degree of integration is closely related to the probability of the country to join the EU. Integration increases substantially over time and seems to be complete by mid-1998. Rangvid (2001) and Aggarwal *et al*. (2003) found evidence of increasing convergence between major EU equity markets since 1982. Tahai *et al*. (2004) investigated the financial integration of G7 equity markets and found evidence of co-movement among these markets.

High levels of co-movement between developed markets has encouraged the flow of portfolio investment to many emerging markets, as investors seek to capitalize upon potential diversification benefits (Bekaert, 1995; Bekaert and Harvey, 1995, 1997; Korajczyk, 1996). However, as international investment increases, the emerging markets become more integrated with world markets (Goetzmann and Jorion, 1999; Syriopoulos, 2004). As such an emerging market may be able to provide significantly enhanced diversification benefits for a relatively short time.

In recent years a large body of literature has investigated the linkages between emerging stock markets and the larger developed markets of the world. Pioneering work by Bailey and Stultz (1990) found that up to 25% of US investors' risk could be reduced, if the stocks of Asian companies were included in their portfolios. In others studies, Cheung and Mak (1992) and Phylaktis and Ravazzolo (2001) examine stock market linkages of a group of Pacific-Basin emerging countries with the USA and Japan. They found that investors have opportunities for portfolio diversification by investing in most of the Pacific-Basin stock markets. However, other empirical studies (such as Chan and Leung, 1989; Lee *et al*. 1990; Chung and Liu, 1994; Janakiramanan and Lamba, 1998), found that some Asian-Pacific countries were segmented from the rest of the world.

However, the focus of a majority of the emerging markets studies is on linkages between the USA and the various emerging stock markets in East Asia. To the present authors' knowledge only very few papers explicitly investigated the ASE's linkages with the overseas markets. Both Darrat *et al*. (2000) and Neaime and Hakim (2002) examined the market linkages in the case of a sample of emerging markets in the Middle East included Jordan. Using cointegration and error-correction techniques, they found that the emerging markets in the Middle East were largely segmented from the US market. While both papers explicitly take into account the long-run linkage, they ignore the time-varying nature of the correlations and the dynamic in second order moments of the data. This study tries empirically to identify the impact of the FTA on the linkages of the ASE with the US equity markets.

III. Econometric Methodology

In recent decades there has been an exponential growth of studies focusing on the time-varying behaviour of correlations and covariance between financial markets. It is now widely accepted that financial volatilities and correlations move together over time across assets and markets. This study therefore examines the time-varying correlation structure between the ASE and the US markets. For this purpose, a recently proposed class of multivariate GARCH models of Engle (2002) were used, called the Dynamic Conditional Correlation model (DCC).

To exemplify Engle's (2002) DCC model for the purpose of this study, let $\mathbf{r_t} \equiv [r_{1t}, r_{2t}]'$ be a 2×1 vector containing the equity market returns series in a conditional equation as:

$$\mathbf{A}(\mathrm{L})\,\mathbf{r_t} = \varepsilon_t, \quad \text{where } \varepsilon_t \sim N(\mathbf{0}, \mathbf{H_t}) \quad \forall\ t = 1, 2, \ldots, T \tag{1}$$

where $\mathbf{A}(\mathrm{L})$ is a polynomial matrix in the lag operator L, and $\varepsilon_t = [\varepsilon_{it}, \varepsilon_{2t}]'$ is a vector of disturbance with a conditional variance–covariance matrix $\mathbf{H_t} \equiv \{h_i\}_t$ for $i = 1, 2$.

The multivariate DCC-GARCH structure can be easily understood by first rewriting the conditional variance–covariance matrix as:

$$\mathbf{H_t} \equiv D_t R_t D_t \tag{2}$$

where D_t is the 2×2 diagonal matrix of time-varying standard deviations from univariate GARCH models with $\sqrt{h_{it}}$ on the 2×2 diagonal and R_t is the (possibly) time-varying correlation matrix. The DCC model is designed to allow for two-stage estimation of the conditional covariance matrix $\mathbf{H_t}$: in the first stage univariate volatility models are fixed for each of the reruns and estimates of h_{it} are obtained; in the second stage market returns, transformed by their estimated standard deviations resulting from first stage, are used to estimate the parameters of the conditional correlations. The original DCC estimator had the dynamics of correlation evolving as a scalar process with a single news impact parameters and a single smoothing parameter.

However, a number of studies document that stocks exhibit asymmetry in the conditional second moments, where volatility increases more after a negative shock than after a positive shock of the same magnitude (see for example, Black, 1976; Christie, 1982; Campbell and Hentschell, 1992).[2] Asymmetric effects have also been recently found in conditional correlations (see for example, Kroner and Ng, 1998; Bekaert and Wu, 2000). Recently, Sheppard (2002) has extended Engle's (2002) DCC-GARCH model to allow for asymmetric dynamics in the correlation in addition to asymmetric response in variance. Cappiello *et al.* (2003) utilized this version to investigate whether, in addition to stocks, government fixed income securities also exhibit asymmetry in conditional second moments and explored the dynamic and changes in the correlations in international asset markets. Following Cappiello *et al.* (2003), the model used in this paper is an asymmetric version of the Engle's (2002) model.

As a first step, the univariate volatility models will be chosen using the Schwartz Information Criterion (BIC) from a class of models that are capable of capturing the common properties of equity returns variance. The models included in the plan search are GARCH of Bollerslev (1986), AVGARCH of Taylor (1986) (GARCH on standard deviations instead of variance), followed by GJR-GARCH of Nelson (1991), ZARCH of Zakonian (1994), EGARCH of Glosten *et al.* (1993) (which all allow for threshold effects but use different powers of the variances in the evolution equation) and APARCH of Ding *et al.* (1993) (which encompass both threshold effects and an estimated power for the variance evolution).

Once the univariate volatility models for both markets are estimated, the standardized residuals for each market, $\varepsilon_{it} = r_{it}/\sqrt{h_{it}}$, are used to estimate the dynamics of the correlation. The standard Engle's (2002) specification of the dynamic correlation structure for the two returns may be written as:

$$R_t = Q_t^{*-1} Q_t Q_t^{*-1} \tag{3}$$

$$Q_t = (1 - a - b)\overline{Q} + a\varepsilon_{t-1}\varepsilon'_{t-1} + bQ_{t-1} \tag{4}$$

where a and b are scalar parameters to capture the effects of previous shocks and previous dynamic conditional correlations. $Q_t \equiv \{q_{ij}\}_t$ is the conditional variance–covariance matrix of the standardized errors with its time-invariant (unconditional) variance–covariance matrix $\overline{Q}$ obtained from the first stage of estimation; and Q_t^* is a diagonal matrix containing the squared root of the diagonal elements of Q_t:

$$Q_t^* = \begin{bmatrix} \sqrt{q_{11}} & 0 \\ 0 & \sqrt{q_{22}} \end{bmatrix}$$

For R_t to be positive definite, the only condition that needs to be satisfied is that Q_t is positive definite, since the elements of the matrix R_t are of the form $\rho_{12,t} = q_{i2,t}/\sqrt{q_{11,t}q_{22,t}}$, where $q_{12,t}$, $q_{11,t}$ and $q_{22,t}$ are the elements of Q_t corresponding to the indices. Equations 2 and 3 are referred to as a DCC (m, n) model.

As Engle's (2002) model does not allow for asymmetries, Sheppard (2002) modified the evolution equation to be:

$$\begin{aligned} Q_t = {} & (\overline{Q} - A'\overline{Q}A - B'\overline{Q}B - G'\overline{N}G) \\ & + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'Q_{t-1}B + G'n_{t-1}n'_{t-1}G \end{aligned} \tag{5}$$

where A, B, G are 2×2 diagonal matrices, $I[\cdot]$ is an indicator function and $n_t = I[\varepsilon_t < 0] * \varepsilon_t$($*$ denotes the Hadamard product, i.e. element-by-element multiplication). The matrix $\overline{N}$ is given by $\overline{N} = E[n_t n'_t]$ for $t = 1, \ldots, T$. In the estimation procedure $\overline{Q}$ and $\overline{N}$ are replaced with sample analogues $T^{-1}\sum_{t=1}^{T} \varepsilon_t\varepsilon'_t$ and $T^{-1}\sum_{t=1}^{T} n_t n'_t$. A sufficient though not strictly necessary condition for Q_t to be positive definite is that $(\overline{Q} - A'\overline{Q}A - B'\overline{Q}B - G'\overline{N}G)$ is positive definite.

Four special cases of the above model (Equation 5) exist. These models can be retrieved by imposing restrictions on the parameter matrices A, B and G in Equation 5, see also Engle (2002) and Cappiello *et al.* (2003). For the ease of discussion the restrictions on Equation 5 for each of the models will be given in the following way for $i = 1, 2$:

- Model I: the standard DCC model. This model is given in Equation 4 and is obtained by the restrictions $(A)_{ii} = \sqrt{a}, (B)_{ii} = \sqrt{b}$ and $(G)_{ii} = 0$, where a and b are the corresponding parameters in Equation 4.
- Model II: The generalized symmetric DCC model. The restriction that is needed is $(G)_{ii} = 0$. This yields the representation $Q_t = (\overline{Q} - A'\overline{Q}A - B'\overline{Q}B) + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'Q_{t-1}B$.

[2] The explanation Black (1976) put forth is the leverage effect hypothesis: a drop in price of stock (negative return) increases the value of equity and hence increases the debt-to-equity ratio, which makes the stock riskier and increases its volatility. Another explanation of the leverage effect is volatility feedback. When positive shocks to volatility increase future risk premia and hence drive down current prices, there is a negative contemporaneous correlation between stock prices and unexpected changes in volatility, as reported in Campbell and Hentschell (1992) for index return.

- Model III. The Asymmetric standard DCC model: the restrictions are $(A)_{ii} = \sqrt{a}$, $(B)_{ii} = \sqrt{b}$ and $(G)_{ii} = \sqrt{g}$, where a, b are the correlation parameters in Equation 4 and g is the asymmetry parameter that is the same for each stock. This yields the representation $Q_t = \quad (1-a-b)\overline{Q} - g\overline{N} + a\varepsilon_{t-1}\varepsilon'_{t-1} + bQ_{t-1} + gn_{t-1}n'_{t-1}$.
- Model IV: The generalized asymmetric DCC model as given in Equation 5.

Following Cappiello *et al*. (2003) the DCC model is also extended to allow for possible structural breaks in mean or dynamics. For example, let d be 0 or 1, depending on whether $t > \tau > T$. Then to investigate whether a structural break has occurred, model 5 can be modified as:

$$Q_t = (\overline{Q} - d\tilde{Q} - A'\overline{Q}A + dA'\tilde{Q}A - B'\overline{Q}B + dB'\overline{Q}B - G'\overline{N}G + dG'\tilde{N}G) + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'Q_{t-1}B + G'n_{t-1}n'_{t-1}G \tag{6}$$

where $\overline{Q} = E[\varepsilon_t\varepsilon'_t]$, $t < \tau$, and $\tilde{Q} = \overline{Q} - E[\varepsilon_t\varepsilon'_t]$, $t \geq \tau$, with $\overline{N}$ and $\tilde{N}$ analogously defined, which is corresponding to the following parameterization when mean reversion is enforced

$$Q_t = (\overline{Q}_1 - d\tilde{Q}_1 - A'\overline{Q}_1A + G'\overline{N}_1G) + (\overline{Q}_2 - A'\overline{Q}_2A - B'\overline{Q}_2B - G'\overline{N}_2G) + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'Q_{t-1}B + G'n_{t-1}n'_{t-1}G \tag{7}$$

where $\overline{Q}_1 = E[e_te'_t]$, $t < \tau$ and $Q'_2 = E[e_te'_t]$, $t \geq \tau$. As model 6 nests the standard model (Equation 4), it is a straightforward test for breaks in the mean of the process. The test can be conducted using standard log-likelihood ratio tests (LR) with $k(k-1)/2d \cdot f$. Breaks in dynamics as well as breaks in both dynamics and mean can be tested for analogously.

The estimation of the model can be performed by the quasi-maximum likelihood (QML) method which maximizes the criterion function:

$$L(\theta) = -\frac{1}{2}\sum\nolimits_{t=1}^{T} (\log|H_t(\theta)| + r'_tH_t^{-1}(\theta)r_t) \tag{8}$$

with respect to the parameter vector θ. Under quite general conditions, listed by Engle and Sheppard (2001), these estimators will be consistent and asymptotically normal. If the estimation for the variance (contained in D_t) and the correlations (contained in R_t) is performed simultaneously, the QML estimation is efficient providing the innovations that are indeed Gaussian. However, if the estimation is split up in two parts, where first the variances are estimated, and then the correlations, then the estimators will no longer be efficient but still consistent. Following Engle (2002), the likelihood can be split in two parts:

$$L(\theta) = L_v(\theta_v) + L_c(\theta_c)$$

where

$$L_v(\theta_v) = -\frac{1}{2}\sum\nolimits_{t=1}^{2} (\log|D_t(\theta_V)|^2 + r'_tD_t(\theta_V)^{-2}r_t \tag{9}$$

is the volatility part of the likelihood, and

$$L_C(\theta_C) = -\frac{1}{2}\sum\nolimits_{t=1}^{T} (\log|R_t(\theta_C)| + \varepsilon'_tR_t(\theta_C)^{-1}\varepsilon_t \tag{10}$$

is the correlation part, with $\theta = (\theta'_V, \theta'_C)'$. At the first step, Equation 9 is maximized with respect to θ_V by estimating the univariate GARCH models for r_{it}, $i = 1, 2$. Define the estimate of θ_V by $\hat{\theta}_V = \arg\max L_V(\theta_V)$. Conditional on the first step, standardized residuals can be calculated. At the second step, Equation 10 is maximized with respect to θ_C, giving the estimate $\hat{\theta}_C = \arg \max L_C(\theta_C)$. This estimation procedure and the theoretical and empirical properties of the estimator are extensively discussed in Engle and Sheppard (2001).

IV. Data

Regarding the purpose of this study, using daily closing prices will lead to synchronization problems due to different trading hours of the markets. These synchronization problems cannot be ignored, since they will lead to a downward bias in the estimated correlations. Martens and Poon (2001) show that applying synchronization to non-synchronous daily data gives a significant downward bias in correlation, as compared to *pseudo*-closes (i.e. sampling all prices at the same time (GMT)). To keep away from synchronization problems many studies use weekly data instead of daily data.[3] Using weekly rather than daily data can also avoid the potential biases associated with non-trading, the bid-ask spread effect in daily data (Lo and MacKinlay, 1998; Hung and Cheung, 1995) and the problem of thin trading which is often associated with most emerging markets including the ASE.

Following these studies weekly prices of the ASE index and the S&P 500 index are used for the time period spanning January 1987 to May 2004.

[3] See for example, Cappiello *et al*. (2003) and Longin and Solnik (1995).

Table 1. Descriptive statistics and preliminary statistics

	Mean	Max	Min	Standard deviation	Skewness	Kurtosis	Standardized skewness	Standardizd kurtosis
Panel A: descriptive statistics								
S&P 500	0.1186	7.492	−11.749	2.086	−0.4473	5.687	−0.3707	3.8104
ASE	0.1791	11.150	−7.887	2.521	−0.7728	14.323	−0.8342	16.096
Panel B: markets returns correlation matrix								
	S&P		ASE					
S&P 500	1.000		0.1853**					
ASE	0.1853**		1.000					
Panel 1C: autocorrelation of returns								
	$AC(1)$	$AC(2)$		$AC(3)$	$AC(4)$	$AC(5)$	$AC(6)$	$Q(6)$
S&P 500	−0.032	0.062		−0.035**	−0.026	−0.033	0.061	8.997
ASE	0.027	0.005		0.021	−0.041	−0.058	−0.045	6.531
Panel 2C: autocorrelation of squared returns								
	$AC(1)$	$AC(2)$		$AC(3)$	$AC(4)$	$AC(5)$	$AC(6)$	$Q^2(6)$
S&P 500	0.255*	0.066*		0.107*	0.074*	0.046*	0.092*	77.286*
ASE	0.125*	0.017*		0.106*	0.034*	−0.001*	0.007*	22.379*
Panel E: test for asymmetry (χ^2_m)								
	asymmetry in variance			asymmetry in correlation				
S&P 500	42.643***			22.360***				
ASE	17.904***			12.587**				

Notes: The mean is the equally weighted average of observation over the sample period. JB is the Jarque–Bera (1980) normality test, which follows a Chi-squared distribution, with two degrees of freedom. $AC(1)$ to $AC(6)$ denotes autocorrelation for return series. Q and Q^2 denote respectively the Ljung–Box and Ljung–Box^2 statistics. The Wald test statistic is asymmetrically distributed as (χ^2_m), with m indicating the degree of freedom. *, ** and *** indicate a significance at the 10%, 5% and 1% levels, respectively.

The ASE index is a value-weighted index, which currently has 70 stocks. The stocks included in the index represent around 90% of the aggregate market capitalization of the listed companies. The data are obtained from the ASE database. For the US data, Yahoo's website (http://www.quote.yahoo.com) provides the necessary weekly figures on S&P 500. Following Koutmos (1996) and De Santis and Imrohoroglu (1997), returns are measured in home-country currencies to incorporate hedging activities of investors against foreign exchange-rate risk. In the empirical studies below, returns are expressed in first differences of log prices to approximate continuously compounded returns.

The summary statistics of the data are given in Table 1. The average weekly return of ASE index during the period under consideration was 0.18% as compared with the weekly return of 0.12% for the S&P 500 index. On the other hand the ASE index returns showed a higher standard deviation of 2.5% as compared to the 2.1% standard deviation of the S&P 500 index. The two equity indices returns were negatively skewed, which indicates a long left tail in the empirical distributions and supports the idea that these series have asymmetric distributions. The raw equity index returns also exhibits extreme excess kurtosis. It is well known that while heteroscedastic returns can exhibit skewness and fat-tails, returns standardized by their estimated conditional standard deviation can be normal (or close to normal). To examine the properties of the innovations, the residuals were standardized by the favoured GARCH model. While the residuals standardized by their estimated standard deviation are both less skewed and less fat-tailed, the standardized residuals are highly non-normal. In fact, both equity indices returns, even when standardized by an estimated conditional standard deviation, reject normality using a Jarque–Bera test at the 1% level. The unconditional correlations between the two equity index return series are also reported in Panel B of Table 1. The ASE and the S&P 500 were reasonably unconditionally correlated with a correlation coefficient of 0.1853. In Panel 1C of Table 1 autocorrelation functions were reported for lags from 1 to 6. As shown, both equity markets returns exhibit no serial correlations. A result confirmed by the Ljung–Box statistics.

While equity market returns in levels show no autocorrelations, squared returns exhibit a second order dependence that a GARCH process should be able to capture. As shown in Panel 2C, serial

Table 2. Model selected and parameters estimates for the univariate GARCH models used to standardize each return series

	Model selected	ϖ	α	γ	β	
Panel A: estimation results of univariate asymmetric GARCH (1,1) models						
S&P 500	EGARCH	−0.0766*	0.1701*	−0.0468*	0.9337*	
ASE	GJR-GARCH	0.3108*	0.0585*	−0.0446*	0.8867*	
Panel B: normality test for the standardized residuals						
	Skewness		Kurtosis		Jarque–Bera	
S&P 500	−0.038		3.523		12.518**	
ASE	−0.098		3.145		9.250**	
Panel 1C: autocorrelation test of standardized residuals						
	$AC(1)$	$AC(2)$	$AC(3)$	$AC(4)$	$AC(5)$	$AC(6)$	$Q(6)$
S&P 500	−0.097	0.076	0.000	−0.034	−0.048	0.075	5.328
ASE	0.042	0.018	0.021	−0.018	−0.049	−0.036	5.163
Panel 2C: autocorrelation test of squared standardized residuals							
	$AC(1)$	$AC(2)$	$AC(3)$	$AC(4)$	$AC(5)$	$AC(6)$	$Q^2(6)$
S&P 500	0.036	−0.039	0.038	−0.033	−0.046	0.000	5.965
ASE	0.016	−0.003	0.056	−0.004	−0.014	−0.006	2.940

Notes: *, ** and *** indicate a significant at the 10, 5 and 1% levels, respectively.

$$\text{EGARCH model:} \quad \log(h_t) = \varpi + \alpha\frac{|\varepsilon_{t-1}|}{\sqrt{h_{t-1}}} + \gamma\frac{\varepsilon_{t-1}}{\sqrt{h_{t-1}}} + \beta\log(h_{t-1})$$

$$\text{GJR} - \text{GARCH:} \quad h_t = \varpi + \alpha\varepsilon_{t-1}^2 + \gamma[\varepsilon_{t-1} < 0]\varepsilon_{t-1}^2 + \beta h_{t-1}$$

correlation is pronounced for both the equity market returns series. The null hypothesis of no simultaneous autocorrelation up to six lags is nevertheless rejected for both variables at the 1% significance level.

Finally to capture asymmetries in variances and correlations, the study examines if the variances of equity returns are higher after a negative shock than after a positive one. To explore asymmetries in variances, following Cappiello *et al.* (2003), $E[r_{it}^2/r_{it-1} < 0]$ is calculated and then the null hypothesis that $E[r_{it}^2/r_{it-1} < 0] = E[r_{it}^2/r_{it-1} > 0]$ is tested. If there were an asymmetric increase in the level of variance after a negative shock, one would expect to find that $(\sum_{t=2}^{T} I_{r_{it-1}<0})^{-1}\sum_{t=2}^{T} r_{it}^2 I_{r_{it-1}<0} - (\sum_{t=2}^{T} I_{r_{it-1}>0})^{-1}\sum_{t=2}^{T} r_{it}^2 I_{r_{it-1}>0} > 0$. Following the same line of process, one can examine if the average covariance of the standard residuals (conditional correlation, $\rho_{ij,t}$) after combined negative returns is different than after two positive returns by testing $(\sum_{t=2}^{T} I_{\varepsilon_{it-1}<0} I_{\varepsilon_{jt-1}<0})^{-1}\sum_{t=2}^{T} \varepsilon_{it}\varepsilon_{jt} I_{\varepsilon_{it-1}<0} - (\sum_{t=2}^{T} I_{\varepsilon_{it-1}>0} I_{\varepsilon_{jit-1}>0})^{-1}\sum_{t=2}^{T} \varepsilon_{it}\varepsilon_{jt} I_{\varepsilon_{it-1}>0} I_{\varepsilon_{jit-1}>0} > 0$.

Panel E contains the Wald test statistics of the asymmetric effects. The results show that both equity indices have significantly greater variances after a negative shock than after positive shocks. Likewise, the tests suggest that both equities exhibited significant increases to joint bad news (two negative returns) after joint good news (two positive returns).

V. Empirical Results

This section presents the empirical results of the asymmetric standard DCC model. In the first step the univariate GARCH models for each market are fitted and the best one selected using information criteria. Table 2 contains the specification of the GARCH process selected by the BIC and the estimated parameters from these models.[4] The results imply that both markets' returns contain significant asymmetry terms. This is consistence with the earlier evidence of large conditional difference in variance after negative shocks.

Using the standardized residuals from the first step, the study continues with the second step of the estimation procedure for the asymmetric DCC model. Model IV is estimated for the dynamics of conditional correlation between the S&P 500 and ASE indices returns. This model is a full model where asymmetric terms were induced allowing for different news impact and smoothing parameters across the returns series. The estimation results of the model are given in Table 3. From this table

[4] The tests of significance are computed with the robust standard errors of Bollerslev and Wooldridge (1992).

Table 3. Asymmetric DCC estimates

	A	A^2	B	B^2	G	G^2
S&P 500	0.123**	0.003	0.892*	0.846*	0.259*	0.193**
ASE	0.217**	0.075*	0.944*	0.907*	0.126*	0.092*

Notes: * and ** indicate significance at the 10% and 5% levels, respectively.

one can see that the asymmetry parameter g is significant at the 5% level.

Significant evidence was also found of a structural break since the UJFTA was signed. Both a structure break in the mean and a structural break in both the mean and the dynamics were tested for. The log-likelihood ratios reject the null hypothesis of no structural break in the mean (LR $= -326.1428$). Furthermore, evidence for both a break in the mean and the dynamics (LR $= -825.2778$) is found. Therefore, the remainder of the paper will present the DCC model with a break in the mean and in the dynamics.

While each of the volatility series were assumed to evolve independently of the other series, the model allows one to examine the volatility linkages between the two markets. A simple measure to examine this linkage is the correlation coefficient of the estimated conditional volatility between the two markets.[5] The average correlation between conditional volatility of the two equity markets during the period under investigation was 0.4015. Furthermore, evidence is found supportive to the presence of a stronger link between the ASE and the US market following the UJFTA. This inference was based on the observation that the correlation coefficient between the estimated conditional volatility of the two markets during the period following the UJFTA more than tripled, rising from 0.1495 in the period prior to the UJFTA, to almost 0.6915 thereafter.

Figure 1 contains a plot of the annualized average volatility series for the two markets. Consistent with the previous evidence, the correlation between the volatility of the two countries has clearly increased since October 2000, and this volatility linkage was almost evident during the certain tumultuous periods: terrorist attacks hit the US in September 2001, and the US invasion of Iraq in March 2003. One is tempted to conclude from assessing this correlation

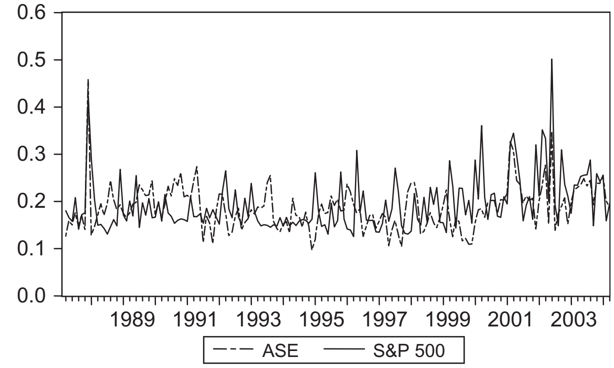

Fig. 1. Annualized conditional volatility

Table 4. Likelihood ratio statistics of time-varying conditional volatility spillovers from US market to the ASE (Dependent variable: time—varying conditional return volatility of the ASE)

All period	Pre-UJFTA period	Post-UJFTA period
2.026	1.079	4.616**

Note: ** indicates a significance at the 5% level.

that the two markets have become more integrated in the post-UJFTA period.

Of course, analysis of contemporaneous volatility correlations between the two markets, although useful for measuring the degree of associations, do not reveal pertinent information on the underlying causal process; the latter requires some evidence on lead-lag relations. To that end, one uses the Granger concept of causality in the context of multivariate volatility model. This mode expresses the conditional based volatility of the ASE as a function of the lagged conditional based volatility of the US market, besides its own lagged innovations. To assess the impact of the UJFTA on volatility spillovers from the US market to the ASE market, the above model is estimated over the periods before and after the UJFTA.

Table 4 reports the likelihood ratio (LR) test statistics for the significance of causal effects from volatility in the US market to volatility in the ASE. These results clearly suggest that volatility spillovers from the US market to the ASE became significant only after the UJFTA. In fact, prior to the UJFTA,

[5] Correlation coefficient between the estimated conditional volatility is calculated as follows:

$$\rho_{ht\,h_{jt}} = \sum_{i=1}^{T} (h_{it} - \bar{h}_i)(h_{jt} - \bar{h}_j) / \sqrt{\sum_{t=1}^{T} (h_{it} - \bar{h}_i)^2 \sum_{t=1}^{T} (h_{jt} - \bar{h}_j)^2}$$

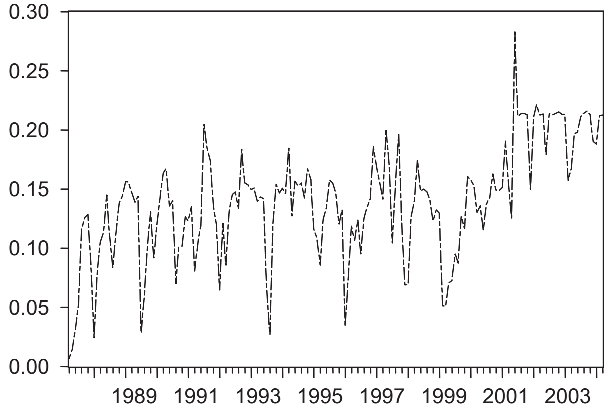

Fig. 2. Conditional correlation of ASE with S&P 500 index

the ASE appeared quite segmented from the US market, where volatility spillovers from this market were statistically non-existent. However, the situation has dramatically changed after the UJFTA. As result in the table, volatility in the ASE in the post-UJFTA period significantly responds to volatility changes in the US market at better than the 5% level. These results provide another testimony supportive of the evidence regarding the ASE linkages with the US market in the post-UJFTA period. Also, provide clear evidence of the presence of significant volatility spillovers from the USA to the Jordanian market.

Interesting empirical observation about volatility not withstanding, the primary motivation of this study was to look at the correlation dynamics between the ASE and the US markets returns. There appear to be significant variations in the correlations between the US and ASE markets during the period of the sample. Figure 2 contains a graph of estimated dynamic market correlation between the two markets. The correlation has clearly increased between the two markets since the UJFTA. Particularly, the correlation has increased considerably from 0.1246 pre-UJFTA period to 0.2304 post-UJFTA period. The increase in correlation is so striking that not only is a mean change obvious, but also correlations appear to be less volatile after the UJFTA. Clearly, these results provide a piece of evidence supportive to the significant role of UJFTA in strengthening equity market linkages of the ASE with the US market.

VI. Conclusion

The main purpose of this study is to investigate whether, and to what extent, the emerging ASE is integrated with the US market. It also addresses the issue of whether such a relationship, if it exists, is affected by the UJFAT. Since its inception in October 2000, the agreement has sought to strengthen the linkages of the ASE with the US market which, if successful, would lessen the appeal of asset diversification across the two markets and promote higher degree of market efficiency in the ASE.

The analysis was carried out using an asymmetric version of the DCC model of Engle (2002) and developed by Sheppard (2002). This model is particularly well suited to examine correlation dynamics among assets, allowing for asymmetries in the correlation as well as the asymmetric response in variances. As expected, the empirical results show that stock market returns exhibit asymmetry in both conditional correlation and volatility.

Taking into account the structural break which was found in the level of conditional correlation as well as in the level of the conditional volatilities, the empirical evidence unambiguously suggests that the UJFTA indeed increased the degree of the linkages of the ASE with the US markets. Particularly, it was observed that volatility correlations and conditional correlations increased substantially and significantly post-UJFTA period. Furthermore, in the post-UJFTA period, the conditional equity correlations were also found to increase dramatically when bad news, such as the September 2001 attack and the US invasion of Iraq in March 2003, hit financial markets. This is an important implication for international investors; diversification sought by investing in the markets is likely to be less when it is most desirable.

Besides the implication for asset allocation and market efficiency, the results of this study support recent attempts in the literature that internationalize the traditional CAPM (e.g. Bekaert and Harvey, 1995; De Santis and Gerard, 1997; De Santis and Imrohoroglu, 1997). Furthermore, the results confirm the argument that direct trade flows are the most important determinant of cross-country linkages in stock markets.

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