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Full Length Article

The impact of macroeconomic and conventional stock market variables on Islamic index returns under regime switching

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Abstract

The objective of this paper is to study the impact of conventional stock market return and volatility and various macroeconomic variables (including inflation rate, short-term interest rate, the slope of the yield curve and money supply) on Islamic stock markets returns for twenty developed and emerging markets using Markov switching regression models. The empirical results for the period 2002−2014 show that both developed and emerging Islamic stock indices are influenced by conventional stock indices returns and money supply for both the low and high volatility regimes. However, the other macroeconomic variables fail to explain the dynamics of Islamic stock indices especially in the high volatility regime. Similar conclusions are obtained by using the MS-VAR model.

JEL classification: G10; G12; C34

Keywords: Islamic index return; Conventional index return; Macroeconomic variables; Markov switching regressions; MS-VAR model

1. Introduction

During the last few decades, the world Islamic financial market has gathered significant momentum in attracting international capital flows from both Muslim and non-Muslim investors. Furthermore, the development of Islamic capital markets continues to show positive trends (Ho, Abd-Rahman, Yusuf, & Zamzamin, 2014). Islamic finance is guided by the principles of Islamic law (*Sharia*), which prohibits interest (*riba*), excessive risk-taking (*gharar*), gambling (*maysir*), and promotes risk-sharing, profit-sharing, asset-backed financial transactions, and ethical investment (Shamsuddin, 2014). One area of Islamic finance that has attracted many investors is the development of Islamic equity indices designed to track the performance of publicly traded Sharia-compliant companies. The Islamic indices are exposed to rigorous screenings for business activities and financial ratios and purification of dividends. Most major global banks and investment firms, both in Islamic and in non-Islamic countries, provide investors with the opportunity to invest in compliance with Sharia principles (Naifar, 2016). Islamic indices were launched for the first time in the late nineties.

The main objective of this paper is to study the impact of macroeconomic and conventional stock market variables on Islamic index returns across financial regimes over the period from June 2002 to June 2014 for ten developed and ten emerging markets which are respectively (Australia, Canada, France, Germany, Japan, Netherlands, Spain, Switzerland, UK and USA) on one side and (Chile, China, Czech Republic, India, Korea, Malaysia, Mexico, Russia, South Africa and Thailand) on the other. We also use Markov switching vector

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autoregressive model (MS-VAR) to study the causation effects from these earlier variables to stock market Islamic indices returns. As to generalize beyond a single country, models are put within a pooled time series framework for developed and emerging markets. We obtained this pooled data by stacking the data for each developed country and estimating one model for all these countries at once. The same is done for emerging markets.

Understanding the behaviour of Islamic equity markets under regime shift will help policy makers and international investors to be more effective in dealing with portfolio diversification strategies. The vast majority of empirical studies investigate the relationship between Islamic and conventional financial markets as well as between Islamic financial markets and macroeconomic variables by using linear models. For instance, Hammoudeh, Mensi, Reboredo, and Nguyen (2014) show that the Dow Jones Islamic Market index exhibits significant dependence with three major global conventional equity indices (Asia, Europe, and USA) and the global factors namely oil prices, stock market implied volatility, the U.S. 10-year Treasury bond interest rate, and the 10-year European Monetary Union government bond index. Prima Sakti and Harun (2013) found a co-integration between Islamic stock prices and macroeconomic variables specifically exchange rate, industrial production, inflation rate, and money supply. Ajmi, Hammoudeh, Nguyen, and Sarafrazi (2014) reveal a significant linear and nonlinear causality between the Islamic and conventional stock markets and between the Islamic stock market and interest rates. They affirm that the nonlinearity results are more credible because of the possible existence of structural breaks, asymmetry and regime switching in the markets and the relevant economic and financial variables.

Our study differs from previous research in several ways. First we explain within a coherent econometric framework the impact of conventional stock market return and volatility, the inflation rate, the short-term interest rate, the slope of the yield curve and the money supply M3 on Islamic index returns over a the recent period from June 2002 to June 2014. Second, given the multiplicity of crises and events that prevailed during the sample period such as the Lehman Brothers collapse (September 15, 2008) and the extreme market movements around the 2008—2009 global financial crisis, we use a Markov switching regression to allow the statistical significance of the selected variables to changes across financial market states. Third, the use of a sample composed of ten developed and ten emerging markets with a pooled analysis framework allows us to know if the heterogeneity of the markets affects the dynamics of Islamic stock indices.

The remainder of this paper is structured as follows. Section 2 provides a short review of the related literature. Section 3 describes the data. In Section 4, we estimate a linear regression model. Section 5 reports and discusses the empirical results of Markov switching regression. In Section 6, we present the Granger causality analysis. Section 7 concludes the article and deduces policy implications.

2. Literature review

The relationship between macroeconomic variables and stock market index has been widely investigated. First, several studies provide a consistent empirical explanation for index price/return-inflation relations across various countries. Differently to the commonly used economic theory which supposes that stock returns should be positively related to both expected and unexpected inflation, Fama and Schwert (1977), Fama (1981), Geske and Roll (1983), and Kaul (1987) among others document negative relations between the stock returns and both the expected and unexpected components of inflation rate. The negative relationship is due to a positive relationship between stock returns and future economic activity. In contrast, other studies reveal that inflation rate and stock market index returns can be positively correlated (Boudoukh & Richardson, 1993; Caporale & Jung, 1997; Choudhry, 2001; Hondroyiannis & Papapetrou, 2006). In fact, contrarily to Fama's view, they find that a positive relationship does exist and the negative effects of inflation on stock prices do not disappear after controlling for output shocks.

Second, the intuition regarding the relationship between interest rates and stock market index prices is well established, suggesting that an increase in interest rates increases the opportunity cost of holding money and thus substitution between stocks and interest bearing securities, and hence falling stock prices. Moreover, any changes in an asset's cash flows should have a direct impact on its price. Thus, the assets expected growth rates which influence its predicted cash flows will affect its price in the same direction (Nishat & Shaheen, 2004). In theory, the relationship between interest rates and stock prices is negative. This is due to the cash flow discounting model according to which, present values of stocks are calculated by discounting the future cash flows at a discount rate. If the discount rate increases, then present values of stocks decline and vice versa. In addition, the rising interest rates reduce cash flows by reducing the profitability of firms. Due to these two reasons, present values of stocks decline and so do current (Panda, 2008). Several studies show that there is a negative correlation between the stock prices and short-long-term interest rates (Chen, Roll, & Ross, 1986; Koch & Saporoschenko, 2001; Shiller & Beltratti, 1992). This implies that the considered stock returns often exhibit significant negative responses to interest rate increases. Differently, Rigobon and Sack (2004) and Bohl, Siklos, and Werner (2003) reveal that stock market movements have a significant impact on interest rates, driving them in the same direction as the change in stock prices. Recently, Martínez, Lapena, and Sotos (2015) exhibit significant interest rate sensitivity, although the degree of interest rate exposure differs considerably across industries and depending on the time horizon under consideration.

Third, monetary policy has been shown to be an important element in the phase of economic cycles (Ghazali & Yakob, 1998). The changes in money supply could have a significant impact on stock prices (Friedman & Schwartz, 1963) and can be explained in two frameworks. From the perspective of the dividend valuation model, Keran (1971), Hamburger and

Kochin (1972) and Homa and Jaffee (1971) reveal that the changes in money supply cause investors to adjust the factors that determine stock prices (the risk-free yield, earnings expectations and risk premium). This adjustment alters the demand for other assets that compete with money balance. An increase in money supply is expected to create excess supply of money balances and, in turn, excess demand for stocks. Taking into account the effects of monetary expansion on the dividend discount model, monetary supply has a direct relationship with stock prices. From the perspective of the Monetary portfolio model developed by Brunner (1961), Friedman (1961), Friedman and Schwartz (1963) and Cagan (1974), a monetary disturbance causes investors to adjust their portfolio holding and thus affecting asset prices. The relationship between money supply aggregate and stock price can exist either positively or negatively. Referring to Fama (1981), Dhakal, Kandil, and Sharma (1993) and Mukherjee and Naka (1995), since the rate of inflation is positively related to money growth rates, an increase in the money supply may lead to an increase in the discount rate. The negative effects may be countered by the economic stimulus provided by money growth. Such stimulus would likely result in increased future cash flows and stock prices.

Despite the increasing attention to Islamic financial markets, the vast majority of empirical studies investigate the performance of Islamic equity indices and mutual funds. Hayat and Kraeussl (2011) analyse the risk and return characteristics of a sample of 145 Islamic equity funds (IEFs) over the period from 2000 to 2009 and show that IEFs are underperformers compared to Islamic as well as to conventional equity benchmarks. Using Sharpe ratio, Treynor index and Jensen alpha, Ho et al. (2014) evaluate the risk-adjusted performances of twelve global conventional and Islamic indices and compare their performances during various crisis and non-crisis periods. Their findings reveal that Islamic indices outperformed their conventional counterparts during crisis periods but results are inconclusive for the non-crises periods. Jawadi, Jawadi, and Louhichi (2014) study the financial performance of Islamic and conventional indices for three major regions: Europe, the USA and the World over the period 2000—2011 and show that, while conventional investments seemed promising before the crisis and during periods of calm, Islamic funds have outperformed them since the subprime crisis began and in turbulent times.

Some recent studies on Islamic finance have dealt with the links between Islamic financial markets and macroeconomic variables as well as between Islamic and conventional financial markets. Mohd. Yusof and Abdul Majid (2007) study the link between the monetary policy volatilities and the volatility of stock returns in both conventional and Islamic stock markets in Malaysia during the period from January 1992 to December 2000. They show that the interest rate volatility affects the conventional stock market volatility but does not have, on the contrarily any effect on the Islamic stock market volatility. They found, however, that for Islamic stock market volatility, the exchange rate is the important indicator of economic instability. Using GARCH models, Albaity (2011) investigates the effect of monetary policy, interest rate and the rate of inflation on the Islamic stock market indices in Malaysia (Kuala Lumpur Syariah Index) and the USA (Dow Jones Islamic Market Index) over the period from April 1999 to December 2007. He finds that in the univariate models, the variance of the two indices is influenced by money supply (M2 and M3) and inflation rate. In addition, in the multivariate model, Dow Jones Islamic Market Index is influenced by the interest rate and the inflation rate in the mean and variance equations. In contrast, Kuala Lumpur Syariah Index is influenced commonly in the mean and variance equations by money supply M3, and the inflation rate.

Abdul Majid and Mohd. Yusof (2009) use the autoregessive distributed lag methodology to explore the extent to which macroeconomic variables such as industrial production index, money supply, real effective exchange rate, interest rate and federal fund rates affect the Islamic stock market behaviour in Malaysia in the post 1997 financial crisis period. They find that, except for the industrial production index, all the macroeconomic variables analysed, have a significant impact on the Islamic stock market returns. Shamsuddin (2014) studies the sensitivity of Dow Jones Islamic equity portfolios to interest rate over the period from January 5, 1996 to April 8, 2011 and reveals that the aggregate portfolio of Islamic stocks is immune to interest rate risk. However, at the sectorial level some Islamic equity portfolios demonstrate exposure to interest rate risk. Wahyudi and Sani (2014) investigate the Toda and Yamamoto (1995) VAR causality relationship between macroeconomic variables and Islamic financial market over the period 2002—2011 and found that the Jakarta Islamic Index is affected by the exchange rate of Rupiah against the U.S. dollar, but not by interest rate.

Hussin, Muhammad, Abu Hussin, and Abdul Razak (2012) use the vector autoregression (VAR) methodology to examine the relationship between exchange rates, oil prices, and the Islamic capital market in Malaysia over the period of January 2007—December 2011. They found that the Islamic share prices had a positive and significant correlation with oil prices, but they had a negative and insignificant relationship with the exchange rate of Malaysian Ringgit to the U.S. dollar. Furthermore, in the short run, only oil price variable Granger cause for Islamic stock returns; this proves that only oil prices could predict short-term sharia stock returns. Prima Sakti and Harun (2013) use the time series techniques of cointegration and VAR to analyse the short-run and long-run relationship between Jakarta Stock Exchange Islamic Index and selected macroeconomic variables; namely exchange rate, industrial production, inflation rate, and money supply from 2000 to 2010. They prove that there is co-integration between Islamic stock prices and all the macroeconomic variables considered.

Albaity and Ahmad (2008) employ the causality and Johansen cointegration tests to examine the short- and long-run relationships between the Kuala Lumpur Syariah Index (KLSI) and the Kuala Lumpur Composite Index (KLCI) during the period 1999—2005. They found that the short-run

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causality, measured by the Granger bivariate test, points to the bidirectional causality between the two indices. This implies that in the short-run both prices move in the same direction and tend to cause each other. Moreover, the long-term relationship indicates that KLCI is adjusting to its long-term equilibrium while KLSI is not. Therefore, predicting one of them on the basis of the other is constructive. Ajmi et al. (2014) use the heteroscedasticity-robust linear Granger causality and nonlinear Granger causality tests to examine the links between the Islamic and global conventional stock markets and between the Islamic stock market and several global economic and financial shocks over the period from January 4, 1999 to October 8, 2010. They reveal a significant linear and nonlinear causality between the Islamic and conventional stock markets but more especially from the Islamic stock market to the other markets where they found that causality is the strongest. They also show a connection between the Islamic stock market and interest rates.

Using a copula approach, Hammoudeh et al. (2014) reveal significant upper and lower tail dependence between the Dow Jones Islamic Market index and the three major global conventional equity indices namely; Asia, Europe and USA and the global factors (oil prices, stock market implied volatility, the U.S. 10-year Treasury bond interest rate, and the 10-year European Monetary Union government bond index). Moreover, this dependence varies over time for all cases except with the S&P 500 index. It is also asymmetric between bear and bull markets in some cases.

Recently Naifar (2016) use quantile regression approach to investigate the co-movement and the dependence structure between Dow Jones Islamic Market Index returns and influential global financial market conditions, macroeconomic indicators and risk factors. He shows that conventional stock market returns, stock market implied volatility and the slope of the yield curve are significant for all the quantiles and display asymmetric tail dependence. The sovereign credit risk factor has also been significant during and after the global financial crisis. Moreover, the impact of oil prices and investor sentiment indicator is positive and significant only for the lower quantiles.

In light of previous researches, our study contributes to the literature by examining the regime switching behaviour of Islamic equity indices in order to known if this behaviour differs from that of conventional indices. Besides, it analyzes the impact of different economic factors on Islamic stock market return across high and low volatility regimes. It is worth noting that most of the current empirical studies have examined the links between the Islamic equity index and conventional counterparts and macroeconomic variables by using linear models.

3. Data and descriptive statistics

The data for our study include ten developed market indices (Australia, Canada, France, Germany, Japan, Netherlands, Spain, Switzerland, UK and USA) and ten emerging markets indices (Chile, China, Czech Republic, India, Korea, Malaysia, Mexico, Russia, South Africa and Thailand). It covers the period from June 2002 to June 2014. These countries are selected according to the importance of their stock market capitalization and to the data disponibility. For the macroeconomic variables, we use the consumer price index, 3-months Treasury bill rates, 10-years government bond yields and Money supply (M3). Monthly indices returns, inflation rate, changes in short-term interest rate and changes in Money supply are computed on a continuous basis as the difference in logarithm between two consecutive observations. The data is compiled from two different sources. Data for the Islamic indices and conventional counterparts are obtained from Morgan Stanley Capital International (MSCI) and the data for macroeconomic variables are obtained from Trading Economics and DataStream databases.

Following Sharia investment principles, MSCI excludes securities using two types of criteria. The first set removes any companies which are directly active in, or derive more than 5% of their revenue from alcohol, tobacco, pork-related products, Gambling, Music, hotels, cinema and conventional financial services. The second set of screens utilizes financial ratios to remove companies based on debt and interest income levels. Fig. 1 displays the monthly indices prices for developed and emerging markets during the sample period.

The Conditional variance of conventional stock market index returns is determined by an AR (P) GARCH (1, 1) model specified as follows:

$$r_t = a_0 + \sum_{s=1}^{p} a_s r_{t-s} + \varepsilon_t; \tag{1}$$

$$\varepsilon_t / \Omega_{t-1} \sim N(0, h_t); \tag{2}$$

$$h_t = \omega_0 + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1}. \tag{3}$$

where r_t is the monthly conventional index return at time (t), h_t is the conditional variance of the residuals from the mean equation andε_t is the error term that follows a normal distribution with mean zero and time-varying variance. As we use pooled data analysis, we calculate descriptive statistics for two series of data; for developed and for emerging markets. Table 1 provides summary statistics for the data.

Table 1 shows that, for developed markets, Islamic indices present a monthly mean return and standard deviation which is more important than their conventional counterpart. For emerging markets, Islamic indices are less profitable and more risky than the conventional ones. The mean of the conditional variance of conventional index returns is equal to 0.0040 and 0.0059 for developed and emerging markets, respectively. Developed markets have the least important mean of inflation. The changes in short-term interest rate have a negative mean for developed and emerging markets. The mean of the changes in the slope of the yield curve is negative for emerging markets and positive for developed ones. The mean of the changes in money supply is less important in developed markets. The Jarque–Bera test shows that the null

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Fig. 1 Stock Market prices of Islamic and conventional indices (developed markets).

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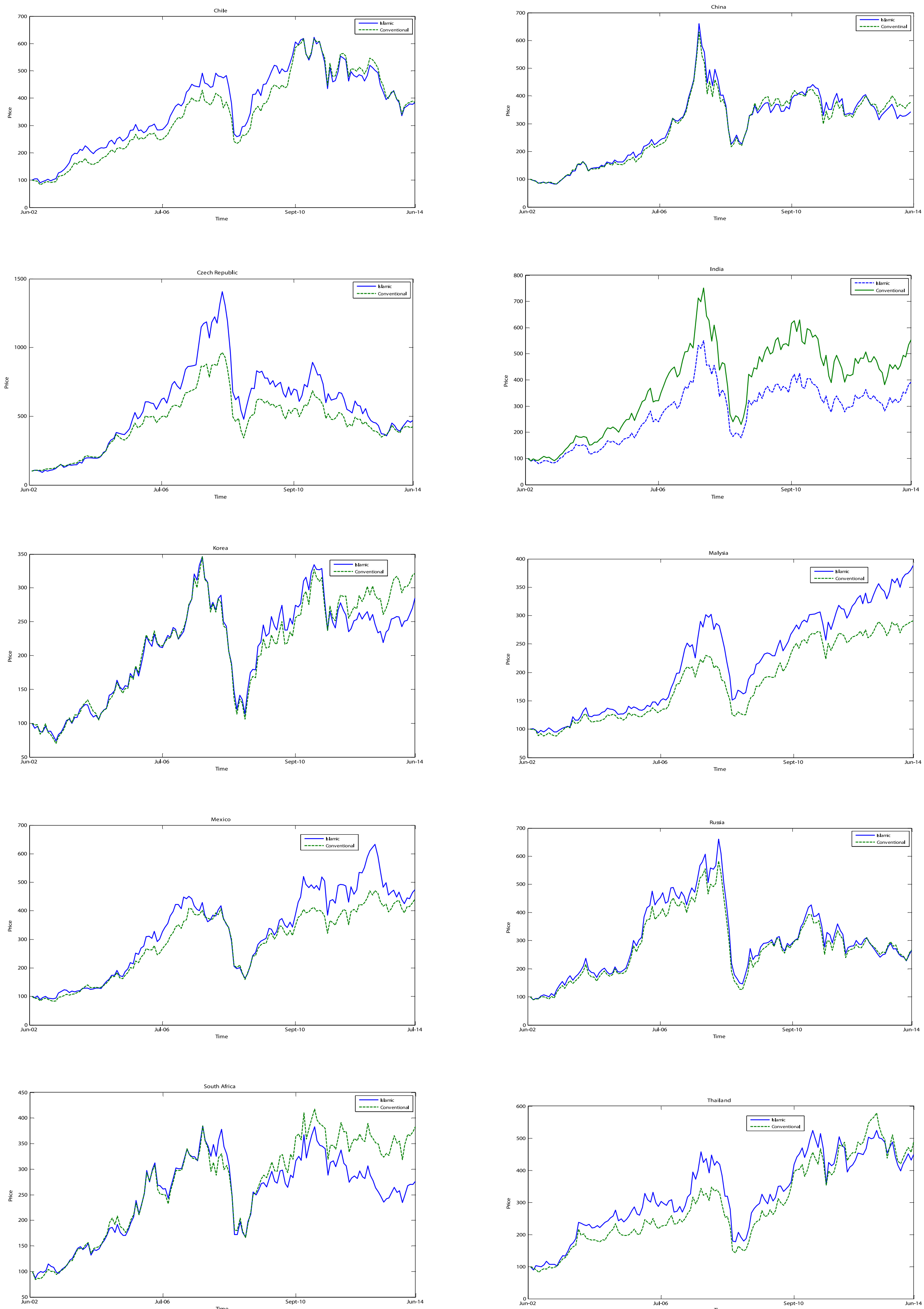

Fig. 1. (*continued*).

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Table 1
Descriptive statistics.

	RISLAM	*RCONV*	*CVAR*	*INF*	*ΔINT*	*ΔSYC*	*ΔMS*
A. Descriptive statistics							
a. Developed markets							
Mean	0.0061	0.0049	0.0040	0.0014	−0.0173	0.00007	0.0046
Std. Dev.	0.0624	0.0619	0.0044	0.0038	0.2385	0.0026	0.0094
Jarque-Bera	1001** (0.00)	728** (0.00)	17503** (0.00)	194** (0.00)	365334** (0.00)	31419** (0.00)	25102** (0.00)
ADF	−34.46** (0.00)	−33.73** (0.00)	−9.90** (0.00)	−4.88** (0.00)	−25.33** (0.00)	−34.14 (0.00)	−6.40** (0.00)
b. Emerging markets							
Mean	0.0092	0.0097	0.0059	0.0037	−0.0025	−0.00002	0.0099
Std. Dev.	0.0812	0.0783	0.0050	0.0060	0.0691	0.0050	0.0098
Jarque-Bera	650** (0.00)	648.10** (0.00)	175681** (0.00)	2056** (0.00)	30766** (0.00)	77677** (0.00)	732** (0.00)
ADF	−33.74** (0.00)	−33.07** (0.00)	−11.24** (0.00)	−3.94** (0.00)	−17.59** (0.00)	−35.35** (0.00)	−456** (0.00)
B. Correlation matrix							
a. Developed markets							
RISLAM	1						
RCONV	0.94	1					
CVAR	−0.04	−0.06	1				
INF	0.06	0.04	−0.10	1			
ΔINT	0.05	0.05	−0.14	0.11	1		
ΔSYC	−0.08	−0.08	0.27	−0.05	−0.24	1	
ΔMS	−0.03	−0.06	−0.05	0.04	−0.01	−0.02	1
b. Emerging markets							
RISLAM	1						
RCONV	0.95	1					
CVAR	−0.07	−0.08	1				
INF	−0.04	−0.04	0.06	1			
ΔINT	−0.11	−0.10	0.01	0.06	1		
ΔSYC	0.10	0.08	−0.08	0.02	−0.59	1	
ΔMS	0.03	0.03	0.07	0.08	0.01	−0.02	1

Note: The sample period is June 2002 through June 2014. *RISLAM* is the monthly Islamic index return; *RCONV* is the monthly conventional index return; *CVAR* is the conditional variance of conventional stock market index returns; *INF* is the inflation rate; *ΔINT* is the changes in short-term interest rate; *ΔSYC* is the first difference of the slope of the yield curve and *ΔMS* is the first difference of money supply. p-values are in parentheses.
** and * indicate significance at the 5% and 10% levels, respectively.

hypothesis of normality is rejected for all series. The ADF test with drift and trend was conducted to check for unit root in all the series. All variables are stationary and the null hypothesis of unit root is rejected.

4. Linear regression analysis

In this section, we use pooled OLS technique to estimate the following linear regression model for developed and emerging markets in our sample:

$$RISLAM_t = \beta_0 + \beta_1 RCONV + \beta_2 CVAR + \beta_3 INF_t + \beta_4 \Delta INT_t + \beta_5 \Delta SYC_t + \beta_6 \Delta MS_t + \varepsilon_t \tag{4}$$

where $(RISLAM)$ is the monthly Islamic index return, $(RCONV)$ is the monthly conventional index return, $(CVAR)$ is the conditional variance of conventional stock market index returns, (INF) is the monthly inflation rate, (ΔINT) is the changes in short-term interest rate, (ΔSYC) is the first difference of the slope of the yield curve, and (ΔMS) is the changes in money supply. The results are given in Table 2.

Tale 2 shows that, except for the interest rate and the slope of the yield curve, all variables have significant effects on Islamic index return in developed markets. For emerging countries, only the conventional index return and the slope of the yield curve have a significant effect. In order to understand how well such a linear regression relation performs, we use both the Ljung-Box Q statistic test on the squared residuals and the cumulative sum of squares of recursive residuals (CUSUMSQ) test as proposed by Brown, Durbin, and Evans (1975). Fig. 2 plots the CUSUMSQ test statistics.

Table 2
Linear regression results.

	Developed markets	Emerging markets
β_0	−0.0005 (0.55)	−0.0007 (0.54)
RCONV	0.9455** (0.00)	0.9864** (0.00)
CVAR	0.2491* (0.07)	0.0109 (0.93)
INF	0.2665* (0.08)	−0.0199 (0.85)
ΔINT	−0.0008 (0.75)	−0.0012 (0.91)
ΔSYC	−0.0541 (0.81)	0.3747** (0.02)
ΔMS	0.1212** (0.04)	0.0380 (0.57)
$Log(L)$	3466.09	3302.09
$Q(12)$	25.44 (0.01)	31.01 (0.00)
$Q^2(12)$	984.1 (0.00)	416.92 (0.00)

Note: Q (12) and Q^2 (12) are the Ljung-Box statistics for serial correlation in the model residuals and squared residuals, respectively, computed with 12 lags. P-values are in parentheses. ** and * indicate significance at the 5% and 10% levels. respectively.

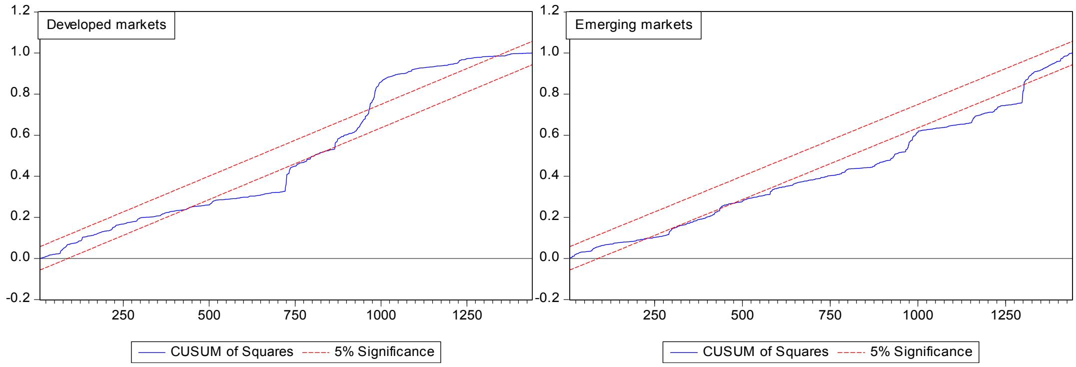

Fig. 2. CUSUM of squares test results.

Fig. 2 shows clearly that the linear model is incapable for addressing either the parameter instability or the variance instability over the sample period for developed and emerging countries. We conclude that Islamic indices may have a regime specific behaviour.

5. Markov regression models

Markov switching regression models allow the influence of explanatory variables to be state-dependent. Within model (4) this approach allows the regression parameters β_1 to change over time according to a particular transition probability and β_i can take different values depending on the market regime or 'state' at time t, which is denoted by S_t. The transition from one state to another is described by an unobservable Markov chain, i.e.

$$RISLAM_t = \beta_0 + \beta_{1,S_t} RCONV + \beta_{2,S_t} CVAR + \beta_{3,S_t} INF_t + \beta_{4,S_t} \Delta INT_t + \beta_{5,S_t} \Delta SYC_t + \beta_{6,S_t} \Delta MS_t + \varepsilon_{t,S_t} \quad (5)$$

Markov switching regressions go back to Goldfeld and Quandt (1973). The formulation used here is due to Hamilton (1989, 1994) where S_t is an unobservable variable governed by a first order Markov process. S_t is assumed to be a two-state first order Markov process with transition probability matrix represented as:

$$P = \begin{bmatrix} P_{11} & P_{21} \\ P_{12} & P_{22} \end{bmatrix} \quad (6)$$

where $p_{ij} = \Pr(S_t = j | S_{t-1} = i)$ and $\sum_{j=1}^{2} p_{ij} = 1$, for $j = 1,2$, and for all i. These probabilities are specified as constant coefficients that are independent of time t.

Table 3
Markov switching regression results.

	Developed markets		Emerging markets	
	Conventional stock market variables	Macroeconomic variables	Conventional stock market variables	Macroeconomic variables
β_0	0.0006 (0.32)	0.0093** (0.00)	0.0004 (0.59)	0.0129** (0.00)
Regime 1				
RCONV	0.9418** (0.00)	–	0.9744** (0.00)	–
CVAR	0.2089* (0.07)	–	−0.1082 (0.28)	–
INF	–	0.7989** (0.03)	–	−0.6670* (0.06)
ΔINT	–	0.0062 (0.22)	–	−0.0756** (0.03)
ΔSYC	–	−1.6702** (0.00)	–	−0.5988 (0.29)
ΔMS	–	0.4551** (0.00)	–	0.3745* (0.06)
σ^2	0.0002** (0.00)	0.0014** (0.00)	0.0003** (0.00)	0.0042** (0.00)
Regime 2				
RCONV	0.9490** (0.00)	–	1.0118** (0.00)	–
CVAR	0.2095 (0.21)	–	−0.0300 (0.87)	–
INF	–	0.3252 (0.75)	–	−1.1651 (0.38)
ΔINT	–	0.0064 (0.82)	–	−0.1944 (0.34
ΔSYC	–	−1.6216 (0.16)	–	2.9696 (0.11)
ΔMS	–	−1.6726** (0.00)	–	−1.4832** (0.05)
σ^2	0.0013** (0.00)	0.0081** (0.00)	0.0013** (0.00)	0.0196** (0.00)
p_{11}	0.98** (0.00)	0.96** (0.00)	0.98** (0.00)	0.98** (0.00)
p_{22}	0.94** (0.00)	0.93** (0.00)	0.95** (0.00)	0.86** (0.00)
$E(d_1)$	46.24	26.03	44.74	46.73
$E(d_2)$	17.58	13.83	21.21	7.22
$Log(L)$	3717.51	2130.75	3454.05	1674.75

Note: $E(d_1)$ and $E(d_2)$ are respectively the expected duration of the low-volatility regime and the high volatility one. P-values are in parentheses. ** and * indicate significance at the 5% and 10% levels, respectively.

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The Markov switching model is estimated using a maximum likelihood procedure.[1] Under the assumption of Gaussian $\varepsilon_{i,t}$ for both states i, the conditional density of y_t is represented as:

$$\eta_{i,t} = f(y_t|S_t = i, \psi_{t-1};\theta) = \frac{1}{\sqrt{2\pi\sigma_i^2}}\exp\left(-\frac{\left(y_t - \beta_0 - X_t^{'}\beta_i\right)^2}{2\sigma_i^2}\right) \quad (7)$$

where ψ_{t-1} refers to information up to time *t-1*. and θ is the vector of parameters $(\beta_0, \beta_{1,S_t}, \beta_{2,S_t}, \beta_{3,S_t}, \beta_{4,S_t}, \beta_{5,S_t}, \beta_{6,S_t}, \sigma^2_{S_t}, P_{11}, P_{22})$ to be estimated. The joint density of y_t and S_t is the product of the conditional and marginal densities:

$$f(y_t, S_t|\psi_{t-1};\theta) = f(y_t|S_t, \psi_{t-1};\theta) f(S_t|\psi_{t-1}) \quad (8)$$

Summing up all possible values for S_t leads to

$$f(y_t|\psi_{t-1};\theta) = \sum_{S_t=1}^{2} f(y_t|S_t, \psi_{t-1};\theta) f(S_t|\psi_{t-1}) \quad (9)$$

As it is impossible to know for sure what regime the Islamic financial market is in, inference about the regime is made by observing the Islamic index returns. The inference comes in the form of filtered probabilities$(\xi_{j,t})$, which are computed recursively using historical information, ψ_{t-1}:

$$\xi_{j,t} = \Pr(S_t = j|\psi_t;\theta) = \frac{\sum_{i=1}^{2} p_{ij}\xi_{i,t-1}\eta_{jt}}{f(y_t|\psi_{t-1};\theta)} \quad (10)$$

The set of optimal parameters $\widehat{\theta}$ can be estimated by maximizing the following log-likelihood function:

$$\ell(\theta) = \sum_{t=1}^{T} \log f(y_t|\psi_{t-1};\theta) \quad (11)$$

The average duration of each regime can be obtained from the transition probability$P_{jj}(j = 1, 2)$ as follows:

$$D_{jj} = 1/\left(1 - P_{jj}\right) \quad (12)$$

The estimated coefficients β_{jS_t} with$S_t \in \{1, 2\}$ and $j = 1, \ldots, 6$ are displayed in Table 3.[2]

Results from Table 3 show that the coefficient estimates are more significant especially in the low volatility regime. All variables, except conventional stock market return and money supply have no impact on Islamic index return in the high volatility regime. This may be explained by the fact that Islamic index returns can be explained more by investors' behaviour in the high volatility period than by economic fundamentals of a country. Looking first at conventional stock market variables, the coefficient of $(RCONV)$ is positive and significant in all regimes and for both developed and emerging markets. As noted by Naifar (2016), this result can be explained by the fact that stocks of Islamic indices are a subset of all the stocks listed on the main board (excluding prohibited sectors) and the benchmarks from which Islamic indices are selected are well-recognized conventional indices. Besides, all Islamic indices follow a stock selection process that is called stock screening regarding the scope of activities of firms and their financial ratios. For $(CVAR)$, the impact is positive and significant only in the first regime and for developed countries. For inflation variable, the coefficient is significant only in the first regime, where an increase in the inflation rate leads to higher Islamic index return in developed countries and to lower those in emerging countries. Similarly to the common used economic theory, the positive relationship of inflation and stock market can be explained through the interest rates. If the inflation rate is high, the interest rate is also high. Therefore, Sharia prohibits acceptance of specific interest or fees for loans, which plays a significant role in prohibiting funds from being invested in debt markets. The estimated coefficient of changes in short term-interest rate has a significant and a negative effect only for emerging markets in the low volatility regime. As noted by Albaity (2011), interest rate is hypothesized to have a negative relationship with stock returns. Based on the stock valuation model, whenever interest rate increases the value of cash flow is worth less after discounting. Therefore, the incentive to invest shrinks as well as the stock return. Interest rates can influence the level of corporate profits which in turn influence the price that investors are willing to pay for the stock through expectations of higher future dividend payment. Besides as substantial amount of stocks are purchased with borrowed money, an increase in interest rates would make stock transactions more costly. Investors will require a higher rate of return before investing. This will reduce demand and lead to a price depreciation. As advanced by Kia (2015), the existence of predetermined interest rate results in a reduction of stock price, the market value of companies, for a given number of shares in circulation and outstanding corporate debt. Concerning the first difference of the slope of the yield curve, the estimated coefficient is negative and significant only for developed markets in the first regime. We confirm the findings of Dotsey (1998) that the explanatory power of the slope of the yield curve for the future economic growth is stronger in periods of price instability. An increase in the changes in the slope of the yield curve has a different impact on Islamic index return according to countries and regimes. In the low volatility regime, there is a statistically significant negative relationship between the first difference of slope of the yield curve and Islamic stock market index return for Germany, Malaysia and South Africa. As the slope of the yield curve can be a proxy of the returns on long duration bonds (Zhang, Hopkins, Satchell, & Schwob, 2009), this negative relation can be explained as proposed by Mukherjee and Naka (1995), by the fact that long-term interest rates seem to be a better proxy for the nominal risk-free component of the discount rate in stock valuation models. The positive relationship is explained by the fact that the slope of the yield curve has significant information content for economic growth (Paya,

[1] See Hamilton (1994) and Kim and Nelson (1999) for more details on the estimation method.

[2] To limit the number of parameters in the Markov switching regression, estimates are done separately for conventional stock market and economic variables.

Table 4a
Bivariate MS-VAR results for developed markets.

	$RISLM_t$	$RISLM_t$	$RISLM_t$	$RISLM_t$	$RISLM_t$	$RISLM_t$
Regime 1						
μ_1	0.026** (0.00)	0.019 (1.00)	0.016** (0.00)	0.013** (0.00)	0.014** (0.00)	0.012** (0.00)
$RISLM_{t-1}$	−0.058 (0.49)	0.170** (0.00)	−0.017 (0.72)	−0.037 (0.23)	−0.057* (0.08)	−0.060 (0.13)
$RISLM_{t-2}$	−0.340** (0.00)	–	–	–	–	–
$RCONV_{t-1}$	−0.167** (0.05)	–	–	–	–	–
$RCONV_{t-2}$	0.202** (0.04)	–	–	–	–	–
$CVAR_{t-1}$	–	−0.137 (1.00)	–	–	–	–
$CVAR_{t-2}$	–	–	–	–	–	–
INF_{t-1}	–	–	−0.98** (0.03)	–	–	–
INF_{t-2}	–	–	–	–	–	–
ΔINT_{t-1}	–	–	–	−0.001 (0.99)	–	–
ΔINT_{t-1}	–	–	–	–	–	–
ΔSYC_{t-1}	–	–	–	–	−0.206 (0.74)	–
ΔSYC_{t-2}	–	–	–	–	–	–
ΔMS_{t-1}	–	–	–	–	–	0.001 (0.87)
ΔMS_{t-2}	–	–	–	–	–	–
σ_1^2	0.001** (0.00)	0.002** (0.00)	0.001** (0.00)	0.003** (0.00)	0.002** (0.00)	0.002** (0.00)
Regime 2						
μ_2	−0.027** (0.00)	0.003 (0.72)	−0.008* (0.08)	−0.014** (0.01)	−0.023** (0.00)	−0.001 (0.99)
$RISLM_{t-1}$	0.396** (0.01)	−0.172** (0.00)	0.086* (0.07)	0.156** (0.02)	0.166** (0.03)	0.150** (0.00)
$RISLM_{t-2}$	0.370** (0.03)	–	–	–	–	–
$RCONV_{t-1}$	−0.205 (0.19)	–	–	–	–	–
$RCONV_{t-2}$	−0.323** (0.04)	–	–	–	–	–
$CVAR_{t-1}$	–	1.073 (0.19)	–	–	–	–
$CVAR_{t-2}$	–	–	–	–	–	–
INF_{t-1}	–	–	0.189 (0.82)	–	–	–
INF_{t-2}	–	–	–	–	–	–
ΔINT_{t-1}	–	–	–	0.008 (0.36)	–	–
ΔINT_{t-1}	–	–	–	–	–	–
ΔSYC_{t-1}	–	–	–	–	0.029 (0.87)	–
ΔSYC_{t-2}	–	–	–	–	–	–
ΔMS_{t-1}	–	–	–	–	–	0.002 (0.83)
ΔMS_{t-2}	–	–	–	–	–	–
σ_2^2	0.007** (0.00)	0.009** (0.00)	0.008** (0.00)	0.008** (0.00)	0.011** (0.00)	0.006** (0.00)
p_{11}	0.81** (0.00)	0.94** (0.00)	0.96** (0.00)	0.95** (0.00)	0.95** (0.00)	0.95** (0.00)
p_{22}	0.62** (0.00)	0.82** (0.00)	0.92** (0.00)	0.81** (0.00)	0.79** (0.00)	0.93** (0.00)
$E(d_1)$	5.20	15.49	24.60	18.89	21.69	19.34
$E(d_2)$	2.60	5.70	13.18	5.25	4.81	13.39
$Log(L)$	4475.45	10124.31	8146.56	3507.36	8959.32	6992.15

Note: $E(d_1)$ and $E(d_2)$ are respectively the expected duration of the low-volatility regime and the high volatility one. P-values are in parentheses. ** and * indicate significance at the 5% and 10% levels, respectively.

Matthews, & Peel, 2005). Concerning the first difference of the money supply, the estimated coefficient is significant in all regimes and for developed and emerging markets. An increase in the changes in the money supply leads to increase or decrease Islamic index return in the low and high volatility, respectively. The negative relationship can be explained, as noted by Fama (1981), by the fact that an increase in money supply would lead to inflation, and may increase discount rate and reduce stock prices. The negative effects might be countered by the economic stimulus provided by money growth, also known as the corporate earnings effect, which may increase future cash flows and stock prices. Nevertheless, a possible explanation for the positive relationship might be the government's active role in preventing prices escalation. The hold of stocks might be an effective hedge against inflation and hence the "Fisher effect" would explain this positive correlation. As concluding by Marshall (1992), if inflation is caused by money shock, it would lower the rate of interest and investors would shift their cash holdings to stocks and bonds in order to maximize potential capital gains. The increase in demand would in turn raise stock prices. Increases in expected inflation may also signal a potential increase in real activity, production and hence higher stock returns (Fama & Gibbons, 1982).

6. Granger causality test

In addition to the study of the impact of conventional stock market and macroeconomic variables, we perform Granger causality tests (Granger, 1969) to investigate the effects of unidirectional causality from these variables to Islamic stock market return across regimes. For this reason, we use a bivariate Markov switching vector autoregressive model (MS-VAR) developed by Krolzig (1997).

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Table 4b
Bivariate MS-VAR results for emerging markets.

	$RISLM_t$	$RISLM_t$	$RISLM_t$	$RISLM_t$	$RISLM_t$	$RISLM_t$
Regime 1						
μ_1	0.040** (0.00)	0.023** (0.00)	0.022** (0.00)	0.005* (0.09)	0.013** (0.00)	0.019** (0.00)
$RISLM_{t-1}$	−0.161** (0.03)	0.043 (0.22)	−0.083** (0.03)	−0.149** (0.00)	−0.070** (0.02)	−0.063* (0.06)
$RISLM_{t-2}$	–	0.038 (0.24)	–			0.030 (0.35)
$RCONV_{t-1}$	0.306** (0.00)	–	–			
$RCONV_{t-2}$	–	–	–			
$CVAR_{t-1}$	–	1.814 (0.25)	–			
$CVAR_{t-2}$	–	−3.095** (0.05)	–			
INF_{t-1}	–	–	−2.029** (0.00)			
INF_{t-2}	–	–	–			
ΔINT_{t-1}	–	–	–	−0.068 (0.30)		
ΔINT_{t-1}	–	–	–			
ΔSYC_{t-1}	–	–	–		−1.564* (0.09)	
ΔSYC_{t-2}	–	–	–			
ΔMS_{t-1}	–	–	–			0.028 (0.99)
ΔMS_{t-2}	–	–	–			−0.45** (0.00)
σ_1^2	0.003** (0.00)	0.004** (0.00)	0.004** (0.00)	0.005** (0.00)	0.004** (0.00)	0.004** (0.00)
Regime 2						
μ_2	−0.100** (0.00)	−0.018 (0.25)	−0.006 (0.31)	−0.025** (0.00)	−0.005 (0.45)	−0.019 (0.90)
$RISLM_{t-1}$	−0.241 (0.18)	−0.009 (0.86)	0.252** (0.00)	−0.212** (0.00)	0.309** (0.00)	0.132** (0.03)
$RISLM_{t-2}$	–	−0.190* (0.08)	–			−0.079 (0.17)
$RCONV_{t-1}$	0.724** (0.00)	–	–			
$RCONV_{t-2}$	–	–	–			
$CVAR_{t-1}$	–	−3.148** (0.05)	–			
$CVAR_{t-2}$	–	1.589 (0.21)	–			
INF_{t-1}	–	–				
INF_{t-2}	–	–	0.374 (0.48)			
ΔINT_{t-1}	–	–	–	0.058 (0.15)		
ΔINT_{t-1}	–	–	–			
ΔSYC_{t-1}	–	–	–		0.438 (0.38)	
ΔSYC_{t-2}	–	–	–			
ΔMS_{t-1}	–	–	–			−0.031 (0.98)
ΔMS_{t-2}	–	–	–			0.285 (1.00)
σ_2^2	0.004** (0.00)	0.019** (0.00)	0.019** (0.00)	0.008** (0.00)	0.013** (0.00)	0.013** (0.00)
p_{11}	0.71** (0.00)	0.93** (0.00)	0.94** (0.00)	0.37** (0.00)	0.96** (0.00)	0.95** (0.00)
p_{22}	0.85** (0.00)	0.67** (0.00)	0.89** (0.00)	0.73** (0.00)	0.86** (0.00)	0.86** (0.00)
$E(d_1)$	3.48	13.96	15.78	1.60	24.64	21.40
$E(d_2)$	6.67	3.06	9.09	3.71	6.93	6.92
$Log(L)$	3675.08	8896.67	7250.61	31166.24	7870.65	6440.32

Note: $E(d_1)$ and $E(d_2)$ are respectively the expected duration of the low-volatility regime and the high volatility one. P-values are in parentheses. ** and * indicate significance at the 5% and 10% levels, respectively.

The MS-VAR model can be written as follows:

$$\begin{aligned} Y_t &= \mu_{1S_t} + \sum_{j=1}^{p} a_{jS_t} y_{t-j} + \sum_{j=1}^{p} b_{jS_t} x_{t-j} + \varepsilon_{1t} \\ x_t &= \mu_{2S_t} + \sum_{j=1}^{p} c_{jS_t} y_{t-j} + \sum_{j=1}^{p} d_{jS_t} x_{t-j} + \varepsilon_{2t} \end{aligned} \tag{13}$$

where y_t represents the monthly stock market index returns, x_t is a conventional stock market or macroeconomic variable and ε_{it}, $i = 1, 2$ is a white noise process with mean zero and a variance depending on regime S_t.P is the order of the respective lag variable. Granger (1969) defines causality between the two variables, x and y, in terms of predictability. Thus, we can determine that x_t Granger causes y_t in a specific regime if at least one of the $b_i, j = 1 \ldots p$ is significantly different from zero in that regime and y_t Granger causes x_t in a given regime if at least one of the $c_i, j = 1 \ldots p$ is significantly different from zero in that regime. The optimal lag length is determined using the Akaike information criteria (AIC). Empirical results of causality tests for developed and emerging countries are reported in Table 4.[3] The optimal lag length is determined using the Akaike information criteria (AIC).

Table 4 shows that except for the conventional stock market return ($RCONV$) which has causation effects on developed and emerging Islamic stock market return in the two regimes; other variables have limited effects. Our results confirm those of Ajmi et al. (2014) and suggest the rejection of the hypothesis of decoupling of the Islamic market from

[3] Only the results of the unidirectional causation from conventional stock market and economic variables to Islamic stock market return has been reported; others results are available on request.

their conventional counterparts. These imply that the Islamic financial system may not provide a large diversification benefits for portfolios managers. Furthermore, none of economic variables Granger cause Islamic index returns in the high volatility regime. In the low volatility regime, we find that inflation rate (INF), the changes in the slope of the yield curve (ΔSYC) and the first difference of the money supply (ΔMS) Granger cause emerging Islamic return but only the inflation rate Granger cause developed Islamic index return. Finally, the conditional variance of the conventional index return ($CVAR$) Granger cause emerging Islamic index return in the two regimes; which isn't the case for developed countries.

7. Conclusion and policy implications

In recent years, the world Islamic financial market has attracted international capital flows from both Muslim and non-Muslim investors. Islamic finance is not only a fast-growing field but has now officially moved into mainstream financial markets. With the continual growth of transactions volume in Islamic equities, an on-going discussion has emerged on whether the Islamic and conventional securities are affected differently by economic variables.

In this paper, we study the impact of conventional index return and volatility, the inflation rate, the short-term interest rate, the slope of the yield curve and the change in the money supply on Islamic stock market return across financial regimes over the period from June 2002 to June 2014 for ten developed and ten emerging markets. Results show that conventional index return and the changes in money supply have significant impact on Islamic index return in low and high volatility regimes for developed and emerging markets. For other economic variables, estimated coefficients are significant especially in the low volatility regime. Besides, the difference in the impact of studied variables on Islamic index return for developed and emerging market is limited. Finally, similar conclusions are obtained by using Granger causality analysis across regimes especially for the great impact of conventional index return.

The results of this paper are relevant and valuable for both investors and financial analysts and give new insights about the dependence relationships between Islamic stock market return and macroeconomic variables across market regimes. We can define whether there exists opportunities for profit from the inefficiencies of stock market mechanisms in the transfer of information between stock markets. The knowledge of financial co-movement among Islamic and conventional stock markets is important for portfolio diversification and risk management. More precisely, the most important finding of our paper is that the co-movement between Islamic stock market return and macroeconomic variables vary across regimes. Finally the significant coefficients in both Markov switching regressions and MS-VAR models may be useful to policy makers to help predict to what degree macroeconomic policy can be a good way to stabilize Islamic financial markets across market regimes.

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