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Exploring the Stock Price Correspondence to Oil Price Shocks In the Gulf Cooperation Council Countries: Evidence from Linear (Symmetric) Model

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ABSTRACT

This study investigates the linkage between oil price index (OPI) and stock price index (SPI) in six Gulf Cooperation Council (GCC) countries in two folds. Firstly, it studies the long-run relationship linking the SPI and the OPI for the time span, beginning February 2002 to May 2015. After confirming the dependency of the SPI across cross-sectional units for the six GCC countries three types of panel cointegration tests were used. Pedroni; Kao which is an Engle-Granger two step residual based test, and Fisher which is a combined Johansen test. Secondly, it investigates the linear short-run effect of OPI shocks on these markets by using bootstrapped resample residuals. The findings reveal a robust long-run relationship amid OPI and SPI of six GCC country members. Furthermore, the linear short-run results indicate significant and positive consequence on Oman, Qatar, and UAE of OPI shocks. The findings also indicate that Qatar is the most oriented market for the oil price changes within a short time period.

Keywords: Stock Price, Oil Price, Panel Cointegration, Bootstrapped Resample Residuals, Linear Short-run Model
JEL Classifications: C65, C33, G14

1. INTRODUCTION

Oil plays a significant role in the economy. Its prices have been subject to drastic increases and decreases over the last years. These fluctuations in oil prices influence the economic activity of both emerging countries and developed. The Gulf Cooperation Council (GCC) is comprised of the leading oil-exporting countries. One of the economic features of these countries is their sensitivity to oil price fluctuations, which stem from their heavy dependency on oil and the lack of diversification in their respective economies. This feature is a potential source of interest, to study these markets from different perspectives, especially the association connecting stock markets with oil price. Investors possibly need such studies to rationalize their investment assessments and policy-makers may need them to regulate stock price markets.

In addition, another issue that drew the attention of investors and decision makers is whether the stock price reacts asymmetrically to the oil prices shock. Detecting asymmetry in long-term equilibrium remains a valuable asset that investors, authorities and companies can use to manage their portfolios and strategies, minimizing the risk of exposure to oil prices (Rafailidis and Katrakilidis, 2014). Pioneer in Sadorsky (1999), there is a number of empirical studies suggesting that aggregate prices on the stock exchange and/or stock prices of listed oil and gas companies may be asymmetrically responsive to changes in oil prices. Therefore, the question of the effect of heterogeneity, which is not visible in Filis et al. (2011) and Jammazi and Aloui (2012), this is particularly important in a generalized manner, because there can be no differential effects of oil prices on equities markets in countries in the same group (within group differences) and possible dichotomies between the two groups. Such generalizations may be needed to understand the international dichotomy (if any) between the export of oil and the import of oil in relation to the oil stock.

More evidence for cooperation between the oil price movement and the stock price (SPI) is an important attraction for financial analysts, investors and decision makers who are interested in the dynamics of the stock market. Literature has suggested various channels to help oil prices influence stock prices. If there is an unprecedented increase in oil prices, energy costs for many companies will likely increase. As a result, the profits may decrease and thus the current cash flow. Although the actual value of the stock depends on future cash flows, equity investors and analysts anticipate further increases in crude oil prices and estimated future cash flows are lower in stock evaluations. But the fact that these future cash flows will react differently to positive innovation and the negative oil price means that the effect of oil price shocks on stock prices should also depend on the nature of the asymmetric shock in terms of size and shock signal.

Meanwhile, we have estimated the predictability of symmetric model based on the expected yield of the sample, which is additionally justified by the importance in this modeling symmetry of oil stocks. We have considered estimated periods beyond the sample and different projection horizons for robustness. This factor is particularly important for analysts and investors in the financial market, who regularly develop forecasts of possible ways to track stock prices in the face of shocks such as oil price.

After this section the structure of the rest of the article is as follows. Section 2 contains a brief presentation of the literature and a contribution to the study, while Section 3 explains the oil price and price index data. In section 4, we describe the methodology used for estimating and performing preliminary analyzes; present a dynamic diagram of heterogeneous data groups, including the estimation procedure. Discussing results, including diagnosis tests, and strategic implications, is provided in Section 5. Section 6 conclusion and policy implications.

2. BRIEF REVIEW OF THE LITERATURE AND CONTRIBUTION

Previous literature has focused on the association of stock markets, in developed and developing countries, with oil prices. Sadorsky (1999) noted that increases in oil price negatively affected U.S. stock market activities. Implementation of a multivariate vector-autoregression (VAR) method, Papapetrou (2001) conducted a study on Greece. Papapetrou affirmed significant association between oil price alterations and economy, employment and stock price markets. The same model was implemented by Park and Ratti (2008) as an attempt to check-out the influence of shocks of oil prices on stock exchange returns in thirteen European countries and the USA. Results showed that shocks of oil prices strongly affected all the selected stock returns apart from USA.

Existence has been confirmed by Chiou and Lee (2009) adopting autoregressive conditional jump intensity model of a significant negative association amid fluctuations of US stock market and oil price shocks. In a study done by Apergis and Miller (2009) on a sample consisting of eight major countries, significant results were found but oil market shocks had a small magnitude of force on international stock market returns.

Aloui and Jammazi (2009) found that a significance aspect in determination of the volatility of stock returns in particular countries such as UK, Japan and France is played by escalation of oil prices. A non-positive correlation in a study implemented on six OECD countries was found by Miller and Ratti (2009) amidst stock returns and oil prices. Narayan and Narayan (2010) regard the Vietnamese stock prices showed compelling positive linkage with oil prices. Nguyen and Bhatti (2012) regard the Chinese stock market found no relationship among stocks returns and fluctuations in international oil prices.

Never the less academic literature covering the topic of stock markets and deviation of oil prices noticeably mentioned the GCC members. For example a study conducted on all GCC countries except Qatar by Hammoudeh and Aleisa (2004) using a VAR model found that only the stock market index in Saudi Arabia is correlated and heavily depends on oil prices. Again Basher and Sadorsky (2006) applied the same analysis on stock markets of GCC members and found solely in Saudi Arabia and Oman markets influence of a projecting fashion of increases in oil prices is evident. Maghyereh and Al-Kandari (2007) results showed a non-linear way oil prices possess direct force on indices of stock prices. Arouri and Fouquau (2009) focused on the short-run association in the midst of GCC stock markets and oil prices. Their study indicated that in Qatar, Oman, and UAE a link of alteration in oil prices exists along with selected stock market returns. On the contrary, such force of oil price deviation was not evident on returns of Bahrain, Kuwait, and Saudi Arabia stock market.

Salisu and Isah (2017) advocated a positive long term and short term relationship between oil prices and stock prices for both oil exporting and oil importing countries via applying panel ARDL model. You et al. (2017) confirmed that oil price shocks and economic uncertainty are asymmetric and highly related to stock market condition. Also, before the crises, rising the oil prices have greater negative effect on stock return. Nusair and Alkhasawneh (2017) stated that rising oil prices only increase stock return when stock markets are highly quintiles and medium quintile and the failing of oil prices only lower stock returns when stock markets are low quintile and medium quintile. Antonakakis et al. (2017) found that, for both oil exporting and oil importing countries, volatility suggest that connectedness varies across different time periods. This time varying is stand with certain development in the global economy.

Al Janabi et al. (2010) showed evidence pointing towards essence of long-run relationship amid alteration of oil prices and stock markets after implementing their study on oil-stock nexus for GCC members Mohanty et al. (2011) also confirmed the attendance of this long-run association. Moreover Mohanty et al. (2011) indicated that returns of GCC stock markets were directly impacted by fluctuation in oil prices at both country and industry levels.

Arouri and Rault (2012) checked the long term association within stock markets of GCC members and oil price deviations. They applied an advanced analysis technique called bootstrap panel cointegration and seemingly unrelated regression (SUR). This study's findings revealed that a positive correlation among stock prices of GCC member countries and the oil prices exists with the exception to Saudi Arabian stock market. The SUR test results gave indication that escalation of oil price has a positive influence on stock prices and again with the exception of Saudi Arabian stock market.

Unlike the previous studies, this study employs a panel cointegration technique after clarifying the dependency a cross cross-sectional units for GCC countries. The use of this technique enables the inspection of long-run association among OPI and SPI. This study also distinguished the liner individual deterministic short-run shocks of OPI. The model is also very advantageous as it uses bootstrapped resample residuals characterized by deriving robust estimates of standard errors and confidence interval of the parameter with no outliers, as good as normally distributed. This is sufficient for measuring the linearity shocks that might affect the stock market's performance during short time period. In this case, it is supposed that stock markets are underperforming if they respond to sudden shocks of oil price, while stock markets are top performing if they do not respond to sudden shocks of oil price.

2.1. Data

The study aims to look-into the long-run and short-run association amid the logarithm SPI (LLSPI) and the logarithm OPI (LOPI) in GCC countries. The intended time span begins on February 2002 to May 2015 for the six GCC countries. The indices were retrieved from the MSCI world market index and international financial statistics, descriptive statistics for SPI and OPI are outlined in Table 1.

From Table 1 the results show that the six GCC countries are within the same mean interval, which insures the common characteristics between the GCC countries. Meanwhile, Oman and UAE SPIs are more volatile; (highest standard deviations of 0.818 and 0.961 respectively) compared to Saudi Arabia, Bahrain, Qatar and Kuwait. Jarque-Bera normality null hypothesis can be rejected in all GCC countries, except for Saudi Arabia**.** Skewness is negative for all cases, which indicates that all GCC countries have the same asymmetry of the probability distribution in their mean.

3. METHODOLOGY

This preceding segment illustrates the research methodology conducted. It spells out the research design, study area, techniques and combination of methods, data collection, processing and analysis as well as data presentation methods. The study seeks to investigate long-run "co-movement" cointegration between SPIs (LSPI, T = 161), and LOPI for the cross country. Hence, in order to verify the long-run relationship "co-movement" the cross section independence assumption in panel data should be tested. The assumption can be verified by (Pesaran, 2004) CD and LM test as follows:

$$CD=\sqrt{\frac{2}{N(N-1)}}\left\{\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\sqrt{T_{ij}}\,\hat{}\,\rho_{ij}\right\}$$

Where, N denoted the number of SPIs for the cross-country (6), T_{ij} donated the number of observations for which the correlation coefficient for the cross-country, ρ_{ij} donated the par-wise correlation coefficient involving the SPIs i and j. The null hypothesis is the cross-section independence. We apply LM test to verify the independence assumption in the panel data heterogeneous that has small cross sectional unit (N, 6 GCC countries) and large number of time period (T) from February 2002 to May 2015. The formula is expressed as follow:

$$LM_{BP}=T\sum_{i=1}^{n-1}\sum_{j=i+1}^{n}\rho_{ij}^{2}$$

Where, T denoted total observations for the cross country, ρ_{ij} donated the correlation coefficient amid the SPIs i and j. Cross-section independence reflects the null hypothesis is, H_0: $\rho_{ij} = 0$ for i≠j. Therefore, rejection of null hypothesis may be because of heteroskedasticity or cross-sectional dependence or the joint of both. This indicates the variance-covariance matrix is proportional to identity matrix. After verifying the cross-sectional independence assumption, the null hypothesis is rejected. This proves nonstationary for the cross-sectional of different countries which tend to be not contemporaneously correlated. Hence, second generation unit root test Cross-sectional Augmented Dickey-Fuller (CADF) regressions should be applied, (Pesaran, 2007). Given by

$$\Delta Y_{it}=\alpha_i+\pi_i\ Y_{i,t-1}+C_{1i}\ \bar{Y}_t+C_{2i}\ \Delta\bar{Y}_{t+\varepsilon_{it}}$$

Where $\bar{Y}_t = N^{-1}\sum_i y_{it}$, and C_{1i}, C_{2i} are nuisance parameters. In order to test the unit root hypothesis which is $\pi_i = 0$ for all i, the null hypothesis assumes all the variables are non-stationary while the alternative or substitute is stationary for at least a single time series. However, the average of the N individual CADF t-statistics on π_i is a combination of p-values of the individual cross-section, with additional lags of ΔY_{it} and $\Delta\bar{Y}_t$ that capture the serial

Table 1: Basic descriptive statistics

Descriptive	Bahrain	Kuwait	Oman	Qatar	UAE	Saudi.A	LOPI
Mean	4.479	4.567	4.623	4.553	3.718	4.670	7.531
Minimum	3.068	3.225	4.341	3.103	1.741	3.636	6.802
Maximum	5.281	5.400	6.340	5.224	5.439	5.721	8.076
Standard deviation	0.572	0.474	0.818	0.499	0.961	0.452	0.311
Skewness	−0.892	−0.955	−0.617	−1.263	−0.794	−0.442	−0.893
Kurtosis	2.881	4.038	2.569	3.997	2.503	3.206	2.859
JB	21.448*	31.75*	11.46*	49.21*	18.57*	5.541	21.57*

JB is the Jarque-Bera test for normality based on excess Skewness and Kurtosis. *, and ** indicate significance at the 1%, and 5% level, respectively. Kurtosis has been normalized to zero correlation. Furthermore, the stationary of the panel data is checked by applying the common unit root process Levin et al. (2002) presume the continuous parameters are common across cross-sectional units, while the continuous parameters freely move across cross-sectional units. The individual unit root should be applied, Ima et al. (1997), ADF-Fisher, and PP-Fisher tests are founded on this form. Levin et al. (2002), common unit root process is expressed as follow:

$$\Delta Y_{it} = \delta Y_{it-1} + \varepsilon_{it}$$

Where, Y_{it} donated the stochastic process for a panel individual i = 1, 2,….,N and each cross-sectional unit contain t = 1,2,…,T time series. To determine the integrated of Y_{it} for each individual cross-sectional units of the panel. The residual ε_{it} is distributed independently across cross-sectional units and follow ARMA process for each cross-sectional unit.

Im et al. (2003), distinguishes a separate ADF for each cross-sectional unit with separate effect without time trend. Expressed preceding finite-order AR(P_i+1) processes:

$$Y_{it} = \mu_i \theta_i (1) + \sum_{j=1}^{p_i+1} \theta_{ij} Y_{i,t-j} + \varepsilon_{it}$$

Where, $\theta_i(1) = 1 - \sum_{j=1}^{p_i+1} \theta_{ij}$, i = 1,…,N and t = 1,…,T. Null hypothesis assume variables are non-stationary $H_0{:}\mu_i\theta_i(1) = 0$ for all cross-sectional units i. while the substitute hypothesis $H_1{:}\mu_i\theta_i(1) < 0$, for i=1,2,…, N_1, $\mu_i\theta_i(0) \neq 0$ for i = N_1 +1, N_1 +2,…,N.

ADF-Fisher, and PP-Fisher test are based on this form that consider the Ƥ autoregessive process, were expressed as follow:

$$\Delta Y_t = \vartheta_0 + \varnothing Y_{t-1} + \sum_{i=2}^{\rho} \tau_i \Delta Y_{t-i+1} + \varepsilon_t$$

Where $\varnothing = (1 - \sum_{i=1}^{\rho} \vartheta_1)$ and $\tau_i = \sum_{j=i}^{\rho} \vartheta_j$, for i = 1,2,…, Ƥ-1. Hence, the H_0: $\varnothing = 0$ and H_1: $\varnothing < 0$. While the PP-Fisher test under the null unit root hypothesis anticipated nonparametric transformation of the t-statistics for the Dickey and Fuller and the regression is:

$$\Delta Y_t = \Delta D_t + \tau Y_{t-1} + \Psi_t$$

The PP test adjust for any heteroskedasticity and serial correlation in residual ε_t of the tested regression, where Ψ_t is I(0) might be heteroskedastic under the null hypothesis that $\tau = 0$ assume that stationary across the cross-sectional units "individual effects."

After having confirmed the assumption all the variables are non-stationary, which is allow a large degree of the heterogeneity and dependence within and across the cross-sectional countries for the cointegration and the co-movement within the SPIs and OPI. For the co-movement between the variable we used three types of panel cointegration test, Pedroni (1999), Kao (1999) which is Engle and Granger (1987) two step residual based test, and Fisher which is a combined Johansen test. Therefore, estimated the panel cointegration assume that η_{it} is the residual for the differenced regression based on this form, (Pedroni, 1999).

$$\Delta Y_{it} = \phi_{1i} \Delta X_{1i,t} + \phi_{2i} \Delta X_{2i,t} + \eta_{i,t}$$

To ascertain clarification of lag length for the system, we calculate Ω as the long-run covariance (LRCOV) matrix of the estimated $\hat{\eta}_{i,t}$ using nonparametric kernel (Newey and West, 1987; Andrews, 1991), with sequence of mean-zero random p-vectors $\{V_t(\theta)\}$ depend on vector parameters θ were the LRCOV matrix is

$$\Omega = \sum_{j=-\infty}^{\infty} \Gamma(j)$$

Where $\Gamma(j) = E(V_t V_{t-j'})$, j≥0 and $\Gamma(j) = \Gamma(-j)$ j<0 is the autocovariance matrix of V_t at lag j. When V_t is second-order stationary, Ω equals 2π times the spectral density matrix of V_t evaluated at frequency zero (Hansen, 1982; Andrews, 1991; Hamilton, 1994). Therefore, we determined the lag length "white noise residuals" for the panel cointgreation regression, based on nonparametric Kernel with covariance matrix estimators in Anderws (1991) followed as:

$$\Omega\hat{} = \frac{T}{T-k} \sum_{j=-\infty}^{\infty} k\left(\frac{j}{b_T}\right).\Gamma\hat{}(j)$$

Where the autocovariance $\Gamma\hat{}$ (j) are,

$$\Gamma\hat{}(j) = \frac{1}{T} \sum_{t=j+1}^{T} \tilde{V}_t \tilde{V}_{t-j'} \quad , j \geq 0$$

$$\Gamma\hat{}(j) = \Gamma\hat{}(-\hat{j})', \, j < 0$$

K is the lag window or a symmetric kernel function in which the persistence among other conditions are |K(x)|≤1 for all x with K (0) = 1, and b_T > is a bandwidth parameter. Therefore, the leading T/(T-K) is a correction term for the degrees-of-freedom linked with the estimation of the K parameter in θ. Under the null hypothesis of no cointegration for the panel data we apply Kao (1999) cointegration test, expressed by the formula as follows:

$$Y_{it} = \alpha_i + \beta x_{it} + e_{it}, \text{ i=1,…,N,t=1,…,T}$$

Where

$$Y_{it} = Y_{it-1} + u_{it}$$

And

$$X_{it} = x_{it-1} + \varepsilon_{it}$$

α_i donated the fixed differing effect across cross-sectional units between the cross-country's, β is slope of parameter, Y_{it} and X_{it} are the independent random walk for all i, with the residual e_{it} I(1) series. The lag length selected subject on the weight of long-run covariance matrix of $w_{it} = (u_{it}, \varepsilon_{it})$ that implies the absence of the serial correlation among the cross-sectional units. Lastly, using the combined individual test (Fisher/Johansen), Johansen cointegration (1988) with two different approaches which are likelihood ratio trace statistics and maximum eigenvalue statistics respectively, in order to verify the existence of cointegration vectors in non-stationary time series based on this form.

$$\lambda_{trace}(r) = T\sum_{i=r+1}^{n}\ln(1-\hat{\lambda}_i)$$

And

$$\lambda_{max}(r,r+1) = T\ln(1-\hat{\lambda}_{r+1})$$

Where T denoted sample size, n−1 variables OPI, and $\hat{\lambda}_i$ is the i^{th} largest correlation among residual from the SPIs and OPI and the residual that derive from one dimensional differentiate processes. While, the trace test null hypothesis assumes at most r cointegration vector among the alternative hypothesis of r-n cointegration vector, and maximum eigenvalue statistics is to check the null hypothesis r cointegration vectors among the alternative hypothesis of r+1 cointegration vectors. However, after applying the long-run cointegration tests we then determine Individual deterministic and short-run coefficients, where liner model applied by using Bootstrapped resample residuals, based on this form.

$$\hat{\epsilon}_i = Y_i - \hat{Y}_i \; (i = 1,\ldots,n)$$

Create synthetic response variables and add a randomly resample residuals $\hat{\epsilon}_j$, to the response variable $Y_i \cdot Y_i^* = \hat{Y}_i + \hat{\epsilon}_j$, (j=1,…,n) for every i. Where $\hat{\epsilon}_j$, substituted in the model as follow:

$$Y_t = \alpha_i + X_t^{'}\beta_i + \sum_{i=-\rho}^{\rho=1}\hat{\epsilon}_{j+i}\, C_i + u_t$$

Where $X_t^{'}$ is a (j × 1) vector of oil piece index factor, β_i is the corresponding vector of factor, and C_i denoted the corresponding shock of OPI to SPIs, ρ is lag length "white noise residual." While $\hat{\epsilon}_j$ is the residuals of the OLS regression amid alteration of OPI on the individual SPIs for (6 GCC) countries. In order to determine the individual short-run effects from OPI to SPIs.

4. RESULTS AND DISCUSSIONS

Primarily, CD and LM tests were applied to approve the cross-section independence assumption for the SPIs between the GCC members, of which share a fair number of common economic or political characteristics. Table 2 shows the summary of the Pesaran (2004) CD test and LM test for the cross-section dependence between individual GCC SPI.

Table 2 reveals the null hypothesis of cross-section independence across GCC countries at a 1% level of significance is rejected. Therefore, the SPIs is found to be noticeably related to OPI across cross-section units is dependence, considering the oil price to have equivalent values for all cross-section units. Due to the rejection of the cross-section independence assumption across 6GCC countries, there are three types of panel unit roots that are used to examine the stationary of the data. Second generation unit root test (CADF), common unit root process test, the individual unit root test, Im, Pesaran and Shin, ADF-Fisher, and PP-Fisher.

Probabilities tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. Lag length selection based on individual LM test white noise residuals.

Results in Table 3 indicate the variables are not integrated on level because the null hypothesis of unit root existence cannot be rejected on level for all the variables for p = 1,…,3 "white noise residuals" at 1% level of significance. While null hypothesis of unit root existence is rejected on first difference I (1) for all variables for p = 1,…,3 "white noise residuals" at 1% level of significance. According to prior findings, SPIs (LSPI) and LOPI for the (6GCC) countries are integrated. However, in order to verify whether the variables are freely moving across cross-sectional units, the individual unit root test (Im, Pesaran and Shin, ADF-Fisher, and PP-Fisher) was applied. Whereas, Levin Lin and Chu (LLC) tests are applied to verify if the persistence parameters are common across cross-section unit, due to the OPI which has the same values for all the (6GCC) countries, as summarized in Table 4.

Findings in Table 4 declare SPIs and OPI for (6 GCC) countries are not stationary on the level I (0). This means, rejection of null hypothesis of the existence of unit root in both cases of common panel unit root and individual unit root. This indicates variables are not integrated in the level I (0). After taking the first difference of the series in both cases, the null hypothesis of unit root is rejected for all the variables at 1% level of significance. This shows that the SPIs and OPI are integrated by first difference I (1), for the (6 GCC) countries. After having confirmed the non-stationary of the LSPI and LOPI series, we verify the long-run cointegration "co-movement" between the variables using three types of panel cointegration tests as summarized in Tables 5-7.

Pedroni residual cointegration test results in Table 5 imply rejection of null hypothesis of no cointergration at a significance level of 1%. In other words, a long-run relationship amid LOPI and LSPI for the (6 GCC) countries in all cases accepts the null hypothesis at significance level of 5% except the panel v-statistic.

The result in Table 6 shows that Kao residual cointegration test rejects null hypothesis of no cointegration at P = 0.000, noting that this is a high significant level. This shows the robust evidence that the series has long-run relationship for the (6 GCC) countries.

Table 2: Pesaran (2004) CD and LM test, cross section dependence for panel 6 GCC country

Variable	CD test statistic	P value*	LM test statistic	P value*
LSPI	5.416	0.000	4.178	0.000

*Null hypothesis: Cross-sectional independence

Johansen Fisher Panel cointegration test in both cases of trace test and max-Eigen test as shown in Table 7, that the series has long-run association among SPIs and OPI. Therefore, after having confirmation of the companionship of the long-run relationship in the midst of the series, the linear short-run effect or the shocks from OPI on SPIs for the six GCC members are checked individually. Whereas, the linear model was applied in light of the confirmation of the stationary of SPIs and OPI at the same level I (1). This means they are within-group variance, (Pinheiro and Bates, 2006).

Table 8's results indicate significant linear short-run association amid SPIs and OPI in Oman, Qatar, and UAE at 1% level of significance. Furthermore, the stock markets of these countries are responding to the effect of OPI at a 5% level of significance. From the coefficients for these countries we can observe a positive linear short-run association amid stock price and oil price. Whereas, Oman's correspondence with oil price shocks by 0.76% would change the market value by approximately 0.117. UAE is correspondence with oil price shocks by 0.7% would change the market value by 0.785 which has a greater sensitivity to sudden shocks of oil price within short term period. Meanwhile, Qatar is correspondence with oil price shocks by 0.3% would change the market value by 0.612, which indicates a lower correspondence of oil price shocks with high volatility of its market values. This represents a change in approximately double its market value, compared with Oman and UAE market values. Thus, it can be implied that GCC stock markets are segmented from the world market, since they represent dominant world energy market players, (Arouri et al., 2010). In other words, it is considered to be a high international portfolio diversifications opportunity for investors from developed countries and likewise from emerging ones. Due to oil dependent economic factors, absence of linear short-run association amid SPIs and OPI in Saudi Arabia, Kuwait, and Bahrain seems to be counterintuitive. Moreover, Figure 1 shows that the stock markets of Oman, Qatar, and UAE are underperforming because of sudden shock responses of oil price, while stock markets of Saudi Arabia, Kuwait, and Bahrain are top performers due to the sudden shocks avoidance of oil price within short time period.

5. CONCLUSION AND POLICY IMPLICATIONS

This examination expands and illuminates the relationship amid stock market values in the six GCC members and oil prices. In view of the fact that these country members are the major global players in energy possessions; their stock markets are apparently vulnerable and have influence the oil crisis. Short and long run have been applied for two unit's dependencies. In terms of short-run analysis, strong positive correlations involving stock markets in UAE, Qatar, and to some extent Oman with alteration in oil prices. Interestingly, weaknesses for Qatar market of sudden shock responses of oil price, which is affected by 0.3% which represents change in the market value by 0.612. While, Oman and UAE have responses of oil price shocks approximately by 0.67% and 0.7% which affect in their market values by 0.117 and 0.785 respectively, which indicates these markets are underperforming. In the case of Bahrain, the Kuwait, and Saudi Arabia their markets are not responded of oil price shocks within a short time period which indicates a top performing market. Furthermore, long-run analysis shows indication of a long-run relationship in the midst of GCC country member's stock markets and oil price's shocks.

Our results underwent through great concernment of market participants and policymaker. Principally, GCC countries should, as OPEC politicians, take in their consideration the impact of fluctuations of oil prices on their particular security markets and

Table 3: Cross-sectionally CADF unit root results

Variable	Specification	*P=1	*P=2	*P=3
LSPI	Intercept	0.974	0.970	0.971
Δ LSPI	Intercept	0.000	0.0031	0.0014

Null: Unit root. *Lag length selection based on individual LM test white noise residuals

Table 4: Common and individual panel unit root tests

Variable	K*	LLC *P* value**	Im, Pesaran and Shin *P* value**	ADF-Fisher Chi-square *P* value**	PP-Fisher Chi-square *P* value**
LSPI	3	0.4024	0.1436	0.1444	0.2762
Δ LSPI	3	0.000	0.000	0.000	0.000
LOPI	3	-	0.0373	0.1217	0.1311
Δ LOPI	3	-	0.000	0.000	0.000

Null: Unit root, Exgenous variable: Individual effect. LLC test: Assumes common unit root process. Im, Pesaran and Shin: Assumes individual unit root process. ADF-Fisher Chi-square: Assumes individual unit root process. PP-Fisher Chi-square: Assumes individual unit root process. *Lag length selection based on individual LM test white noise residuals. **Probabilities for Fisher test are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. LLC: Levin Lin and Chu

Table 5: Pedroni residual cointegration test

Series	Panel v-stat		Panel rho-stat		Panel pp-stat		Panel ADF-stat	
	Statistic	P	Statistic	P	Statistic	P	Statistic	P
LSPI, LOPI	−1.24	0.89	−62.34	0.000	−25.32	0.000	−7.95	0.000
Group rho-stat			Group pp-stat		Group ADF-stat			
Statistic		P	Statistic	P	Statistic	P		
−52.26		0.000	−26.70	0.000	−7.81	0.000		

Null hypothesis: No cointegration. Trend assumption: No deterministic intercept or trend. Lag length=5 selected based on SIC and ACI long-run covariance prewhithening=3 white noise residuals

Table 6: Kao residual cointegration test

Series	ADF	
	t-stat	P
LSPI, LOPI	−9.43	0.000

Null hypothesis: No cointegration. Trend Assumption: no deterministic intercept or trend. Lag length=5 selected based on SIC and ACI long-run covariance prewhithenin g=3 white noise residuals

Table 7: Johansen Fisher panel cointegration test

Series	No of CE (s)	Fisher stat*(from trace test)	P	Fisher Stat*(from max-Eigen test)	P
LSPI, LOPI	None	171.1	0.000	98.14	0.000
	At most 1	131.2	0.000	131.2	0.000

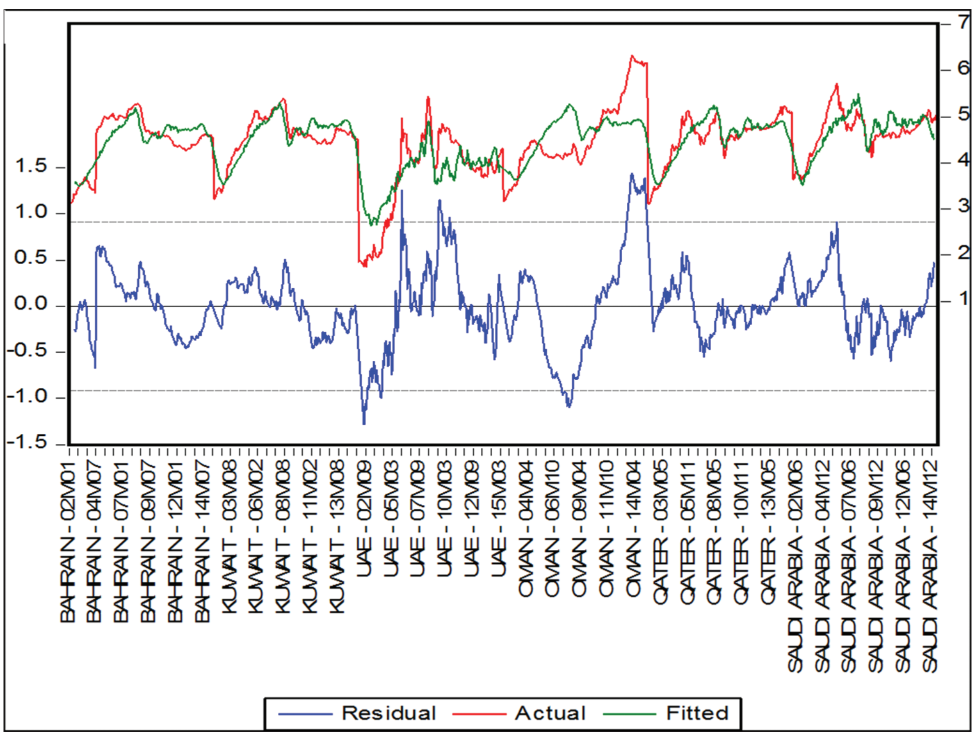
Trend assumption: No deterministic intercept or trend. Lag length=5 selected based on SIC and ACI long-run covariance prewhithening=3 white noise residuals. *Probabilities are computed using asymptotic Chi-square distribution

Table 8: Short-run results-linear models

Parameters	Bahrain	Kuwait	Oman	Qatar	UAE	Saudi. A
αi	0.008	0.006	0.017	0.01	0.012	0.007
βi	0.247	0.335	0.117	0.612	0.785	0.366
	(0.146)	(0.071)	(0.075)**	(0.100)*	(0.201)**	(0.112)
Ci	0.0037	0.0018	0.0067	0.003	0.007	0.002
	(0.0027)	(0.0013)	(0.0032)**	(0.002)**	(0.003)**	(0.001)
R^2	0.026	0.12	0.018	0.19	0.092	0.077
Log likelihood	128.106	242.514	234.629	187.251	78.515	170.746
AIC	−1.563	−2.993	−2.895	−2.303	−0.943	−2.096

Bootstrapped coefficient estimates and standards errors (10000 repetitions). Resample residuals type, *, ** at 1% and 5% level of significance

Figure 1: The residual volatility across cross-sectional units based on the effect from oil price index on stock price index



overall economies. From investors' perspective, the essential link amid stock and oil prices, in term of the stock markets respondents of oil prices shocks involve a level of predictability and anticipation in the GCC markets. The achieved results propose several options for further inquiries and examinations. Oil and values of the GCC countries to be expected between markets to diversify the economic ties of various fields.

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