How do Macroeconomic Changes Impact Islamic and Conventional Equity Prices? Evidence from Developed and Emerging Countries

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How do Macroeconomic Changes Impact Islamic and Conventional Equity Prices? Evidence from Developed and Emerging Countries

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ABSTRACT

In theory, the price of equity is determined by the dividend yields and growth potentials of the firms. There exists established empirical proof of the impact of macroeconomic changes to the equity markets. With the advent of Islamic equities, and the recent surge of interest in them have raised the question of whether the same theoretical framework and relationship be considered for Shariah compliant equities or not. This study explores the impact of macroeconomic changes on Islamic and conventional indices for a large set of 37 countries, classifying them according to developed and emerging countries. The study finds a higher impact of Industrial production on the Islamic equities, while the interest rate and money supply have a lesser impact as compared to the impact on conventional counterparts. This lends support to the argument that Shariah screening methodology provides a set of Islamic equities which are more founded on the real sector of the economy. In addition the adjustment process during the crisis is faster for the Islamic equities in both regions. These results provide initial empirical proof for further research on the impact of specific economic variables on the changes in Islamic equity prices.

Key Words: Islamic finance, Stock Market, Equity, Emerging Countries

JEL Classification Codes: C22, C58, E44, G15

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1. Introduction

Literature in finance has explored the pricing of equities for decades. According to established works of theory the price is determined by the dividend yields and growth potentials of the firms. While the latter is dependent on internal factors of the firm, it is also heavily dependent on the economic health and prospects of the business environment. A well-functioning stock market is widely considered as an indicator of the future potential of the economy. Based on this premise the price of a firm in the stock market should reflect the future potential which is indicated by the discounted dividend payouts. A standard pricing equation in finance literature is the Dividend Discount Model proposed in Gordon (1959) theoretically underpinned in "The Theory of Investment Value" by John Burr Williams in 1938.

Revisiting the valuation of securities, the dividend payout forms the numerator of the equation while the denominator is the discount rate. The discount rate is affected by the interest rate and inflation in the economy alongwith the policy measures the regulators take into account. This brings us to our key study area; since the pricing of equity is impacted by macroeconomic variables, do they impact in similar manner to conventional and Islamic equities.

Several studies have focused on inquiring the validity of macroeconomic fundamentals on equity prices. Fama (1990), Barro (1990), and Schwert (1990) among others, found strong short-run correlations between the two variables for the United States (US) and amongst these variables and other financial variables. Significant number of studies have also been undertaken for several major international stock markets (e.g., Cheung and Ng, 1998; Chung et al., 1998), documenting the same positive linkage between stock returns and real (aggregate) economic activity.

A plethora of literature exists raising questions on the validity of macroeconomic fundamentals as the predictors of equity prices and returns. Carlson and Sargent (1997) and Shiller (2005) argue that in US most of the rise in equity prices during the second part of the 1990s cannot be attributed to fundamental values such as projected earnings growth or dividends but to exogenous shocks and/or irrational market behavior. Along the same lines Lee (1995, 1998) and Chung and Lee (1998) report that fundamental variables like discount rates, earnings, dividends and industrial production did not explain price movements. More recent evidences of violation of the inter-linkage between stock returns and real economic activities are chronicled in the studies by Binswanger (2000, 2001, 2004) and Laopodis (2006), who found no consistent economic behavior and argue that stock prices may be affected owing to stock markets bubbles and irrational behavior in the markets. Moreover, a series of recent financial crises including 2008 global financial crisis again challenged the nexus between macroeconomic fundamentals and equity prices. Equity markets in majority of the developed and some developing countriesexperienced a sharp downturn believed to have been triggered by excessive speculation and amplified by the higher level of economic integration during the recent financial crisis.

Quite a few studies have been undertaken to understand the conflicting findings in recent literature on equity prices/returns and macroeconomic fundamentals; however there are very few studies observing this relationship in the context of Islamic equity markets. The current trend of Islamic finance's move towards the global markets, especially expansion from a merely banking-based industry into a wider spectrum of financial market-based instruments, has made Islamic capital markets, the fastest growing sector in the Islamic finance industry. Such an important question as to whether *Shariah*-compliant (i.e. Islamic law compliant) equities have similar risk-return profile relative to those of conventional counterparts remains unclear. This is solely due to the role of *Shariah* (Islamic) rules, which have distinguished between *halal* (lawful) and *haram* (unlawful), resulting in the unique characteristics of Islamic equities Derigs & Marzban, (2008). The qualitative *Shariah* screening excludes firms with any non-compliant activity (i.e. liquor, gambling, interest-based financial institutions, etc.) while the quantitative *Shariah* screening strictly imposes the zero interest-based leverage. Since only a small number of today's listed firms fit into this requirement, thus some certain degree of tolerance is required[6]. As a result, the filtering criteria will take out the large non-compliant firms from the pool of investable equities, leaving the remaining Shari'ah compliant firms available to become smaller and portray more volatile returns (Hussein &Omran, 2005). In other words, the lower leverage, smaller size of firms, and under-diversification of the market, will be the main distinctive features that potentially lead Islamic stocks to behave differently compared to the conventional counterparts.

To the best of our knowledge, the existing literature lacks is any rigorous empirical study focused on comparing the Islamic vis-à-vis conventional equity indices with regard to the impact of key macroeconomic variables on the stock price. Therefore, this study sets out to perform a comparative analysis by examining the relationship between major macroeconomic fundamentals (industrial production, money supply and consumer price index) and equity prices (both Islamic and conventional). The main contribution of the study is to observe this inter-linkage by using a large number of countries in order to obtain a wide coverage. We use a panel data of 37 countries, both developed and emerging countries, with monthly observations from January 2008 till December 2011. Since it is only MSCI which have Islamic equity indices that comprise a large number of countries, hence the relatively short numbers of years are used taking into account the limited availability of the length of observations for these indices. Moreover, the study investigates the influence of changes in macroeconomic fundamentals on (i) Shariah compliant equities and (ii) mainstream equities in (i) developed, (ii) emerging, and (iii) whole countries. As to the methodology, we apply Panel unit root, cointegration, while the estimation will use the dynamic heterogeneous panel techniques that allows coefficients to vary across different individual groups.

The paper is organized as follows. Section 2 presents some literature reviews associated with the macroeconomic impacts as well as the issue of Islamic equity markets. Section 3 presents the data and methodology. Section 4, presents and discusses the empirical results, while Section 5 provides conclusions and policy implications.

2. Literature review

2.1. The impact of macroeconomic variables

The seminal work of Fama (1990) investigated the extent to which changes in future cash flows and discount rates explain variations in stock returns in the US market. Other significant studies on the same issue for several international markets was undertaken by Barro (1990), Schwert (1990), Cheung and Ng (1998) and Chung et al (1998). Recently, Jangkoo et al (2011) developed a conditional version of the consumption capital asset pricing model (CCAPM) using the conditioning variable from the cointegrated macroeconomic variables such as dividend yield, term spread, default spread, and short-term interest rate. The study used quarterly data starting from 1963:Q3 to 2005:Q4. Findings of the study suggest

[6](i) a company's debt financing is not more than 33 percent of its capital, (ii) interest-related income of a company is not more than 10 percent of its total income, (iii) the composition of account receivables and liquid assets (cash at banks and marketable securities) compared to total assets is minimum at 51 percent while a few cite 33 percent as an acceptable ratio.

that conditioning variable has a strong power to predict market excess returns in the presence of competing predictive variables.

If the stock is efficiently priced, both cash flows of a firm and the discount rate will theoretically be dependent on the macroeconomic fundamentals over time. Therefore, empirically testing linkages between equity prices and macroeconomic fundamentals is one of the major areas of interest in finance literature. Many studies have investigated the impact of key macroeconomic variables on stock prices. Laopodis (2006) investigated the dynamic linkages among stock prices, interest rates, inflation, and economic activity for the United States since the 1970s. Analysis of this study indicates absence of dynamic nexus between real economic activity and stock prices across different monetary regimes during the last thirty years. In another study, Pesaran and Timmermann (1995) investigated the robustness of the evidence on predictability of US stock returns and addressed the issue of whether this predictability could have been historically exploited by investors to earn profits in excess of a buy-and-hold strategy in the market index. The research used monthly data (of stock price, annualized dividends and earnings, 1-month T-bill rate, Inflation rate, change in industrial production, excess return on stocks) from 1954:M1 to 1992:M12 and found that the predictive power of various economic factors over stock returns changes through time and tends to vary with the volatility of returns.

Recalling the importance of economic integration, the impact of cross-country macroeconomic variables on stock price in respective country has been observed in literature. Verma and Ozuna (2005) examined the responsiveness of Latin American Stock markets to movements to changes in cross-country macroeconomic variables. The study used monthly data of stock price, money supply, consumer price index, interest rates and exchange rates of four Latin American countries such as Mexico, Argentina, Brazil and Chile. They found the presence of Mexican stock market's influence on Latin American stock markets and found little evidence that Latin American stock markets are responsive to these changes. Yang et al (2009) studied the time-varying stock–bond correlation over macroeconomic conditions (the business cycle, the inflation environment and monetary policy stance) by applying a class of bivariate AR(1)–GARCH (1,1) models for conditional correlations between stock and bond premiums. Results of the study came up with different patterns of time variation in stock–bond correlations over the business cycle between US and UK. The study, furthermore, argued that higher stock–bond correlations tend to follow higher short rates and (to a lesser extent) higher inflation rates.

Some studies have been undertaken specifically for key developed countries. By using monthly data from 1990:01 to 2009:12, Nikiforos (2011) studied the dynamic linkages between equity price and major macroeconomic indicators of France, Germany, Italy, UK, and the US. Applying rolling cointegration and VAR approaches, the study revealed different ways of responsiveness of stock prices to changes in macroeconomic fundamentals. Results of the study suggested that stock markets move more independently in the long run particularly in the post-euro period. For instance, equity prices were not much stimulated by industrial production or interest rates. Furthermore, the study found that European consumers are more concerned about the general economic conditions and personal financial situations rather than inflationary pressures in both pre- and post-Euro sub-periods.

While analyzing the impact of macroeconomic news on stock returns, Briz and John (2011) have investigated how stock returns are responsive to newspaper stories about the releases of new macroeconomic information in the economy. Results of the study indicated that newspaper interpretation of the GDP news does affect stock returns. As to the impact of consumer price index, Suk-Joong et al (2004) investigated the impact of scheduled government announcements of six different macroeconomic variables (nominal foreign international trade balance, gross domestic product, unemployment rate, retail sales growth, consumer price index and producer price index) on the risk and return of three major US financial markets (stock, bond and foreign exchange markets). Using GARCH modeling, results of the study suggested that these markets do not respond in any meaningful way to the act of releasing information by the government rather the 'news' content of these announcements cause the market to react. News related to the internal economy was found to be important for the bond market and stock market is more influenced by the consumer and producer price information release. Nikkihen and Sahlstrom (2004) studied the impact of the scheduled Federal Open Market Committee (FOMC) meetings and the scheduled macroeconomic news releases on stock market uncertainty by employing regression and GARCH approaches. The study used monthly reports of employment, CPI and PPI, and FOMC meeting days on the US market covering the period from January 1996 to December 2000. Results suggested that the employment report has the largest impact on stock price uncertainty, whereas, investors regard the information content of the PPI and CPI together as significant.

2.2. Empirical studies in Islamic equity indices

There have been a number of empirical studies that compare Islamic assets to their conventional counterparts. Al-Zoubi & Maghyereh (2007), applying Risk Matrices, Student-t APARCH and skewed Student-t APARCH, show that the DJIM (Dow Jones Islamic Market index) is less risky than its respective benchmark. Another study focuses on bubble formation by applying duration dependence tests of survival analysis, and find none of the evidence of speculative bubbles for weekly and monthly returns of AMANX, AMAGX and DJIMI (Hassan & Tag El-Din, 2005).

Apart from asset pricing model, Hakim & Rashidian (2002) use CAPM and find that the DJIMI performs well as compared to the Dow Jones World Index (DJW), but underperforms the Dow Jones Sustainability World Index (DJS). By capturing the effects of industry, size, economic conditions, and performance measures, some studies also show that Islamic indices outperform during bull period while underperform during bear period, with the reasons of investing in growth and small-cap firms (Hussein 2004, 2005; Girard & Hassan, 2005).

Other studies have focused on mutual funds' performance, and find that Islamic funds perform averagely similar to other conventional counterparts, and even are subject to multiple regimes (Hassan, Antoniou & Paudyal, 2005; Elfakhani, Hassan & Sidani, 2005; Hassan & Antoniou, 2006; Abdullah, Hassan, & Mohamad, 2007). Hoepner, Rammal, & Rezec (2011) find that Islamic funds from Malaysia or GCC neither significantly underperform their respective benchmarks nor are significantly affected by small-size stocks.

To the best of our knowledge, what is lacking in the existing literature is any rigorous empirical study focused on comparing the Islamic vis-à-vis conventional equity indices with regard to the impact of key macroeconomic variables on the stock price. Therefore, our paper is the first study using a panel data of 37 countries, both developed and emerging countries.

3. Data and methodology

We use monthly data of Islamic and conventional stock indices of 37 countries. The data is collected from MSCI equity indices. The observations consist of monthly data extending over four years starting with January 2008 due to the availability of MSCI Islamic indices that cover all countries of our interest.

On the other hand, the key macroeconomic variables include real production index, real money supply, consumer price index and short-term interest rate. We take M1 and 3-

month interbank middle rate as a proxy for money supply and short-term interest rate. This allows us to obtain a standard measure across all countries since we cannot get M2 as well as 3-month T-bill rate for some countries. All dataset are collected from *Datastream*. We classify the developed and emerging countries according to IMF criteria (see Table 1).

	Developed Countries	**Emerging Countries**
1	Australia	Argentina
2	Belgium	Brazil
3	Canada	Chile
4	Denmark	China
5	Finland	Colombia
6	France	Czech
7	Germany	Egypt
8	Greece	Hungary
9	Ireland	India
10	Italy	Indonesia
11	Japan	Malaysia
12	Netherlands	Peru
13	Norway	Philippines
14	Singapore	Poland
15	Spain	Russia
16	Sweden	South Africa
17	Taiwan	South Korea
18	UK	Turkey
19	USA	

Table 1: List of Countries in the Sample

3.1. Panel Unit Root tests

Panel unit root tests are performed in order to investigate stationarity of data in panel format. LL (1993) proposed a panel based ADF test that restricts parameters γ_i by keeping them identical across cross sectional regions as follows:

$$\Delta y_{it} = \alpha_i + \gamma_i y_{it-1} + \sum_{j=1}^{k} \alpha_j \Delta y_{it-j} + e_{it} \qquad (1)$$

, where $t = 1 \ldots\ldots, T$ time periods and $i = 1 \ldots\ldots, N$ members of the panel. LL tests the null hypothesis of $\gamma_i = \gamma = 0$ for all i, against the alternative of $\gamma_1 = \gamma_2 \ldots\ldots.. = \gamma < 0$ for all i. The shortcoming of this test is that γ is restricted by being kept identical across regions under the null and alternative hypothesis.

In order to overcome the shortcomings of the LL test, IPS (1997) relaxed the assumption of the identical first-order autoregressive coefficients of the LL test and allow γ to vary across regions under the alternative hypothesis. IPS test the null hypothesis of $\gamma_i = 0$ for all i, against the alternative of $\gamma_i < 0$ for all i. The IPS test is based on the mean-group approach, which uses the average t_{γ_i} statistics to perform the following $\bar{Z}$ statistic:

$$\bar{Z} = \sqrt{N}(\bar{t} - E\bar{t}))/\sqrt{Var(\bar{t})} \qquad (2)$$

, where $\bar{t} = (1/N)\sum_{i=1}^{N} t_{\gamma_i}$, the terms $E(\bar{t})$and $Var(\bar{t})$ are, respectively, the mean and variance of each t_{γ_i} statistic, and they are generated by simulations and are tabulated in IPS (1997). The $\bar{Z}$ converges to a standard normal distribution. Based on the Monte Carlo experiment results, IPS demonstrates that their test has more favorable finite sample properties than the LL test.

Hadri (2000) argues differently that the null should be reversed to be the stationary hypothesis in order to have a stronger power test. His Lagrange multiplier (LM) statistic can be written as below:

$$LM = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{\frac{1}{T^2}\sum_{t=1}^{T}S_{it}^2}{\hat{\sigma}_\varepsilon^2}\right), S_{it} = \sum_{j=1}^{t}\widehat{\varepsilon_{ij}} \quad (3)$$

, where $\hat{\sigma}_\varepsilon^2$ is the consistent Newey and West (1987) estimate of the long-run variance of disturbance terms.

3.2. Panel Cointegration tests

Perdroni (1999) considers the following time series panel regression:

$$y_{it} = \alpha_{it} + \delta_{it}t + X_i\beta_i + e_{it} \quad (4)$$

where y_{it} and X_{it} are the observable variables with dimension of $(N^*T) \times 1$ and $(N^*T) \times m$, respectively. He develops asymptotic and finite sample properties of testing statistics to examine the null hypothesis of non-cointegration in the panel. The test allow for heterogeneity among individual members of the panel, including heterogeneity in both long-run cointegrating vectors and in the dynamics, since there is no reason to believe that all parameters are the same across countries.

Pedroni suggested two types of tests. The first type is based on the within dimension approach, which includes four statistics. They are panel υ-statistic, panel ρ statistic, panel PP-statistic, and panel ADF-statistic. These statistics pool the autoregressive coefficients across different members for the unit root tests on the estimated residuals. The second test by Pedroni is based on the between-dimension approach, which includes three statistics. They are group ρ-statistic, group PP-statistic, and group ADF-statistic. These statistics are based on estimators that simply average the individually estimated coefficients for each member. Following Pedroni (1999), the heterogeneous panel and heterogeneous group mean panel cointegration statistics are calculated as follows:

Panel υ-statistic:

$$Z_v = \left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{it-1}^2\right)^{-1} \quad (5)$$

Panel ρ statistic:

$$Z_p = \left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{it-1}^2\right)^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{\cdot 2}(\hat{e}_{it-1}\Delta\hat{e}_{it} - \hat{\lambda}_i) \quad (6)$$

Panel PP-statistic:

$$Z_t = \left(\hat{\sigma}^2 \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^2\right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i) \qquad (7)$$

Panel ADF-statistic:

$$Z_t^* = \left(\hat{s}^{*2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{*2}\right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*} \qquad (8)$$

Group ρ-statistic:

$$\hat{Z}_\rho = \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{e}_{it-1}^2\right)^{-1} \sum_{t=1}^{N} (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i) \qquad (9)$$

Group PP-statistic:

$$\tilde{Z}_t = \sum_{i=1}^{N} \left(\hat{\sigma}^2 \sum_{t=1}^{T} \hat{e}_{it-1}^2\right)^{-1/2} \sum_{t=1}^{T} (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i) \qquad (10)$$

Group ADF-statistic:

$$\tilde{Z}_t = \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{s}_i^2 \hat{e}_{it-1}^{*2}\right)^{-1/2} \sum_{t=1}^{T} (\hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*}) \quad (11)$$

Here, $\hat{e}_{it}$ is the estimated residual from equation (4) and $\hat{L}_{11i}^2$ is the estimated long run covariance matrix for $\Delta \hat{e}_{it}$. Similarly, $\hat{\sigma}_i^2$ and $\hat{s}_i^2$ $\left(\hat{s}_i^{*2}\right)$ are, respectively, the long-run and contemporaneous variances for individual i. The other terms are properly defined in Pedroni (1999) with the appropriate lag length determined by the Newey-West method. All seven tests are distributed as being standard normal asymptotically. This requires a standardization based on the moments of the underlying Brownian motion function. The panel v-statistic is a one-sided test where large positive values reject the null of no Cointegration. The remaining statistics diverge to negative infinitely, which means that large negative values reject the null. The critical values are also tabulated by Pedroni (1999).

3.3. The MG and PMG estimators

Assume an autoregressive distributive lag (ARDL) (p; q_1…. q_k) dynamic panel specification of the form

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=o}^{q} \delta'_{ij} X_{i,t-j} + \mu_i + \epsilon_{it} \qquad (12)$$

where the number of groups i = 1; 2; : : : ;N; the number of periods t = 1; 2; : : : ; T; X_{it} is a k 1 vector of explanatory variables; δ_{it} are the k*1 coefficient vectors; λ_{ij} are scalars; and $\mu_i$is the group-specific effect. T must be large enough such that the model can be fitted for each group separately. Time trends and other fixed regressors may be included.
If the variables in (1) are, for example, I(1) and cointegrated, then the error term is an I(0) process for all i. A principal feature of cointegrated variables is their responsiveness to any deviation from long-run equilibrium. This feature implies an error correction model in which the short-run dynamics of the variables in the system are influenced by the deviation from equilibrium. Thus it is common to reparameterize (1) into the error correction equation.

$$\Delta y_{it} = \emptyset_i\left(y_{i,t-1} - \theta'_i X_{it}\right) + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-1} + \sum_{j=o}^{q-1} \delta_{ij}^{\prime *} \Delta X_{i,t-j} + \mu_i + \epsilon_{it} \qquad (13)$$

where $\emptyset_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right)$, $\theta_i = \sum_{j=0}^{q} {}^{\delta_{ij}}\big/_{(1 - \sum_k \lambda_{ik})}$, $\lambda_{ij}^* = -\sum_{m=j+1}^{p} \lambda_{im}$ $j = 1,2,\dots,p-1$,

$and\ \delta_{ij}^* = \sum_{m=j+1}^{q} \delta_{im}$ $j = 1,2,\dots,q-1$.

The parameter $\emptyset_i$ is the error-correcting speed of adjustment term. If $\emptyset_i = 0$, then there would be no evidence for a long-run relationship. This parameter is expected to be significantly negative under the prior assumption that the variables show a return to a long-run equilibrium. Of particular importance is the vector $\theta_i^{'}$ which contains the long-run relationships between the variables.

The recent literature on dynamic heterogeneous panel estimation in which both N and T are large suggests several approaches to the estimation of (2). On one extreme, a _fixed-effects (FE) estimation approach could be used in which the time-series data for each group are pooled and only the intercepts are allowed to differ across groups. If the slope coefficients are in fact not identical, however, then the FE approach produces inconsistent and potentially misleading results. On the other extreme, the model could be fitted separately for each group, and a simple arithmetic average of the coefficients could be calculated. This is the MG estimator proposed by Pesaran and Smith (1995). With this estimator, the intercepts, slope coefficients, and error variances are all allowed to differ across groups.

More recently, Pesaran, Shin, and Smith (1997, 1999) have proposed a PMG estimator that combines both pooling and averaging. This intermediate estimator allows the intercept, short-run coefficients, and error variances to differ across the groups (as would the MG estimator) but constrains the long-run coefficients to be equal across groups (as would the FE estimator). Since (2) is nonlinear in the parameters, Pesaran, Shin, and Smith (1999) develop a maximum likelihood method to estimate the parameters.

Expressing the likelihood as the product of each cross-section's likelihood and taking the log yields

$$l_T = (\theta', \varphi', \sigma') = -\frac{T}{2}\sum_{i=1}^{N} \ln\left(2\pi\sigma_i^2\right) - \frac{1}{2}\sum_{i=1}^{N} \frac{1}{\sigma_i^2}\{\Delta y_i - \emptyset_i \xi_i(\theta)\}' H_i \{\Delta y_i - \emptyset_i \xi_i(\theta)\} \quad (14)$$

For $\mathrm{i} = 1, \dots, \mathrm{N}$, where, $\xi_i(\theta) = y_{i,t-1} - X_i\theta_i$,

$H_i = I_T - W_i(W_i'W_i)W_i, I_T is\ an\ identity\ matrix\ of\ order\ T, and\ W_i = \left(\Delta y_{i,t-1}, \dots, \Delta y_{i,t-1}, \Delta y_{i,t-p}, \Delta X_i, \Delta X_{i,t-1}, \dots, \Delta X_{i,t-q+1}\right).$

Beginning with an initial estimate of the long-run coefficient vector, $\hat{\theta}$, the short-run coefficients and the group-specific speed of adjustment terms can be estimated by regressions of $\Delta y_i on\ (\widehat{\xi}_{i,}, W_i)$. These conditional estimates are in turn used to update the estimate of θ. The process is iterated until convergence is achieved.

The parameter estimates from iterated conditional likelihood maximization are asymptotically identical to those from full-information maximum likelihood. But the estimated covariance matrix is not. However, since the distribution of the PMG parameters is known, we can recover the full covariance matrix for all estimated parameters. As shown in Pesaran, Shin, and Smith (1999), the covariance matrix can be estimated by the inverse of

$$\begin{bmatrix} \sum_{i=1}^{N} \frac{\hat{\phi}_i^2 X_i' X_i}{\hat{\sigma}_i^2} & \frac{-\hat{\phi}_1 X_1' \hat{\xi}_1}{\hat{\sigma}_1^2} & \cdots & \frac{-\hat{\phi}_N X_N' \hat{\xi}_N}{\hat{\sigma}_N^2} & \frac{-\hat{\phi}_1 X_1' W_1}{\hat{\sigma}_1^2} & \cdots & \frac{-\hat{\phi}_N X_N' W_N}{\hat{\sigma}_N^2} \\ & \frac{\hat{\xi}_1' \xi_1}{\hat{\sigma}_1^2} & \cdots & 0 & \frac{\hat{\xi}_1' W_1}{\hat{\sigma}_1^2} & \cdots & 0 \\ & & \ddots & \vdots & \vdots & \ddots & \vdots \\ & & & \frac{\hat{\xi}_N' \hat{\xi}_N}{\hat{\sigma}_N^2} & 0 & \cdots & \frac{\hat{\xi}_N' W_N}{\hat{\sigma}_N^2} \\ & & & & \frac{W_1' W_1}{\hat{\sigma}_1^2} & \cdots & 0 \\ & & & & & \ddots & \vdots \\ & & & & & & \frac{W_N' W_N}{\hat{\sigma}_N^2} \end{bmatrix}$$

The MG parameters are simply the un-weighted means of the individual coefficients. For example, the MG estimate of the error correction coefficient, ϕ, is

$$\hat{\emptyset} = N^{-1} \sum_{i=1}^{N} \hat{\emptyset} \qquad (15)$$

with the variance

$$\hat{\Delta}_{\hat{\emptyset}} = \frac{1}{N(N-1)} \sum_{i=1}^{N} (\hat{\emptyset}_i - \hat{\emptyset})^2 \qquad (16)$$

The mean and variance of other short-run coefficients are similarly estimated.

3.4. The Hausman test

We perform Hausman test to make a choice between the PMG and MG approaches. Hausman (1978) assumes that there are two estimators $\hat{\beta}_0$ and $\hat{\beta}_1$ of the parameter vector β and he added two hypothesis testing procedures. We test null hypothesis (H_0), where PMG are consistent and efficient against alternative hypothesis (H_1), where PMG are inconsistent (as the MG are always consistent). Hausman test uses the following test statistic:

$$H = (\hat{\beta}^{PMG} - \hat{\beta}^{MG})' \left[Var(\hat{\beta}^{PMG}) - Var(\hat{\beta}^{MG})\right]^{-1} (\hat{\beta}^{PMG} - \hat{\beta}^{MG}) \sim \chi^2(k) \qquad (17)$$

The difference between the estimates is significant if the value of the statistic is large. Accordingly, we reject the null hypothesis that the PMG model is consistent and we use the MG estimators. In contrast, a small value of the Hausman statistic indicates that the PMG estimator is more appropriate.

4. Empirical results

Our empirical study performs a comparative analysis in different models of estimation according to different group of countries. As both pooled mean group and mean group estimators involve the long-run equilibrium relationship, the variables cannot be consistently estimated when all the single variables have unit roots or are non-stationary of order one, unless the variables in the long-run relationship are co-integrated. Hence we need to perform panel unit roots test for all the variables and test whether co-integrating equilibrium relationship between variables exist.

We split the samples into six different models. Panel I and II will observe the impact of macroeconomic determinants on the Islamic and conventional stock indices for all our sample countries, respectively. While Panel III and IV use the samples of the developed countries, Panel V and VI perform the comparative study in emerging countries. We apply the tests for each model of interests. Looking at Table 2, we carry out three different panel unit roots tests, which include the LLC (Levin et al., 2002), the IPS (Im et al., 2003), the Fisher ADF (Maddala and Wu, 1999). The advantage of the IPS over the LLC is by allowing heterogeneity on the autoregressive coefficient, as well as different specifications of the parametric values, the residual variance and the lag lengths, while the Fisher-type test minimizes the size of distortions due to the cross-sectional correlations. As we can see, the results accept the null of unit roots for all the variables in the level form, whereas the variables are stationary in the first-difference form.

Table 2. Panel Unit Root Test

		Level Form			**First Difference**		
		Levin, Lin & Chut	**Im, Pesaran and Shin W-stat**	**ADF - Fisher Chi-square**	**Levin, Lin & Chut**	**Im, Pesaran and Shin W-stat**	**ADF - Fisher Chi-square**
Panel I	**Total Islamic**						
	$RSP_{i,t}$	-0.4758	2.98248	31.3985	-9.15343	-7.55398	219.215
		(0.3171)	(0.9986)	(1.0000)	(0.0000)*	(0.0000)*	(0.0000)*
	$RIP_{i,t}$	0.28912	0.75596	71.2188	-17.9748	-20.5144	568.794
		(0.6138)	(0.7752)	(0.5701)	(0.0000)*	(0.0000)*	(0.0000)*
	$RM_{i,t}$	1.78368	0.84376	82.0615	-29.0041	-31.3818	1043.36
		(0.9628)	(0.8006)	(0.2437)	(0.0000)*	(0.0000)*	(0.0000)*
	$CPI_{i,t}$	3.22677	3.21432	31.6621	-1.1043	-5.79984	157.316
		(0.9994)	(0.9993)	(0.9993)	(0.1347)	(0.0000)*	(0.0000)*
	$RIR_{i,t}$	5.81191	8.28088	22.4405	-6.82559	-5.91353	190.636
		(1.0000)	(1.0000)	(1.0000)	(0.0000)*	(0.0000)*	(0.0000)*
Panel II	**Total Conventional**						
	$RSP_{i,t}$	-0.96466	-0.64664	70.4143	-11.9352	-7.99459	230.784
		(0.1674)	(0.2589)	(0.5966)	(0.0000)*	(0.0000)*	(0.0000)*
	$RIP_{i,t}$	0.28912	0.75596	71.2188	-17.9748	-20.5144	568.794
		(0.6138)	(0.7752)	(0.5701)	(0.0000)*	(0.0000)*	(0.0000)*
	$RM_{i,t}$	1.78368	0.84376	82.0615	-29.0041	-31.3818	1043.36
		(0.9628)	(0.8006)	(0.2437)	(0.0000)*	(0.0000)*	(0.0000)*
	$CPI_{i,t}$	3.22677	3.21432	31.6621	-1.1043	-5.79984	157.316
		(0.9994)	(0.9993)	(0.9993)	(0.1347)	(0.0000)*	(0.0000)*
	$RIR_{i,t}$	5.81191	8.28088	22.4405	-6.82559	-5.91353	190.636
		(1.0000)	(1.0000)	(1.0000)	(0.0000)*	(0.0000)*	(0.0000)*
Panel III	**Developed Islamic**						
	$RSP_{i,t}$	-0.56543	1.83822	17.1778	-12.2133	-8.52824	165.455
		(0.2859)	(0.9670)	(0.9985)	(0.0000)*	(0.0000)*	(0.0000)*
	$RIP_{i,t}$	-0.84864	-0.65498	53.0345	-15.249	-16.511	324.574
		(0.1980)	(0.2562)	(0.0534)	(0.0000)*	(0.0000)*	(0.0000)*
	$RM_{i,t}$	1.75657	-0.16576	52.5988	-24.3245	-26.0796	585.000
		(0.9605)	(0.4342)	(0.0579)	(0.0000)*	(0.0000)*	(0.0000)*
	$CPI_{i,t}$	3.22039	2.36275	15.3922	3.63333	-2.97463	60.8550
		(0.9994)	(0.9909)	(0.9996)	(0.9999)	(0.0015)*	(0.0000)*
	$RIR_{i,t}$	5.48591	7.84685	9.42741	-4.59591	-4.03591	82.7107
		(1.0000)	(1.0000)	(1.0000)	(0.0000)*	(0.0000)*	(0.0000)*
Panel	**Developed**						

IV	**Conventional**						
	$RSP_{i,t}$	-0.65902 (0.2549)	2.48411 (0.9935)	13.1906 (0.9999)	-4.47741 (0.0000)*	-2.98338 (0.0000)*	86.1954 (0.0000)*
	$RIP_{i,t}$	-0.84864 (0.1980)	-0.65498 (0.2562)	53.0345 (0.0534)	-15.249 (0.0000)*	-16.511 (0.0000)*	324.574 (0.0000)*
	$RM_{i,t}$	1.75657 (0.9605)	-0.16576 (0.4342)	52.5988 (0.0579)	-24.3245 (0.0000)*	-26.0796 (0.0000)*	585.000 (0.0000)*
	$CPI_{i,t}$	3.22039 (0.9994)	2.36275 (0.9909)	15.3922 (0.9996)	3.63333 (0.9999)	-2.97463 (0.0015)*	60.8550 (0.0000)*
	$RIR_{i,t}$	5.48591 (1.0000)	7.84685 (1.0000)	9.42741 (1.0000)	-4.59591 (0.0000)*	-4.03591 (0.0000)*	82.7107 (0.0000)*
Panel V	**Emerging Islamic**						
	$RSP_{i,t}$	0.01299 (0.5052)	2.38635 (0.9915)	14.2207 (0.9996)	-3.78594 (0.0001)*	-2.7741 (0.0028)*	65.3281 (0.0020)*
	$RIP_{i,t}$	1.52998 (0.9370)	1.75420 (0.9603)	18.1843 (0.9941)	-10.1281 (0.0000)*	-12.4878 (0.0000)*	244.220 (0.0000)*
	$RM_{i,t}$	0.88196 (0.8111)	1.38046 (0.9163)	29.4627 (0.7712)	-16.4616 (0.0000)*	-18.352 (0.0000)*	458.358 (0.0000)*
	$CPI_{i,t}$	1.42250 (0.9226)	2.18094 (0.9854)	16.2699 (0.9981)	-4.67547 (0.0000)*	-5.25677 (0.0000)*	96.4609 (0.0000)*
	$RIR_{i,t}$	2.31262 (0.9896)	3.81361 (0.9999)	13.0131 (0.9998)	-5.54588 (0.0000)*	-4.32608 (0.0000)*	107.925 (0.0000)*
Panel VI	**Emerging Conventional**						
	$RSP_{i,t}$	-0.33468 (0.3689)	2.27054 (0.9884)	14.7304 (0.9994)	-2.89265 (0.0019)*	-2.45545 (0.0070)*	65.63 (0.0018)*
	$RIP_{i,t}$	1.52998 (0.9370)	1.75420 (0.9603)	18.1843 (0.9941)	-10.1281 (0.0000)*	-12.4878 (0.0000)*	244.220 (0.0000)*
	$RM_{i,t}$	0.88196 (0.8111)	1.38046 (0.9163)	29.4627 (0.7712)	-16.4616 (0.0000)*	-18.352 (0.0000)*	458.358 (0.0000)*
	$CPI_{i,t}$	1.42250 (0.9226)	2.18094 (0.9854)	16.2699 (0.9981)	-4.67547 (0.0000)*	-5.25677 (0.0000)*	96.4609 (0.0000)*
	$RIR_{i,t}$	2.31262 (0.9896)	3.81361 (0.9999)	13.0131 (0.9998)	-5.54588 (0.0000)*	-4.32608 (0.0000)*	107.925 (0.0000)*

Panel I uses Islamic stock indices in developed and emerging countries; Panel II uses conventional stock indices in developed and emerging countries; Panel III uses Islamic stock indices in developed countries; Panel IV uses conventional stock indices in developed countries; Panel V uses Islamic stock indices in emerging countries; Panel VI uses conventional stock indices in emerging countries.
RSP denotes the real Stock price/ RIP denotes the Real Industrial Production. RM denotes Real Money Supply (M1). CPI denotes the Consumer Price Index. RIR denotes the Real Short-term Interest Rate. *indicates significance at the 5% level.

Using these results, we proceed to test all variables in each model for co-integration in order to determine whether there is a long-run relationship. Firstly, we implement Pedroni co-integration test from the following equation:

$$RSP_{it} = \alpha_i + \gamma_{it} + \beta_{1i}RIP_{it} + \beta_{2i}RM_{it} + \beta_{3i}CPI_{it} + \beta_{4i}RIR_{it} + \varepsilon_{it}$$

where RSP is the real stock price; RIP is the real industrial production; CPI is the consumer price index; and RIR is the real short-term interest rate. This co-integration test allows for co-integrating vectors of differing magnitudes between countries, country (α) and time (γ) fixed effects. Looking at Table 3, the estimation results for each model demonstrate that, except for the panel v-statistic, ρ-statistic, and group ρ-statistic, all remaining statistics significantly reject the null of no co-integration. Hence it may imply that there is a long-run steady state relationship between all variables in each of our model after allowing for a country-specific effect. The next step is an estimation of such a relationship.

Table 3: Pedroni Residual Cointegration Test

	Panel I	Panel II	Panel III	Panel IV	Panel V	Panel VI
Panel v-Statistic	-0.567013 (0.7146)	0.027219 (0.4891)	-1.469127 (0.9291)	-0.692814 (0.7558)	0.825009 (0.2047)	0.874503 (0.1909)
Panel rho-Statistic	0.059964 (0.5239)	-0.819545 (0.2062)	0.165297 (0.5656)	-0.305601 (0.3800)	-0.098651 (0.4607)	-0.909931 (0.1814)
Panel PP-Statistic	-7.024253 (0.0000)*	-9.848777 (0.0000)*	-4.492425 (0.0000)*	-6.496249 (0.0000)*	-5.539752 (0.0000)*	-7.535807 (0.0000)*
Panel ADF-Statistic	-7.841582 (0.0000)*	-9.984421 (0.0000)*	-4.773458 (0.0000)*	-5.881715 (0.0000)*	-6.497795 (0.0000)*	-8.44948 (0.0000)*
Group rho-Statistic	0.525170 (0.7003)	0.087770 (0.5350)	0.889593 (0.8132)	0.564439 (0.7138)	-0.161023 (0.4360)	-0.454068 (0.3249)
Group PP-Statistic	-11.65844 (0.0000)*	-17.95677 (0.0000)*	-8.126557 (0.0000)*	-14.81904 (0.0000)*	-8.365706 (0.0000)*	-10.51988 (0.0000)*
Group ADF-Statistic	-10.72162 (0.0000)*	-13.11539 (0.0000)*	-6.887886 (0.0000)*	-8.878152 (0.0000)*	-8.29518 (0.0000)*	-9.682368 (0.0000)*

Panel I uses Islamic stock indices in developed and emerging countries; Panel II uses conventional stock indices in developed and emerging countries; Panel III uses Islamic stock indices in developed countries; Panel IV uses conventional stock indices in developed countries; Panel V uses Islamic stock indices in emerging countries; Panel VI uses conventional stock indices in emerging countries. *, **, *** indicates, significance at 1%, 5 % and 10% level respectively.

The below equation is estimated under pooled mean group and mean group estimators under the following error correction model:

$$\Delta RSP_{i,t-1} = \Phi_i\,(RSP_{i,t-1} - \Theta_{0i} - \Theta_{1i}\,RIP_{i,t-1} - \Theta_{2i}\,RM_{i,t-1} - \Theta_{3i}\,CPI_{i,t-1} - \Theta_{4i}\,RIR_{i,t-1}) + \beta_{11i}\Delta RIP_{i,t-1} + \beta_{21i}\,\Delta RM_{i,t-1} + \beta_{31i}\,\Delta CPI_{i,t-1} + \beta_{41i}\,\Delta RIR_{i,t-1} + dummycrisis_i + \xi_{i,t}$$

$$\text{where } \emptyset_i = -(1-\lambda_i), \theta_{0i} = \frac{\mu_i}{1-\lambda_i}, \theta_{it} = \frac{\delta_{10i}+\delta_{11i}}{1-\lambda_i}, \theta_{2i} = \frac{\delta_{20i}+\delta_{21i}}{1-\lambda_i}$$

Loayza and Ranciere (2006) mentioned that a common lag structure can be imposed across countries instead of using some consistent information criteria like Schwartz Bayesian criterion due to the limitation of the data. Hence we follow Pesaran et al. (1999) by parameterizing the error correction equation using ARDL (1,1,1). In addition, Pesaran et al. (1999) suggested that we may allow higher order lags by including lagged changes in explanatory variables. Our study then takes only one lagged changes of regressors without including current changes in order to avoid contemporaneous impact since it may involve endogeneity problem.

Looking at above equation, the pooled mean group and mean group estimates are presented as a two-equation model, which are the normalized co-integrating vector and the short-run dynamic coefficients, in order to uncover the long- and short-run consequences of our macroeconomic determinants on the stock price. The primary interest will include the error-correction speed of adjustment parameter, Φ_i, and the long-run coefficients, Θ_{1i} and Θ_{2i}, Θ_{3i}, and Θ_{4i}. By including Θ_{0i}, we allow a nonzero mean of the co-integrating relationship. We will expect Φ_i to have a negative sign so that the variables exhibit a return to a long-run equilibrium relationship. Since the PMG estimator put a constraint over the long-run coefficients to be equal across all countries, it will generate consistent and efficient estimates is this restriction is met. However, if the long-run coefficients are heterogeneous, MG estimates will be preferred as the PMG estimates are inconsistent. We apply the Hausman test to test the difference between the PMG and MG estimates where, for each of our six models, the results accept the null of no difference in these two models. Therefore, the PMG estimator is used in our study.

Table 4 presents the PMG estimator on different groups of countries according to our six models of interest. Focusing our analysis on the developed and emerging countries, we can see from the long-run coefficients, as our primary interest, that there are significant relationships between all macroeconomic variables and stock price for both Islamic and conventional equity indices. The signs of explanatory variables are also the same for these two indices albeit the magnitudes are different, thereby implying the similar characteristics of Islamic indices in response to the macroeconomic factors.

Table 4: Pooled Mean Group Estimation.

	I	II	III	IV	V	VI
Long Run coefficients						
RIP	0.3490322 (0.001)*	0.229346 (0.052)***	0.9526466 (0.000)*	0.9396849 (0.000)*	0.88485253 (0.000)*	0.236266 (0.079)***
RM	0.2400154 (0.048)**	0.5401298 (0.010)*	0.7817279 (0.000)*	0.7350334 (0.000)*	0.7252126 (0.009)*	1.516926 (0.000)*
CPI	1.206508 (0.000)*	1.274314 (0.000)*	0.5901437 (0.000)*	1.194756 (0.000)*	2.110102 (0.000)*	1.654602 (0.000)*
RIR	-0.056974 (0.000)*	-0.127835 (0.000)*	-0.062077 (0.011)**	-0.229212 (0.000)*	-0.49537 (0.044)**	-0.056663 (0.026)**
Short Run coefficients						
Error-correction coefficient	-0.280838 (0.000)*	-0.251654 (0.000)*	-0.257125 (0.000)*	-0.285048 (0.000)*	-0.3455792 (0.000)*	-0.334802 (0.000)*
ΔRIP	0.0441636 (0.711)	0.0355382 (0.732)	0.1115882 (0.511)	0.1227879 (0.456)	-0.2136897 (0.261)	-0.071891 (0.677)

ΔRM	0.0901835 (0.514)	0.1534615 (0.222)	0.1204007 (0.456)	0.1337529 (0.420)	-0.1719584 (0.394)	-0.234000 (0.230)
ΔCPI	1.552105 (0.001)*	1.491875 (0.000)*	0.970262 (0.055)***	1.129885 (0.015)**	0.5966971 (0.145)	0.9516804 (0.010)*
ΔRIR	0.1591704 (0.014)**	0.1321724 (0.031)**	0.2017334 (0.311)	0.125175 (0.378)	0.0225253 (0.666)	-0.034702 (0.345)
dummy crisis	-0.037521 (0.039)**	0.0063279 (0.671)	0.2017334 (0.311)	-0.033713 (0.006)*	-0.0481775 (0.015)**	-0.036467 (0.092)***
Intercept	-0.896746 (0.015)**	-1.781673 (0.000)*	-0.766693 (0.001)*	-1.751273 (0.000)*	-4.920332 (0.000)*	-6.279999 (0.000)*
Hausman Prob>chi2	0.975	0.3642	0.6318	0.9032	0.7309	0.0364

Model 1 is Total Islamic Indices. Model II is Total Conventional Indices. Model III is Emerging Markets Islamic Indices. Model IV is Emerging Market Conventional Indices. Model V is Developed Markets Islamic Indices. Model VI is Developed Markets Conventional Indices. *, **, *** indicates, significance at 1%, 5 % and 10% level respectively.

As to the impact of individual variable, the impact of the real industrial production is higher for Islamic as compared to conventional indices, whilst the response of Islamic indices to the short-term interest rate is significantly lower for these two groups of countries. This is understandable as the *Shariah* screening has set the certain low-threshold of interest-based debt in the capital structure. The firms included in the Islamic index therefore are those which depend more on the internal and external equity rather than fixed-income instruments. As a result, the compliant firms will not fully benefit from the lower interest rate by raising funds from the debt financing and, at the same time, they are not substantially negatively impacted during the higher interest rate regime. This will lead to the less exposure of these firms to the interest rate movement over time. On the other hand, the *Shariah* screening has removed a large number of highly-leveraged firms and especially those that operate under non-compliant business activity such as interest-based financial institutions and entertainment. This leads to Islamic index to be concentrated more on the production sector of the economy, particularly agriculture, manufacturing, oil and gas, telecommunication, technology, and so on. This segmentation may explain a higher dependence of Islamic stocks on the industrial production in the country.

While comparing the magnitude of the long-run coefficients between Islamic and conventional indices, the higher impact of industrial production on Islamic stocks is substantially greater in the developed countries as compared to, the emerging countries. This may indicate the closer link between Islamic stocks and the real production in the developed countries. Also, the less exposure of Islamic stocks to the interest rate is evident in the emerging countries, which is understandable since these countries generally do not have a well-functioning stock market. In the underdeveloped stock markets, even though a country's financial system becomes more sophisticated and credit becomes more available, the allocation of resources is still inefficient (misallocation of resources) (Chaiechi, 2012; Deidda and Fattouh, 2008; and others). The economy may have less alternative to raise the funds in response to an increase in the demand for output, hence it depends more on credit channel to promote growth. As a result, the non-Islamic firms in emerging countries may have a greater opportunity to grow by increasing leverage so that this will exhibit a substantially higher sensitivity to the interest rate movement. On the other hand, while analyzing money supply and consumer price index, we observe that the impact of the former on Islamic indices is higher in the emerging countries, whilst the later have higher impact on conventional indices in the developed countries.

Relating to the impact of the US-born subprime crisis, the dummy crisis demonstrate the stronger effect of the global crisis on Islamic, as compared to, conventional indices in both the developed and emerging countries. In the understandiong of the authors this is owing to the role of *Shariah* screening that have taken out the large non-compliant firms from the pool of investable equities, leaving the remaining Shari'ah compliant firms universe to be smaller and less diversified (Hussein & Omran, 2005). It is true from the theoretical underpinning that the lower leverage will imply a lower fixed financial commitment out of uncertain revenues, which consequently decrease the risk of the cash flow to equity (Hamada, 1972; Rubenstein, 1973; Christie, 1982; Mandelker & Rhee, 1984). It means that, during economic downturn, *Shariah*-compliant equities with a lower leverage theoretically will have lower systemic risk. The compliant firms also will be less volatile since leverage effect suggests that firms with lower debt/equity ratios should have a lesser negative relation between current returns and stock volatility (Black, 1976; Christie, 1982). Nonetheless, it seems that the disadvantage of smaller size, and less diversified, of Islamic firms may offset the advantage of lower leverage. Some prior studies mention that a relatively smaller size of a firm can lead to a higher systemic risk via size effect (Breen & Lerner, 1973; Kim et al., 2002; and so on). The size effect also plays an important role, where smaller firms are more exposed to a greater increase in their volatility, following a percentage fall in their stock price compared to those of larger firms (Black, 1976; Christie, 1982; Cheung & Ng, 1992).

When we observe the error-correction coefficient, the empirical results show that Islamic stocks tend to show a faster speed of adjustment to the equilibrium in the emerging countries. Notwithstanding the larger impact of the crisis on Islamic indices, this evidence may imply a higher stability of compliant firms during the recovery. This is in the opposite of what we found in the developed countries, whereby Islamic stocks portray a slightly slower adjustment as compared to the conventional ones. The plausible reason can be attributed to the nature of shocks absorption in the developed countries. For some countries, the shocks have involved the extreme severity of flight-to-quality created by self-fulfilling expectation along with the massive panicked deleveraging, and been followed by the frozen credit markets. This has led to a structural problem that substantially affects the real economy, where policy makers should not rely merely on monetary adjustment and fiscal stimulus to sustain. It is observed that the industrial production suffered more in the developed countries, thereby suggesting a greater and longer deviation from the equilibrium. Since this production variable has a substantial contribution to the Islamic indices within these countries, this evidence may explain the slower speed of adjustment of Islamic stocks to the long-run equilibrium relationship.

Finally, a proper analysis should be derived from the samples covering all the countries since the Hausman test accept the common long-run coefficients across different groups. We observe that the long-run coefficients are significant for both indices. The higher impact of the real industrial production, together with the lower impact of interest rate, on Islamic indices also remain in this model, thereby applying the same underlying reason that we described earlier. In addition, the impact of money supply is lower for Islamic stocks, which further suggest that compliant firms may have lower benefit from monetary expansion recalling the limit of taking higher financial leverage.

While observing the dummy crisis, the impact of the global crisis is only significant for Islamic indices. The finding may emphasize the major role of the size effect and less diversification on increasing the systemic risk as well as volatility of Islamic stocks during economic downturn. In addition the lower coefficient of error-correction term for Islamic indices may signify the relatively stability for these stocks to return to the long-run macroeconomic equilibrium relationship.

5. Conclusion

This study had set out to analyze and compare the impact of macroeconomic factors on the Islamic and Conventional equity prices. An insight into the behaviour of the equity prices would allow for a better understanding of how the Shariah screening for Islamic securities modify the relationship between the variables.

In line with our expectations, the Islamic indices are more grounded in the real sector owing to the focus of Islamic finance in the real part of the economy. The industrial production's impact is much higher for Islamic indices as compared to conventional indices while the response of Islamic indices to the short-term interest rate is significantly lower for these two groups of countries which can be attributed to the Shariah screening methodology as it sets a certain low-threshold of bearing the interest-based debt. The firms included in the Islamic index therefore are those which depend more on the internal and external equity rather than fixed-income instruments.

Also, there is evidence of a less exposure of Islamic stocks to money supply and consumer price index that can be evidenced in emerging countries. The results of this study provides some interesting insights and provides strong foundation for the argument that the Islamic equities are primarily impacted by changes in the real sector of the economy, and not highly dependent on the money supply and interest rates.

The impact of the recent crisis on Islamic indices is significant as compared to the conventional counterparts, but at the same time the adjustment to equilibrium is quicker for Islamic indices. This signifies the relative stability of the Islamic indices in the face of economic shocks. This study aimed to explore the behaviour of economic variables' impact on the Islamic indices in comparison with that on the conventional indices. There is a difference in which economic variables impact the equity prices of Shariah compliant stocks, which is an area that requires further research.

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