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RATIONAL SPECULATIVE BUBBLES IN THE OIC (ORGANISATION OF ISLAMIC CONFERENCE) STOCK MARKETS

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ABSTRACT

While OIC countries play a vital economic role in oil exports, tourism, and international investments, the stock market behavior in OIC countries has been largely unrevealed and bypassed by financial market research. To fill this gap, we examine the possibility of rational speculative bubbles in OIC stock markets by employing alternative bubble tests, such as fractional bubble tests and duration dependence tests. We confirm that log dividend yields of OIC stock markets are fractionally integrated and none of the nonparametric Nelson-Aalen smoothed hazard functions is monotonically decreasing. Therefore, we do not find strong evidence of rational speculative bubbles in OIC stock markets without regard to currency denominations.

JEL Classification: C14, C41, G15

Key words: Rational speculative bubbles, OIC stock markets, Fractional integration tests, Duration dependence tests

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1. INTRODUCTION

The OIC (Organisation of the Islamic Conference) is an inter-governmental organization grouping of 57 mostly Islamic nations in the Middle East, North and West Africa, Central Asia, Southeast Asia, the Indian subcontinent and South America. These OIC member countries decided to promote Islamic solidarity by coordinating social, economic, scientific, and cultural activities.[1] However, despite their important economic role to international community through oil exports, tourism, and international investments, the stock market behavior in OIC countries has been largely unrevealed and bypassed by financial market research.

Over the decades, there has been considerable academic interest in rational speculative bubbles in which price deviates from fundamental price for the mature and developing stock markets (e.g., Brooks and Katsaris 2003; Gürkaynak, 2005, among others). From the perspective of the US investors, it is helpful to investigate rational speculative bubbles in OIC countries for the purpose of international diversification. It is well-established that international portfolio in developing countries can Pareto-improve the risk-return tradeoff and the unconditional mean-variance frontier shifts significantly upwards (Divecha, Drach, and Stefak, 1992).

In addition, even though the volatility of the individual emerging stock market is high, the pair-wise correlations and cointegration are relatively low among OIC stock markets. Therefore, significant benefits are possible in diversifying into OIC stock markets because the low correlations should reduce portfolio volatility and the relatively low or no cointegration among these stock markets also lead to further opportunities for long-term portfolio risk diversification (Hassan, 2003).

In a similar vein, from the perspective of OIC investors, it is also imperative to investigate the possibility of rational speculative bubbles to attract significant amount of international investments into OIC countries. As Claessens (1995) points out, foreign equity investment can benefit developing countries by diversifying the sources of external finance, increasing the risk-bearing by investors, reducing the cost of capital, improving incentives for managing the investment process, assisting in the development of domestic capital markets, and enhancing the mobilization of domestic resources.

However, despite the potential benefits of portfolio diversification in OIC countries, there is a lack of research and relatively much less is known about the existence of rational speculative bubbles of OIC stock markets. Access to most of OIC stock markets has been limited to international investors who are seeking portfolio diversification. For example, among OIC stock markets in the MENA (Middle East and North Africa) region, only Morocco and Egypt offer unrestricted access to foreign investors while Jordan allows foreigners to hold up to 50% of a company's capital. Although the market capitalization of Saudi Arabia's stock market is the 12th-largest among emerging markets in the world ($82 billion at the June of 2002), Saudi Arabia's stock market is still closed to direct investment by non-GCC (Gulf Cooperation Council: Bahrain, Oman, Saudi Arabia, Kuwait, UAE, and Qatar) nationals (Yu and Hassan, 2008).

Furthermore, OIC countries in South Asia (Bangladesh and Pakistan) and Sub-Saharan (Nigeria and Côte d'Ivoire) regions are classified as low income with one major agribusiness, while their stock markets are less developed, relatively small, and illiquid. As a result, information on OIC stock markets to international investors is generally less available than in other mature and emerging financial markets with the exception of OIC stock markets in the East Asian & Pacific (Malaysia and Indonesia) region, that have attracted a significant amount of both foreign direct investments and portfolio investments. Therefore, for the purpose of international portfolio diversification and stock market development, it is essential to identify whether the recent OIC stock market movements are based on correct pricing or whether they are driven by rational speculative bubbles.

In this paper, to be consistent with recent literature and to overcome the shortcomings of traditional bubble tests, such as unit root tests and cointegration tests, we employ fractional integration tests based on ARFIMA model (e.g., Cuñado, Gil-Alana, and Gracia, 2005; Koustas and Serletis, 2005, among others) and duration dependence tests based on survival analysis (e.g., McQueen and Thorley, 1994; Cochran and Defina, 1996; Chan, McQueen, and Thorley, 1998; Cameron and Hall, 2003; Tudela, 2004; Buehler, Kaiser, and Jaeger, 2006, among others) to examine the possibility of rational speculative bubbles in OIC stock markets. Traditional bubble tests have mainly relied on integer orders of integration to the log dividend yield or have tested for integer cointegration between stock dividends and prices by examining expectations of future streams of dividends. However, Koustas and Serletis (2005) indicate that the notion of fractional integration allows more flexible modeling of the low frequency dynamics of stock prices, dividends, and their equilibrium relationship, while allowing significant deviations from equilibrium in the short-run. Therefore, we check the log dividend yields for a fractional exponent in the differencing process based on ARFIMA models since the unit root and cointegration tests allow for only integer orders of integration.

For duration dependence tests, researchers have traditionally preferred fitting parametric hazard functions such as log-logistic, exponential, Weibull, and Gompertz specifications along with semiparametric tests based on Cox regression model. However, we find that nonparametric hazard functions have much better small-sample properties and are more intuitive to interpret whether hazard functions is decreasing or increasing. Therefore, we estimate nonparametric Nelson-Aalen smoothed hazard functions together with traditional parametric and semi-parametric hazard specifications to investigate duration dependence in runs of positive excess returns of OIC stock markets because we have relatively small sizes of samples to fit parametric hazard functions. We firmly believe that our approach to plot nonparametric smoothed hazard functions is more reliable to obtain the robust empirical results of duration dependence tests to identify the existence of bubbles in OIC stock markets.

We summarize the main results as follows. Although there have been frequent extreme fluctuations of OIC stock markets, we do not find strong evidence of rational speculative bubbles in the perspective of both OIC and US investors. Fractional integration tests built on ARFIMA models do not support the possibility of bubbles in the OIC stock markets. Similarly, duration dependence tests derived from nonparametric Nelson-Aalen hazard functions strongly reject the existence of bubbles. In addition, it appears that both OIC and US investors do not observe rational speculative bubbles in OIC countries, evidenced by the statistically identical hazard functions across them.

The organization of the paper is as follows. In Section 2, we provide sample selection criteria and summary statistics on OIC stock markets data. Section 3 points out the empirical shortcomings of traditional bubble tests, such as unit root tests and cointegration tests, to detect rational speculative bubbles. In Section 4, we describe fractional integration tests and duration dependence tests for rational speculative bubble identification. Section 5 analyzes the empirical results. Section 6 concludes.

2. OIC STOCK MARKETS DATA AND SUMMARY STATISTICS

2.1 DATA AND SAMPLE SELECTION CRITERIA

At present, although there are 57 OIC member countries across different geographic regions, we focus on the following 14 representative OIC stock markets because of data availability and their vital economic role to international community through oil exports, tourism, and financial markets. For illustration purposes, we provide the topography of the 57 OIC member countries in Appendix A. To investigate whether rational speculative bubbles exist or not in OIC stock markets, we collect monthly S&P/IFCG price indexes[2] of 14 OIC stock markets of Indonesia (1990:01–2003:03), Malaysia (1985:01–2003:03), Turkey (1987:01–2003:03), Bahrain (1999:01–2003:03), Egypt (1996:01–2003:03), Jordan (1979:01–2003:03), Morocco (1996:01–2003:03), Oman (1999:01–2003:03), Saudi Arabia (1998:01–2003:03), Tunisia (1996:01–2003:03), Bangladesh (1996:01–2003:03), Pakistan (1985:01–2003:03), Nigeria (1985:01–2003:03), and Côte d'Ivoire (1996:01–2003:03).

The sample periods for indexes are chosen on the basis that they represent the longest periods over which reliable data for OIC stock markets are available. All monthly S&P/IFCG price indexes are expressed in local currencies and US dollar denominations to consider both OIC and US investors' perspectives. Then, to perform duration dependence tests to examine the existence of rational speculative bubbles in OIC stock markets, we compute the monthly simple return (R_{t+1}) on OIC stock markets index held from time t to $t + 1$ like the following:

$$(1) \qquad R_{t+1} = \frac{I_{t+1} - I_t + D_{t+1}}{I_t} = \frac{I_{t+1} + D_{t+1}}{I_t} - 1$$

where I_t is the S&P/IFCG price indexes of OIC stock markets in local currency and US dollar denominations considered in time t, and D_{t+1} is the dividends in period $t + 1$.

In addition, reliable dividend yields data are essential to perform formal econometric bubble tests such as fractional integration tests. Therefore, we consider slightly different sample periods of dividend yields for OIC stock markets from those of S&P/IFCG price indexes: Indonesia (1992:01–2003:03), Malaysia (1985:11–2003:03), Turkey (1987:11–2003:03), Bahrain (2000:01–2003:03), Egypt (1996:12–2003:03), Jordan (1984:12–2003:03), Morocco (1996:12–2003:03), Oman

(2000:01–2003:03), Saudi Arabia (1998:11–2003:03), Tunisia (1996:12–2003:03), Bangladesh (1996:12–2003:03), Pakistan (1985:11–2003:03), Nigeria (1985:12–2003:03), and Côte d'Ivoire (1996:12–2003:03).

The source of data is the Emerging Markets Data Base (EMDB) published by Standard & Poor's, then we categorize each OIC country into different geographic regions and income groups based on World Bank classifications from World Bank's *World Development Indicators* 2006 (WDI) database. Out of 57 OIC full members, our data set consists of total 14 OIC countries including 2 (Indonesia and Malaysia) in East Asia & the Pacific, 1 (Turkey) in Europe & Central Asia, 7 (Bahrain, Egypt, Jordan, Morocco, Oman, Saudi Arabia, and Tunisia) in the Middle East & North Africa, 2 (Bangladesh and Pakistan) in South Asia, and 2 (Nigeria and Côte d'Ivoire) in Sub-Saharan Africa. Most of the OIC members belong to low or lower-middle income groups based on World Bank classifications, especially in South Asia and Sub-Saharan African regions. However, within the same OIC members, many of the oil-rich kingdoms are classified as high-income non-OECD or upper middle income groups as listed in Appendix B.

2.2 OIC STOCK MARKET CHARACTERISTICS AND SUMMARY STATISTICS

In Table 1, we provide an overview of OIC stock market characteristics. Local price indices changed a lot during our sample periods, especially for Turkey and Nigeria. In most cases, OIC countries experienced a huge decrease in market capitalization in millions of US dollars except Saudi Arabia, Pakistan, and Nigeria. However, there were only a few changes in the number of listed domestic companies and monthly value traded in millions of US dollars with some exceptions. Unlike other OIC stock markets, it appears that Saudi Arabia, Pakistan, and Nigeria experienced an increase in both sizes and liquidity of their stock markets, evidenced by enhanced local price indices, market capitalizations, and monthly value traded.

In Table 2, we also report summary statistics of monthly S&P/IFCG price index returns for 14 OIC stock markets in both local currency and US dollar denominations. During our sample periods, most of OIC stock markets experienced severe fluctuations. In local currency denominations (Panel A), Turkey and Bangladesh investors experienced extreme stock market movements, evidenced by 19.64% and 14.27% of standard deviations. Maximum (minimum) monthly returns for Turkey and Bangladesh are 79.33% (-39.27%) and 90.65% (-30.15%), respectively.

TABLE 1
OIC Stock Market Characteristics

	Local Price Indices		Market Capitalization		Number of Listed Domestic		Monthly Value Traded	
	Feb. 1999	Feb. 2003	Feb. 1999	Feb. 2003	Feb. 1999	Feb. 2003	Feb. 1999	Feb. 2003
Indonesia	103.59	91.55	9910.05	8209.30	58	58	294.57	355.78
Malaysia	146.69	176.25	42611.24	36284.93	147	105	981.01	896.91
Turkey	208082.2	638230.6	23296.10	15453.50	58	46	3881.24	3165.63
Bahrain	100.64	81.64	4126.04	1726.55	15	11	21.03	7.03
Egypt	132.60	63.88	6690.79	1971.66	66	52	458.71	124.89
Jordan	327.95	293.58	4486.17	3725.61	41	29	54.03	33.05
Morocco	280.13	177.02	9701.66	4262.15	18	19	150.15	49.07
Oman	102.41	88.83	2607.79	1794.27	34	27	16.52	18.13
Saudi	69.13	124.19	22451.89	37565.46	21	27	503.30	842.99
Tunisia	66.20	50.08	1340.39	961.43	13	19	26.19	6.09
Bangladesh	68.18	65.80	557.85	484.24	49	59	8.54	13.91
Pakistan	299.25	569.80	2228.36	2735.31	53	41	766.97	1651.01
Nigeria	5956.69	20106.29	1835.80	3364.29	28	31	4.95	42.21
Côte d’Ivoire	143	118.87	1133.18	293.34	12	12	5.05	0.65

Note: The table describes the characteristics of 14 OIC stock markets. Market summary data is collected from S&P’s Emerging Market Data Base (EMDB).

TABLE 2
Summary Statistics

Panel A: OIC Investors

	Indonesia	Malaysia	Turkey	Bahrain	Egypt	Jordan	Morocco	Oman	Saudi Arabia	Tunisia	Bangladesh	Pakistan	Nigeria	Côte d'Ivoire
Sample Periods	1990M1 ~2003M3	1985M1 ~2003M3	1987M1 ~2003M3	1999M1 ~2003M3	1996M1 ~2003M3	1979M1 ~2003M3	1996M1 ~2003M3	1999M1 ~2003M3	1998M1 ~2003M3	1996M1 ~2003M3	1996M1 ~2003M3	1985M1 ~2003M3	1985M1 ~2003M3	1996M1 ~2003M3
Obs	159	219	195	51	87	291	87	51	63	87	87	219	219	87
Mean	0.0053	0.0066	0.0616	-0.0036	-0.0025	0.0076	0.0071	0.0005	0.0052	-0.0068	0.0025	0.0128	0.0257	0.0026
Median	-0.0023	0.0039	0.0263	-0.0031	-0.0122	-0.0011	0.0030	-0.0237	0.0028	-0.0081	-0.0128	0.0062	0.0197	-0.0020
Maximum	0.4278	0.3500	0.7933	0.1085	0.2614	0.1819	0.1513	0.1889	0.1046	0.2446	0.9065	0.3499	0.3182	0.1496
Minimum	-0.3245	-0.3119	-0.3927	-0.1124	-0.1198	-0.1288	-0.1073	-0.1184	-0.1259	-0.1483	-0.3015	-0.3461	-0.2088	-0.1780
Std. Dev.	0.1081	0.0906	0.1964	0.0361	0.0748	0.0446	0.0480	0.0659	0.0447	0.0502	0.1427	0.0950	0.0539	0.0545
Skewness	0.2250	0.2420	1.0156	0.1406	1.0258	0.9979	0.5498	1.0975	-0.2123	1.3036	3.2108	0.5289	1.3892	0.2983
Kurtosis	4.4989	5.3348	4.6300	5.3572	4.2276	4.9725	3.2317	4.1264	3.3981	10.0597	20.8708	5.9347	10.5392	4.2112
Jarque-Bera	16.23	51.88	55.11	11.97	20.72	95.48	4.58	12.93	0.89	205.31	1307.18	88.80	589.10	6.61
	[0.000]	[0.000]	[0.000]	[0.002]	[0.000]	[0.000]	[0.101]	[0.002]	[0.641]	[0.000]	[0.000]	[0.000]	[0.000]	[0.037]
Q(12)	15.35	27.24	11.26	14.06	11.15	6.92	28.09	15.65	8.26	23.21	22.17	14.61	35.66	7.62
	[0.223]	[0.007]	[0.506]	[0.296]	[0.516]	[0.862]	[0.005]	[0.207]	[0.764]	[0.02]	[0.036]	[0.263]	[0.000]	[0.814]

TABLE 2 (continued)
Summary Statistics

Panel B: US Investors

	Indonesia	Malaysia	Turkey	Bahrain	Egypt	Jordan	Morocco	Oman	Saudi Arabia	Tunisia	Bangladesh	Pakistan	Nigeria	Côte d'Ivoire
Sample Periods	1990M1 ~2003M3	1985M1 ~2003M3	1987M1 ~2003M3	1999M1 ~2003M3	1996M1 ~2003M3	1979M1 ~2003M3	1996M1 ~2003M3	1999M1 ~2003M3	1998M1 ~2003M3	1996M1 ~2003M3	1996M1 ~2003M3	1985M1 ~2003M3	1985M1 ~2003M3	1996M1 ~2003M3
Mean	-0.0007	0.0053	0.0226	-0.0036	-0.0086	0.0046	0.0054	0.0005	0.0052	-0.0099	-0.0017	0.0069	0.0098	0.0008
Median	-0.0093	0.0054	-0.0087	-0.0034	-0.0183	-0.0014	0.0040	-0.0237	0.0027	-0.0109	-0.0203	-0.0015	0.0125	-0.0033
Maximum	0.5477	0.5359	0.7129	0.1087	0.2588	0.1637	0.1582	0.1889	0.1052	0.2039	0.9066	0.3566	0.9919	0.2068
Minimum	-0.4012	-0.3135	-0.4066	-0.1121	-0.1262	-0.1288	-0.0980	-0.1184	-0.1260	-0.1739	-0.3087	-0.3520	-0.7074	-0.1969
Std. Dev.	0.1408	0.0997	0.1994	0.0360	0.0730	0.0455	0.0493	0.0659	0.0447	0.0549	0.1427	0.0974	0.1279	0.0629
Skewness	0.4289	0.6665	0.8370	0.1470	1.0477	0.6213	0.3037	1.0988	-0.2114	0.4538	3.2249	0.6986	1.1363	0.6079
Kurtosis	5.0627	7.9116	4.1189	5.3540	4.4816	4.4301	2.9725	4.1291	3.4010	6.1152	21.0740	6.2593	25.3719	4.7497
Jarque-Bera	33.06	236.35	32.94	11.96	23.87	43.52	1.34	12.97	0.89	38.17	1334.97	114.75	4614.22	16.46
	[0.0000]	[0.0000]	[0.0000]	[0.0025]	[0.0000]	[0.0000]	[0.5117]	[0.0015]	[0.6405]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0003]
$Q(12)$	24.32	36.68	11.44	14.00	14.22	7.96	21.60	15.62	8.25	27.42	21.95	9.68	14.54	11.61
	[0.0180]	[0.0000]	[0.4910]	[0.3010]	[0.2870]	[0.7880]	[0.042]	[0.2090]	[0.7650]	[0.0070]	[0.0380]	[0.6430]	[0.2670]	[0.4770]

Note: The table describes summary statistics of monthly S&P/IFCG price index returns for 14 OIC stock markets. The $Q(12)$ is the Ljung-Box portmanteau test statistic for 12 autocorrelations. The p-values are reported in square brackets for Jarque-Bera normality tests and Ljung-Box portmanteau tests

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Similarly, in US dollar denominations (Panel B), Turkey and Bangladesh still had the highest uncertainty in their monthly stock market movements out of OIC stock markets. Maximum (minimum) monthly returns for Turkey and Bangladesh are very close to those of local currency denominations, which amount to 71.29% (-40.66%) and 90.66% (-30.87%), respectively. In general, it appears that OIC and US investors experienced similar market fluctuations without regard to currency risks. However, the only exception is Nigeria where US investors suffered from extreme movements of stock markets due to exchange rate risks, as illustrated in Panel B of Figure 1 through the Box-Whisker plots.

The Jarque-Bera statistics show that, except Morocco and Saudi Arabia, all of monthly index returns for OIC stock markets are far from normally distributed with positive or negative skewness and leptokurtosis. Although some countries include a significant number of outliers as depicted in Figure 1, they might result from temporary or spurious shocks not directly related with bubbles, evidenced by insignificant values of the Ljung-Box portmanteau test statistics for 12 autocorrelations, $Q(12)$, except Malaysia, Morocco, Tunisia, and Bangladesh in Table 2.

We also find that the Ljung-Box portmanteau test statistic is only significant in the US dollar denomination for Indonesia. On the contrary, $Q(12)$ is only statistically significant in local currency denomination for Nigeria. Therefore, it appears that the maximum and minimum values for Nigeria in US dollar denominations might be considered as outliers or abrupt jumps, which commonly occurr in stock or foreign currency markets since extreme observations do not contribute to autocorrelations or persistence of equity returns, thus they should not be representatives of normal market movements of Nigeria. Therefore, we need to be equipped with formal econometric tests to detect the existence of rational speculative bubbles in OIC stock markets for various practical reasons such as international investment, portfolio diversification, risk management, and monetary policy purposes.

3. THE EMPIRICAL SHORTCOMINGS OF TRADITIONAL BUBBLE TESTS

According to theories of bubbles (e.g., Brooks and Katsaris, 2003; Cuñado, Gil-Alana, and Gracia, 2005; Kirman and Teyssière, 2005), the actual OIC stock market indexes deviate from the fundamental values if $Bubble_t > 0$. In this case, the markets indexes are considered to have rational speculative bubbles;

$$OIC_t^a = OIC_t^f + Bubble_t + \mu_t$$

$$(2) \qquad = \sum_{k=1}^{\infty} \frac{1}{(1+i)^k} E_t\left(Div_{t+k}\right) + \frac{E_t\left(Bubble_{t+1}\right)}{(1+i)} + \mu_t$$

where OIC_t^a is the actual index of 14 OIC stock markets considered in time t, Div_t is the dividend at period t, OIC_t^f is the fundamental value of the index in time t, i is the market discount rate, $E(.)$ is the mathematical expectation operator, and ì $_t$ is an identically and independently distributed (*i.i.d.*) stochastic process. $Bubble_t$ is the value of the bubble component in time t and it is entirely consistent with rational expectations and the time path of expected returns.

FIGURE 1
Box-and-Whisker Plots for OIC Countries

Panel A: OIC Investors

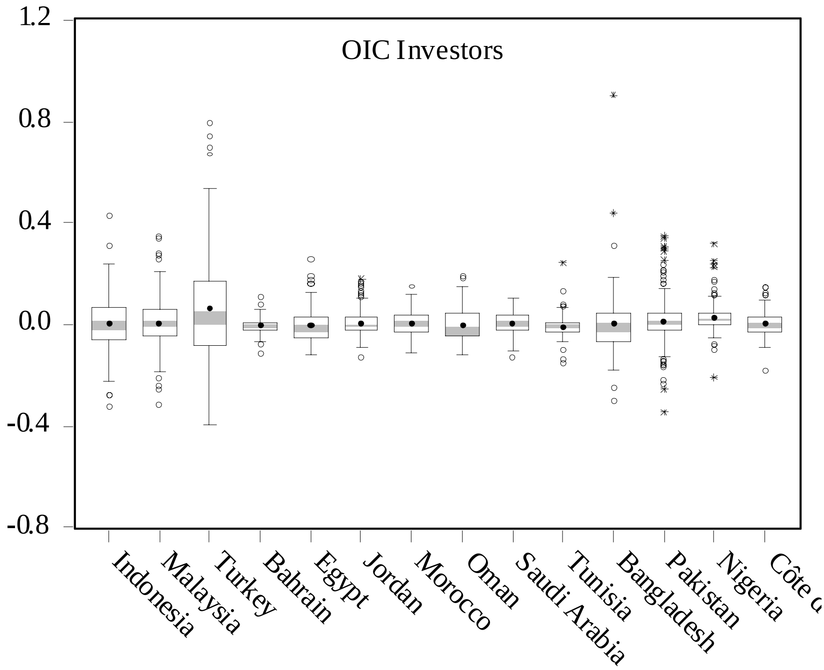
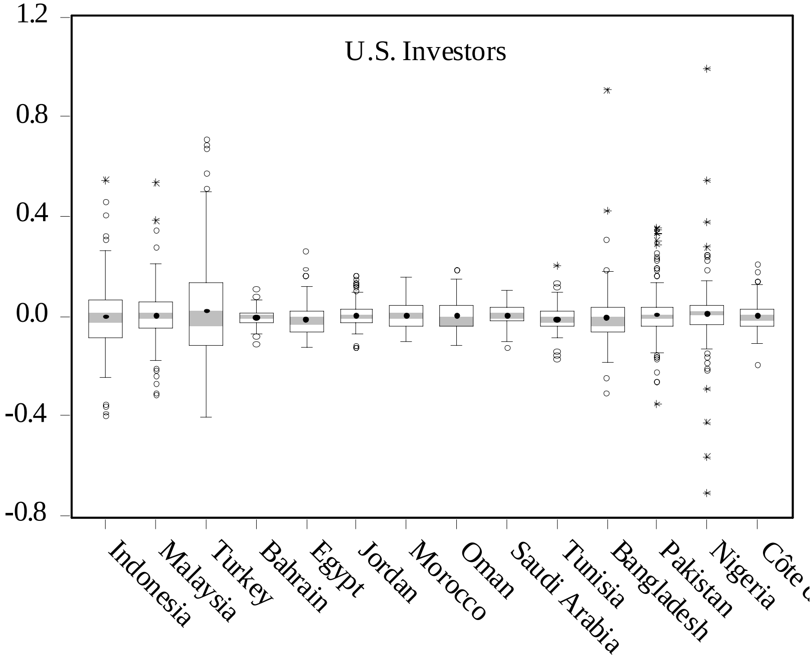


FIGURE 1 (continued)
Box-and-Whisker Plots for OIC Countries

Panel B: US Investors



Note: The Figure illustrates box-and-whisker plots, which explore maximum, minimum, median, upper/lower quartiles, and near/far outliers in the monthly S&P/IFCG price index returns of 14 OIC stock markets for both local currency and US dollar denominations.

However, previous studies on empirical tests to detect rational speculative bubbles concentrating on developed and emerging stock markets still remain inconclusive. Therefore, many financial economists have tried to explain the main sources of the controversy on the tests of bubbles detection. For example, Blanchard (1979) explains that speculative bubbles may take all kinds of shapes and their fundamentals may be stochastic. Kaizoji (2000) argues that bubbles and crashes come from the collective crowd behavior of many interacting agents. In theory, although an asset's fundamental value can be obtained by discounting the asset's future earnings stream, the difficulties in estimating the earnings stream and in proper discounting make the identification of bubbles empirically challenging (See Chen, 2001). Gürkaynak (2005) surveys the formal econometric tests of asset price bubbles and concludes that we cannot distinguish bubbles from time-varying or regime-switching fundamentals.

A traditional approach to test the null hypothesis of rational speculative bubbles is to investigate univariate time series properties of log dividend yields using unit-root tests such as the Augmented Dickey-Fuller (ADF) tests and the Phillips-Perron (PP) tests. The evidence of a unit root in the dividend yield is consistent with rational bubbles in OIC stock markets. In other words, nonstationary log dividend yields are consistent with the existence of rational bubbles, while stationarity implies that deviations from market fundamentals are short lived. Therefore, the log of dividend yields should be a stationary process if there are no rational speculative bubbles. Such a test directly on the dividend yield is actually a test for cointegration between index and dividends.

However, many researchers have also reported that it is very difficult to detect rational speculative bubbles precisely using traditional econometric tests, such as unit root tests and cointegration tests, mainly relying on expectations of future steams of dividends especially in small samples. For example, Taylor and Peel (1998) point out that although rational speculative bubbles imply noncointegration of index or stock prices and dividends, the traditional cointegration tests are subject to size distortion or specification error especially in small samples. Due to these undesirable properties of cointegration tests, they apply the robust noncointegration test with much smaller size distortion and good power characteristics to a long-run data of US real stock price and dividend, then reject the bubbles hypothesis on US data. More recently, using S&P 500 log dividend yields, Koustas and Serletis (2005) show that Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit-root tests are unable to reject a unit root in the price-dividend ratios (dividend yields), which suggests the lack of a cointegrating relationship between stock prices and dividends.

4. ALTERNATIVE ECONOMETRIC TESTS TO IDENTIFY RATIONAL SPECULATIVE BUBBLES

In this section, to overcome the shortcoming of traditional bubble tests we briefly summarize the general idea of alternative econometric tests, such as fractional integration tests and duration dependence tests, to identify rational speculative bubbles in OIC stock markets.

4.1 FRACTIONAL INTEGRATION TESTS

With a fractional integration parameter *d*, the ARFIMA (*p*, *d*, *q*) specification is written as,

(3) $$\Phi(L)(1-L)^d(y_t-\mu_t)=\Theta(L)\varepsilon_t$$

and the fractional differencing operator, $(1\text{-}L)^d$, is defined by the following binomial expansion:

(4) $$\begin{aligned}(1-L)^d &= \sum_{j=0}^{\infty}\binom{d}{j}(-L)^j \\ &= 1-dL-\frac{d(1-d)}{2!}L^2-\frac{d(1-d)(2-d)}{3!}L^3-\ldots\end{aligned}$$

where y_t is log dividend yields, ì $_t$ is the mean of dividend yields and *L* is the lag operator, $L^k = y_{t\text{-}k}$. The ARMA part of the model:

$$\Phi(L)=1-\sum_{i=1}^{p}\phi_i L^i \text{ and}$$

represent invertible and stationary autoregressive and moving average polynomial in the lag operation *L*. The *p* and *q* are integers, but *d* is in real values, respectively. Therefore, the ARFIMA (1, *d*, 1) model is:

(5)

We also say that z_t = y_t– ì$_t$ is $I(d)$, integrated of order *d* and the zero mean z_t is covariance stationary if $d < 0.5$ with autocovariance function (ACF) that decays hyperbolically. When , the ARFIMA process has a long memory. The process is called intermediate memory or overdifferenced for . We check the log dividend yield for a fractional exponent in the differencing process via the Exact Maximum Likelihood (EML) using BFGS algorithm with numerical derivatives. The log-likelihood function and more detailed computational procedures are described in Doornik and Hendry (2001).

If rational speculative bubbles are present in OIC stock markets, the fractional integrating parameter of log dividend yields, *d*, should have a unit root nonstationarity ( ) even in the fractional integration tests with more statistical power and flexibility than traditional unit root tests. However, if the fractional integrating parameter of log dividend yields, *d*, is statistically zero or is fractionally integrated with $d > -1$, then we can reject the possibility of rational speculative bubbles in OIC stock markets.

4.2 DURATION DEPENDENCE TESTS

To identify rational speculative bubbles in the perspective of OIC and US investors, we estimate the following semiparametric Cox proportional hazards,

(6) $$h_i\left(t \mid Currency\right) = h_0\left(t\right)\exp\left\{\beta \cdot Currency\right\}$$

where $h_0(t)$ is the base-line hazard at time *t* and *â* an unknown parameter to estimate. The variable *Currency* is a dummy variable with values 1 or 0; it equals 1 in the case of local currency denominations and 0 otherwise. By definition, exp{*âCurrency*} is called the hazard ratio. We consider a variety of parametric hazard function specifications, such as exponential, Weibull, and Gompertz regressions, depending on the expected shapes of hazard functions by making reasonable assumptions about the shape of the baseline hazard, $h_0(t)$. Then, we estimate parameter values and hazard ratios in each semiparametric and parametric hazard specification to decide whether the bubble identifications between OIC and US investors are affected by different currency denominations.

We also perform bubble tests by plotting an estimate of the nonparametric Nelson-Aalen smoothed hazard function, *h*(*t*), calculated as a Gaussian kernel smooth of the estimated hazard contributions to consider the relative small sample sizes of OIC stock markets (Cleves, Gould, and Gutierrez, 2004). Therefore, we can estimate the hazard function as follows.

(7) $$\hat{h}\left(t\right) = b^{-1}\sum_{j=1}^{D} K\left(\frac{t - t_j}{b}\right)\Delta\hat{H}\left(t_j\right)$$

for some symmetric Gaussian kernel density function and bandwidth *b*. The estimated hazard contribution, $\Delta\hat{H}\left(t_j\right)$, is defined as:

(8) $$\Delta\hat{H}\left(t_j\right) = \hat{H}\left(t_j\right) - \hat{H}\left(t_{j-1}\right)$$

and

(9) $$\hat{H}\left(t\right) = \sum_{j \mid t_j \le t} \frac{d_j}{n_j}$$

where n_j is the number at risk at time t_j, d_j is the number of failures at time t_j, and the sum is over all distinct failure times less than or equal to t. For each observed run ends time , $\hat{H}(t_j)$ is the estimated cumulative Nelson-Aalen estimator.

In any case, the null hypothesis (H_0) of no duration dependence implies that the probability of a run ending is independent of the prior returns or that positive and negative abnormal returns are random. The alternative hypothesis (H_1) of duration dependence suggests that the probability of a positive run ending should have a decreasing function of the run length. Therefore, if bubbles are detected, the hazard rates should decrease as the run length increases.

5. EMPIRICAL RESULTS FOR BUBBLE TESTS

5.1 AUTOCORRELATION TESTS OF MONTHLY S&P/IFCG PRICE INDEX RETURNS

Since rational speculative bubbles must be persistent to survive several months or years until market crashes, we should observe statistically significant positive autocorrelations, skewness, and leptokurtosis due to excess returns during bubble periods if bubbles exist in OIC stock markets. In our study, although summary statistics show excess kurtosis and positive or negative skewness evidenced by significant Jarque-Bera statistics except Morocco and Saudi Arabia, many other factors not directly related with bubbles can affect market returns. Therefore, much care should be taken in associating higher moments of market returns with the possibility of rational speculative bubbles. However, unlike Jarque-Bera normality tests, most OIC stock markets do not show significant positive autocorrelations except Malaysia, Morocco, Tunisia, and Bangladesh based on the Ljung-Box Portmanteau test statistics for 12 autocorrelations, $Q(12)$, supporting no bubbles to some degree in OIC stock markets.

The results of these autocorrelation tests question whether OIC stock markets really experienced rational speculative bubbles during our sample periods even though they have suffered a lot of extreme positive or negative monthly returns. As Koustas and Serletis (2005) insightfully point out, rational speculative bubbles must be continually expanding and persistent in order to survive since stock buyers will pay a price higher than that suggested by the fundamentals if they believe that someone else will subsequently pay an even higher price. Therefore, statistically significant positive autocorrelations among monthly returns are prerequisites for rational speculative bubbles to be present in OIC stock markets.

5.2 UNIT ROOT TESTS AND FRACTIONAL INTEGRATION TESTS OF LOG DIVIDEND YIELDS

For comparison purpose, we report the results of unit root tests of log dividend yields based on the Augmented Dickey-Fuller (ADF) tests and the Phillips-Perron (PP) tests in Table 3. We also provide the optimal number of bandwidths of unit root tests based on the Newey-West automatic bandwidth selection methods. According to the results of the ADF tests as well as the PP tests, the levels of log dividend yields are stationary and do not have a unit root for only 5 (Indonesia, Turkey, Bahrain, Morocco, and Bangladesh) OIC countries.[3] Therefore, our unit root tests, in general, do not reject the null hypothesis (H_0) of bubbles, although the 1st differences of log dividend yields show strong stationarity, suggesting the existence of bubbles in many of OIC stock markets.

However, the estimation results of fractional integration tests via exact maximum likelihood (EML) methods in Panel A of Table 4 for log dividend yields of OIC stock markets strongly support the results of the Ljung-Box autocorrelation test statistics shown in Table 2. Furthermore, all of the values of the fractional integrating parameter, $\hat{d}$, is statistically significant with 0< d <0.5 except for Turkey and Bangladesh, implying that the ARFIMA (1, d, 1) process is covariance stationary and has a long memory. Therefore, the estimation results of ARFIMA models for fractional integration tests strongly deny the possibility of rational speculative bubbles in OIC stock markets.We also test whether fractional integrating parameter, $\hat{d}$, is statistically 0 (no unit root stationarity) or 1 (unit root nonstationarity) by performing linear restriction tests in Panel B of Table 4. Unlike previous studies (e.g., Brooks and Katsaris, 2003; Gürkaynak, 2005, among others) that have tested for integer orders of integration, we are able to obtain robust rejections of both ‘unit root’ and ‘no unit root’ hypotheses in the log dividend yields from the linear restriction tests. Therefore, we confirm that log dividend yields of OIC stock markets are fractionally integrated, thus parameter values, $\hat{d}$, are statistically different from zero or one. Therefore, we cannot conclude that OIC stock markets have experienced rational speculative bubbles based on fractional integration tests built on ARFIMA approaches.

TABLE 3
Unit Root Tests for Log Dividend Yields of OIC Stock Markets

Countries	The Augmented Dickey-Fuller (ADF) test				The Phillips-Perron (PP) test					
	Levels		1st difference		Levels			1st difference		
	t-statistic	Prob.	*t*-statistic	Prob.	adj *t*-statistic	Prob.	Band	adj *t*-statistic	Prob.	Band
Indonesia	-6.7717	[0.0000]	-9.1478	[0.0000]	-6.3005	[0.0000]	6	-9.1812	[0.0000]	4
Malaysia	-3.1225	[0.1038]	-13.8335	[0.0000]	-3.1225	[0.1038]	0	-13.8373	[0.0000]	7
Turkey	-5.2214	[0.0001]	-15.3214	[0.0000]	-5.2214	[0.0001]	0	-17.3610	[0.0000]	15
Bahrain	-5.7341	[0.0002]	-9.5296	[0.0000]	-5.7794	[0.0001]	7	-25.1514	[0.0000]	32
Egypt	-2.7633	[0.2154]	-7.0361	[0.0000]	-3.0893	[0.1164]	2	-6.9453	[0.0000]	6
Jordan	-3.4518	[0.0474]	-17.2275	[0.0000]	-3.2650	[0.0749]	1	-17.3313	[0.0000]	5
Morocco	-3.9797	[0.0134]	-8.9628	[0.0000]	-3.9571	[0.0143]	3	-10.1880	[0.0000]	8
Oman	-2.2517	[0.4488]	-6.8712	[0.0000]	-2.2504	[0.4495]	2	-6.8714	[0.0000]	1
Saudi Arabia	-3.3680	[0.0670]	-8.4659	[0.0000]	-3.1579	[0.1043]	11	-9.3204	[0.0000]	9
Tunisia	-2.2377	[0.4619]	-8.4220	[0.0000]	-2.2486	[0.4561]	2	-8.6269	[0.0000]	7
Bangladesh	-4.2755	[0.0057]	-9.0099	[0.0000]	-4.2683	[0.0059]	5	-9.2259	[0.0000]	10
Pakistan	-0.9984	[0.9409]	-14.6998	[0.0000]	-0.9076	[0.9522]	4	-14.7678	[0.0000]	3
Nigeria	-2.2175	[0.4769]	-15.1455	[0.0000]	-2.4795	[0.3380]	6	-15.1324	[0.0000]	5
Côte d'Ivoire	-2.4733	[0.3402]	-9.6217	[0.0000]	-2.4272	[0.3629]	1	-9.6849	[0.0000]	6

Note: Intercept and trend included in test equations. The null hypothesis (H_0) of ADF and PP tests is that log dividend yields have a unit root. The optimal numbers of bandwidths (Band) of PP tests are chosen based on the values of Newey-West automatic bandwidth selection methods. MacKinnon one-sided *p*-values are reported in square brackets.

TABLE 4
Fractional Integration Bubble Tests in OIC Stock Markets

	Indonesia	Malaysia	Turkey	Bahrain	Egypt	Jordan	Morocco
Sample Periods	92M01 ~03M03	85M11 ~03M03	87M11 ~03M03	00M01 ~03M03	96M12 ~03M03	84M12 ~03M03	96M12 ~03M03
Panel A: Parameter estimates							
$\hat{d}$	0.3769	0.4691	-0.3111	0.4984	0.4713	0.4930	0.4743
	[0.003]	[0.000]	[0.346]	[0.000]	[0.000]	[0.000]	[0.000]
AR1 ($\hat{\phi}$)	0.9327	0.7786	0.9984	0.0415	0.6121	0.8174	0.6482
	[0.000]	[0.000]	[0.000]	[0.955]	[0.001]	[0.000]	[0.005]
MA1 ($\hat{\theta}$)	-0.1154	-0.2845	0.0402	-0.1727	-0.1022	-0.4524	-0.1798
	[0.332]	[0.007]	[0.907]	[0.813]	[0.611]	[0.035]	[0.511]
Panel B: Tests for linear restriction							
$\hat{d}=0$	9.3340	121.26	0.9137	48715.6	136.83	1811.04	148.73
	[0.000]	[0.000]	[0.339]	[0.000]	[0.000]	[0.000]	[0.000]
$\hat{d}=1$	25.4901	155.204	16.2262	49327.4	172.10	1914.95	182.68
	[0.002]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Panel C: Residual tests for model adequacy							
Normality	193.71	73.15	17.89	47.02	67.66	26.88	23.41
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
ARCH	1.1586	1.1632	2.2045	0.3009	0.0113	0.0235	3.5299
	[0.283]	[0.282]	[0.147]	[0.585]	[0.915]	[0.878]	[0.064]
Portmanteau	324.74	218.01	34.11	45.70	88.51	46.64	61.16
	[0.000]	[0.254]	[0.510]	[0.993]	[0.090]	[0.991]	[0.815]

Note: The table reports the estimation results of fractional integration tests via exact maximum likelihood (EML) methods. We also test if fractional integrating parameter ($\hat{d}$) is statistically 0 (no unit root) or 1 (unit root) by performing linear restriction tests. The parameters, $\hat{\phi}$ and $\hat{\theta}$, are the estimators of the first order autoregressive, AR(1), and moving average, MA(1), processes in ARFIMA (1, d, 1) models. To check model adequacy, we also tabulate residual tests for normality, ARCH effects, and serial correlations (Portmanteau tests). The p-values are reported in square brackets.

TABLE 4 (continued)
Fractional Integration Bubble Tests in OIC Stock Markets

	Oman	Saudi Arabia	Tunisia	Bangla-desh	Pakistan	Nigeria	Côte d'Ivoire
Sample Periods	00M01 ~03M03	98M11 ~03M03	96M12 ~03M03	96M12 ~03M03	85M11 ~03M03	85M12 ~03M03	96M12 ~03M03
Panel A: Parameter estimates							
$\hat{d}$	0.4924	0.4816	0.4757	-0.2238	0.4879	0.4521	0.4923
	[0.000]	[0.000]	[0.000]	[0.099]	[0.000]	[0.000]	[0.000]
AR1 ($\hat{\phi}$)	0.5257	0.8693	0.7839	0.9989	0.8423	0.8827	0.7889
	[0.362]	[0.001]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
MA1 ($\hat{\theta}$)	-0.3520	-0.5991	-0.3464	0.1092	-0.3508	-0.4756	-0.4481
	[0.573]	[0.060]	[0.037]	[0.514]	[0.008]	[0.000]	[0.032]
Panel B: Tests for linear restriction							
$\hat{d}=0$	1759.76	125.12	184.78	2.75	709.61	41.07	1588.83
	[0.000]	[0.000]	[0.000]	[0.096]	[0.000]	[0.000]	[0.000]
$\hat{d}=1$	1870.05	144.94	224.34	82.48	781.61	60.32	1688.76
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Panel C: Residual tests for model adequacy							
Normality	18.95	28.64	1042.10	63.43	53.32	79.73	29.10
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
ARCH	0.0599	0.0283	0.0155	1.0271	4.8201	1.1072	1.2916
	[0.808]	[0.867]	[0.901]	[0.313]	[0.029]	[0.294]	[0.259]
Portmanteau	42.95	19.63	125.95	115.84	164.27	170.92	59.52
	[0.715]	[0.983]	[1.000]	[0.825]	[0.981]	[0.693]	[0.853]

Note: The table reports the estimation results of fractional integration tests via exact maximum likelihood (EML) methods. We also test if fractional integrating parameter ($\hat{d}$) is statistically 0 (no unit root) or 1 (unit root) by performing linear restriction tests. The parameters, $\hat{\phi}$ and $\hat{\theta}$, are the estimators of the first order autoregressive, AR(1), and moving average, MA(1), processes in ARFIMA (1, *d*, 1) models. To check model adequacy, we also tabulate residual tests for normality, ARCH effects, and serial correlations (Portmanteau tests). The *p*-values are reported in square brackets.

In a recent study, Koustas and Serletis (2005) also report similar empirical findings as ours using S&P500 log dividend yields in that they also find the possibility of bubbles based on unit root tests, but they reject the null hypothesis of bubbles based on ARFIMA methods. They clarify that fractional integration tests are robust to the choice of parametric estimator of the fractional differencing parameter and data frequency, and bootstrap inference fully supports the estimation results. Therefore, our results of the fractional integration tests are inconsistent with rational speculative bubbles in OIC stock markets.

In Panel A of Table 4, the parameters, $\hat{\phi}$ and $\hat{\theta}$, are the estimators of the first order autoregressive, AR(1), and moving average, MA(1), processes in ARFIMA (1, d, 1) models, respectively. To check for model adequacy, we also tabulate residual tests for normality, ARCH effects, and serial correlations (Portmanteau tests) in Panel C of Table 4 along with the p-values in square brackets. For all of OIC stock markets, we can reject the null hypothesis of normality of residuals even after the ARFIMA fitting, which suggests using alternative fatter-tailed distributions such as skewed t-distribution and generalized error distribution (GED) rather than simply assuming normal distributions. For ARCH tests, it appears that ARFIMA settings are well-suited to fit residuals in all of OIC stock markets even though we assume constant volatility rather than time-varying GARCH-families for the convenience. For Portmanteau tests, it is likely that our ARFIMA (1, d, 1) model successfully captures the serial correlations of residuals except for Indonesia. However, although further complexity of model setups considering alternative fatter-tailed distributions and higher orders of AR or MA processes might improve the overall fit of our ARFIMA models, we do not believe that these additional computational efforts will change our main results of fractional cointegration tests to detect bubbles in OIC stock markets.

5.3 DURATION DEPENDENCE TESTS OF MONTHLY EXCESS POSITIVE RETURNS

In Table 5, to perform duration dependence tests, we compute the actual number of positive runs (sequences of excess returns of the same sign) for monthly positive excess index returns during our sample periods for the perspectives of both OIC and US investors. We find the longest positive runs (17 months) in Nigeria, next 10 months for Egypt and Bangladesh. All the OIC countries experience at least 3-month

TABLE 5

Run Lengths and Run Counts for Positive Runs of Monthly Excess Index Returns

Panel A: OIC Investors

	Indonesia		Malaysia		Turkey		Egypt		Jordan		Morocco		Tunisia		Bangladesh		Pakistan		Nigeria		Côte d'Ivoire	
Runs	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count
	1	38	1	54	1	29	1	8	1	39	1	11	1	22	1	16	1	50	1	35	1	24
	2	16	2	25	2	11	2	5	2	16	2	1	2	11	2	8	2	23	2	18	2	9
	3	10	3	15	3	8	3	1	3	3	3	4	3	5	3	3	3	13	3	13	3	2
	4	4	4	7	5	1	5	1	4	3	4	1	4	4	7	2	4	5	4	8		
	5	2	5	4			10	1	5	1	5	2			10	1	5	3	5	5		
	6	1	6	1					6	2							6	1	7	3		
																			13	1		
Total		71		106		49		16		64		19		42		30		95		83		35

Note: The tables report the results of duration dependence tests to detect the possibility of rational speculative bubbles in OIC stock markets. For this purpose, we compute the actual number of positive runs for monthly positive excess index returns for the perspectives of OIC (Panel A) and US (Panel B) investors. A run is defined as a sequence of excess returns of the same sign. We do not report the results for Bahrain, Oman, and Saudi Arabia because their hazard functions between OIC and US investors are identical as illustrated in Figure 2.

TABLE 5 (continued)
Run Lengths and Run Counts for Positive Runs of Monthly Excess Index Returns

Panel B: US Investors

Runs	Indonesia		Malaysia		Turkey		Egypt		Jordan		Morocco		Tunisia		Bangladesh		Pakistan		Nigeria		Côte d'Ivoire	
	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count	Length	Count
	1	37	1	52	1	27	1	10	1	38	1	12	1	22	1	17	1	51	1	45	1	25
	2	17	2	26	2	11	2	5	2	17	2	2	2	14	2	8	2	23	2	20	2	11
	3	10	3	18	3	7	3	1	3	8	3	2	3	5	3	3	3	11	3	13	3	2
	4	5	4	10	4	2	4	1	4	3	4	2	4	1	7	2	4	5	4	7	5	1
	5	3	5	3	5	1	10	1	5	2	5	1			10	1	5	3	6	5		
	6	2	6	1					6	1	7	1					6	1	7	3		
	7	1																	17	2		
Total		75		110		48		18		69		20		42		31		94		95		39

Note: The tables report the results of duration dependence tests to detect the possibility of rational speculative bubbles in OIC stock markets. For this purpose, we compute the actual number of positive runs for monthly positive excess index returns for the perspectives of OIC (Panel A) and US (Panel B) investors. A run is defined as a sequence of excess returns of the same sign. We do not report the results for Bahrain, Oman, and Saudi Arabia because their hazard functions between OIC and US investors are identical as illustrated in Figure 2.

positive runs for monthly positive excess index returns during our sample periods. However, it seems that those numbers of positive runs are too short, transient, and spurious to be considered as bubbles.

For example, before the worst market crash like Black Thursday (October 24th, 1929), the bull markets lasted about 63 months. The presence of a positive and increasing bubble premium continued about 18 months before the crash of Black Monday (October 19th, 1987). When the NASDAQ index of technology stocks in the US peaked, the market tripled in value between November 1998 and March 2000 (17 months). More recently, the Chinese stock market index had risen 700 percent, propelled by China's double digit growth rates and surging corporate profitability from July 1994 till June 2001 (84 months).

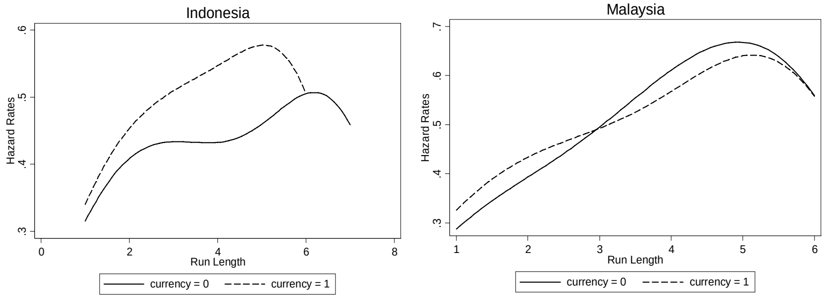
As we can also find from Figure 2, none of the nonparametric smoothed hazard functions is monotonically decreasing, implying no bubbles in OIC stock markets. Even though the nonparametric Nelson-Aalen smoothed hazard functions have different values of hazard rates except Bahrain, Oman, and Saudi Arabia, depending on the types of currency denominations, we observe that Indonesia, Malaysia, Jordan, Turkey, Morocco, Saudi Arabia, Tunisia, and Pakistan generally show increasing hazard functions, which are not acceptable if bubbles exist in those markets. Bahrain, Egypt, Oman, Bangladesh, Nigeria, and Côte d'Ivoire also show distinct patterns which are not completely consistent with bubbles. Nonparametric smoothed hazard functions initially increase then decrease. Therefore, although Nigeria has experienced 17-month positive runs for the perspective of US investors, this record does not appear to be directly related with bubbles because econometric tests do not support the possibility of rational speculative bubbles in Nigeria. Therefore, it is necessary to further examine economic fundamentals of Nigeria to figure out the boom of Nigeria stock market during our sample period.

In Table 6, we provide the comprehensive test results for equality of hazard functions between OIC and US investors. Panel A and B of Table 6 show the estimation results of hazard ratios, $\exp(\hat{a})$, of semiparametric Cox and parametric proportional hazard models, respectively, when the currencies are US dollar denominated. It seems that the hazard ratios are insensitive to model specifications, such as Cox, exponential, Weibull, and Gompertz models. In addition, all the hazard rates are not only very close to 1 but also statistically insignificant,

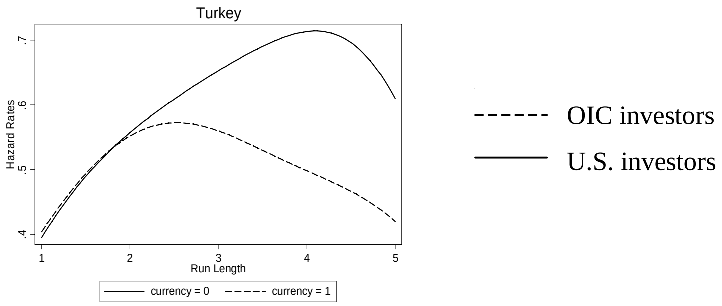
FIGURE 2

Nonparametric Nelson-Aalen Smoothed Hazard Functions

Panel A: East Asia & Pacific



Panel B: Latin America & Caribbean



Panel C: Middle East & North Africa

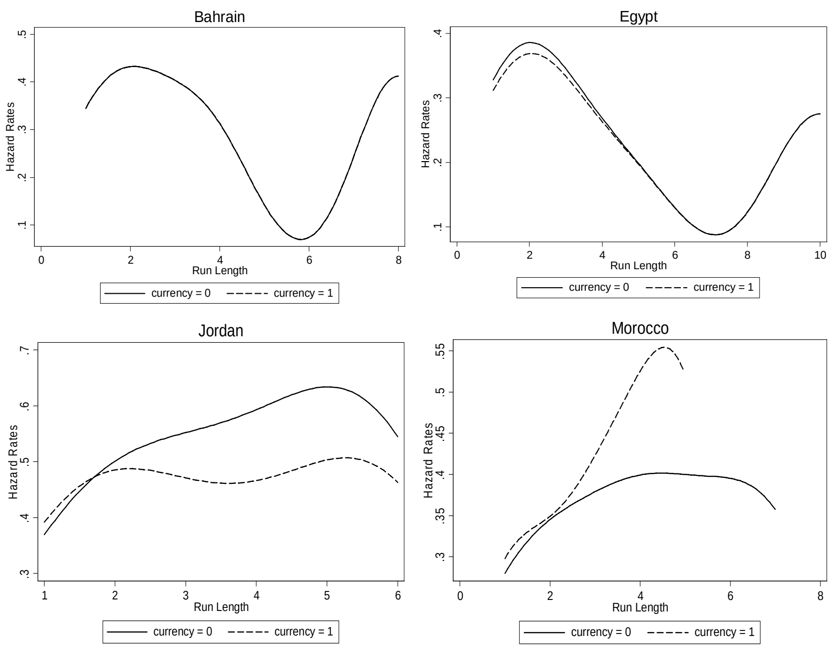
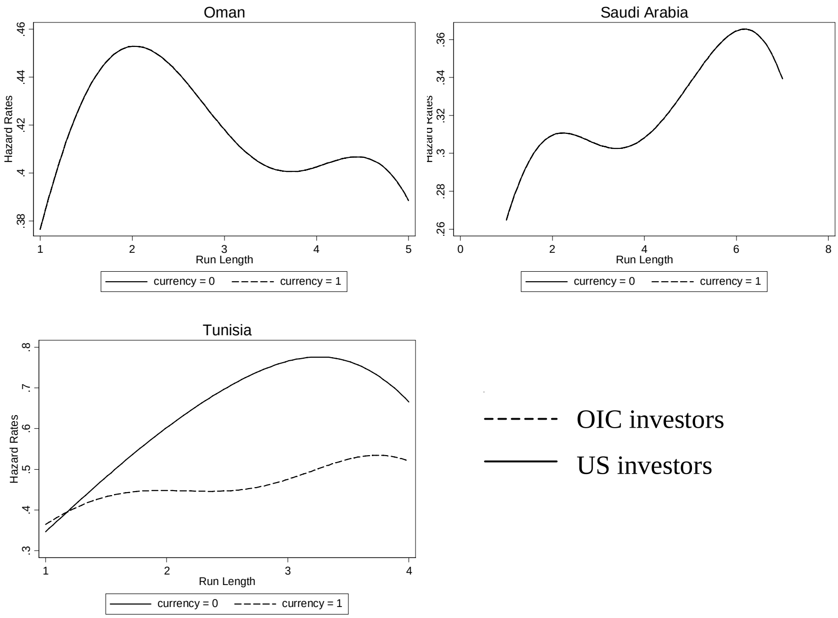


FIGURE 2 (continued)
Nonparametric Nelson-Aalen Smoothed Hazard Functions



Panel D: South Asia

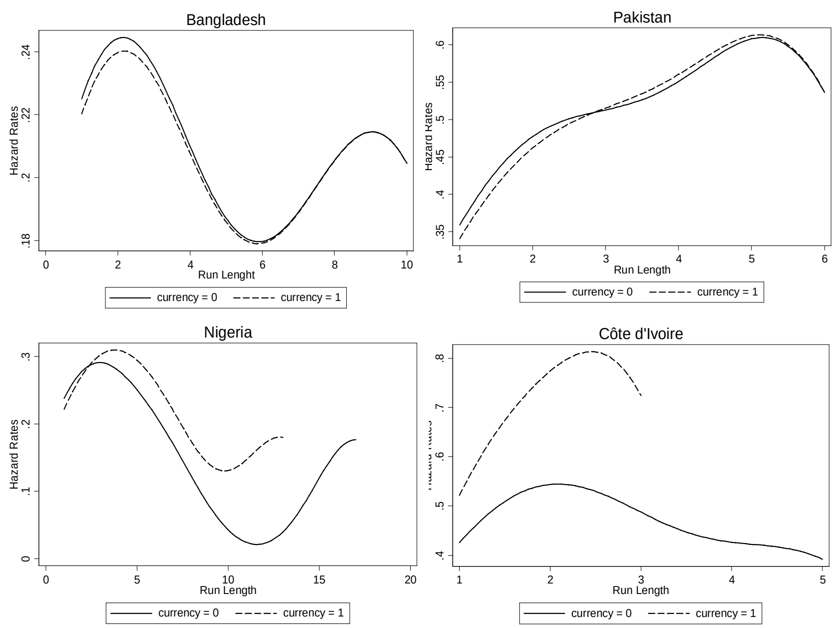


TABLE 6

Tests for Equality of Hazard Functions Between OIC and US Investors

	Indonesia	Malaysia	Turkey	Egypt	Jordan	Morocco	Tunisia	Bangladesh	Pakistan	Nigeria	Côte d'Ivoire
Panel A: Semiparametric tests based on Cox regression model											
Cox	1.0951	1.0751	1.0586	0.9544	1.0437	1.0298	0.9444	0.9710	0.9683	0.9587	1.0958
	[0.6970]	[0.7100]	[0.7800]	[0.8910]	[0.8060]	[0.9280]	[0.8530]	[0.9330]	[0.8720]	[0.8540]	[0.7510]
Panel B: Parametric tests for proportional hazards models											
Exponential	1.0848	1.0776	1.0591	0.9356	1.0551	1.0230	1.0000	0.9659	0.9700	1.0085	1.0971
	[0.7240]	[0.7000]	[0.2800]	[0.8450]	[0.7570]	[0.9430]	[1.0000]	[0.9210]	[0.8790]	[0.9700]	[0.7460]
Weibull	1.1702	1.1216	1.1404	0.9166	1.0626	1.0772	0.8767	0.9603	0.9516	1.0339	1.3632
	[0.4980]	[0.5550]	[0.5180]	[0.7980]	[0.7260]	[0.8170]	[0.6660]	[0.9080]	[0.8030]	[0.8830]	[0.2880]
Gompertz	1.1598	1.1012	1.1276	0.9356	1.0406	1.0928	0.8312	0.9666	0.9605	0.9740	1.3757
	[0.5240]	[0.6200]	[0.5550]	[0.8450]	[0.8190]	[0.7840]	[0.5480]	[0.9220]	[0.8400]	[0.9080]	[0.2860]

Note: This table provides hazard ratios in each of semiparametric (Panel A) and parametric (Panel B) hazard specifications to decide whether different currency denominations make shifts on the bubble identifications between OIC and US investors. We also provide test results for equality of hazard functions (Panel C) between OIC and US investors. We do not report the test results for Bahrain, Oman, and Saudi Arabia because their hazard functions between OIC and US investors are identical as illustrated in Figure 2.

TABLE 6 (continued)
Tests for Equality of Hazard Functions Between OIC and US Investors

	Indonesia	Malaysia	Turkey	Egypt	Jordan	Morocco	Tunisia	Bangladesh	Pakistan	Nigeria	Côte d'Ivoire
Panel C: Nonparametric tests for equality of hazard functions											
Log-rank	0.3000	0.2800	0.2000	0.0400	0.1400	0.0200	0.0700	0.0200	0.0600	0.0600	0.2900
	[0.5816]	[0.5994]	[0.6585]	[0.8394]	[0.7087]	[0.8976]	[0.7884]	[0.9018]	[0.8136]	[0.8084]	[0.5885]
Wilcoxon	0.1200	0.3400	0.1500	0.0400	0.5700	0.0100	0.2000	0.0200	0.0500	0.5000	0.1800
	[0.7340]	[0.5607]	[0.7004]	[0.8422]	[0.4518]	[0.9246]	[0.6522]	[0.8750]	[0.8256]	[0.4783]	[0.6710]
Tarone-Ware	0.1600	0.3600	0.1900	0.0400	0.4500	0.0000	0.0300	0.0200	0.0600	0.3800	0.1900
	[0.6921]	[0.5493]	[0.6613]	[0.8347]	[0.5009]	[0.9809]	[0.8644]	[0.8832]	[0.8060]	[0.5386]	[0.6591]
Peto-Peto	0.1000	0.4100	0.1300	0.0400	0.5200	0.300	0.2600	0.0300	0.0600	0.4900	0.2100
	[0.7472]	[0.5215]	[0.7229]	[0.8488]	[0.4701]	[0.8736]	[0.6077]	[0.8740]	[0.8058]	[0.4853]	[0.6496]

Note: This table provides hazard ratios in each of semiparametric (Panel A) and parametric (Panel B) hazard specifications to decide whether different currency denominations make shifts on the bubble identifications between OIC and US investors. We also provide test results for equality of hazard functions (Panel C) between OIC and US investors. We do not report the test results for Bahrain, Oman, and Saudi Arabia because their hazard functions between OIC and US investors are identical as illustrated in Figure 2.

implying that currency denominations do not make any difference for bubble identification for OIC stock markets. We also obtain nonparametric statistical test results supporting the equality of hazard functions between OIC and US investors from log-rank, Wilcoxon, Tarone-Ware, and Peto-Peto tests in Panel C of Table 6. Therefore, we do not find any statistically significant evidence of bubbles from both fractional integration tests and duration dependence tests for all the 14 OIC stock markets.

6. CONCLUSION

When we consider that OIC countries play an important role to international trade and financial markets, reliable empirical results of bubble tests in OIC stock markets can provide international investors as well as policy-makers with invaluable benchmark to better understand the irregular and highly fluctuating OIC stock markets behavior compared to other developed and emerging stock markets. For OIC and US investors, the formal analysis of OIC stock markets behavior including rational speculative bubbles will help them in their portfolio decisions and hedging purposes. Similarly, the empirical results of bubble tests in this paper will be also helpful to policy-makers in OIC countries to take actions to improve the functioning of these fast growing stock markets.

For this purpose, we extended the rational speculative bubble literature to OIC stock markets from the perspectives of OIC and US investors. This paper has employed fractional integration techniques and duration dependence tests based on the ARFIMA models and nonparametric Nelson-Aalen smoothed hazard functions in OIC stock markets. In this study, we do not find any strong evidence of rational speculative bubbles in OIC stock markets without regard to local currency and US dollar denominations. Fractional integration tests do not support the possibility of rational speculative bubbles, evidenced by fractionally integrated parameter values of log dividend yields. Similarly, duration dependence tests strongly reject the existence of bubbles as well, supported by nondecreasing nonparametric Nelson-Aalen smoothed hazard functions. These test results to identify rational speculative bubbles in OIC countries do not differ between OIC and US investors.

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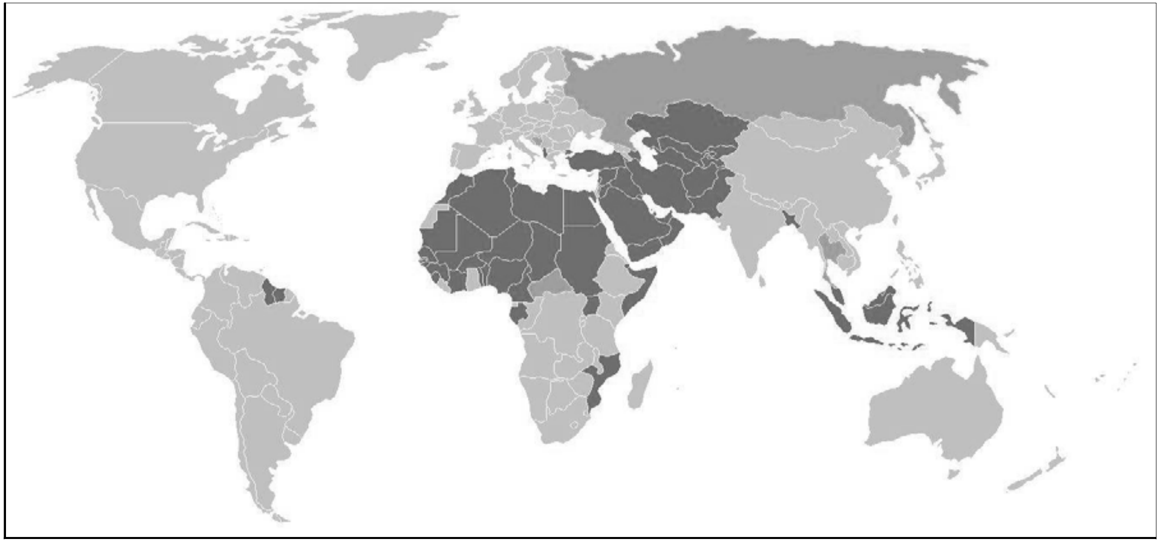
APPENDIX 1
OIC (Organisation of the Islamic Conference) Member Countries



Note: The OIC (Organisation of the Islamic Conference) is an inter-governmental organization grouping of 57 mostly Islamic nations in the Middle East, North and West Africa, Central Asia, Southeast Asia, the Indian subcontinent and South America. The 57 OIC members are illustrated in gray-black. The detailed list is tabulated in Appendix B.

Source: Wikipedia (http://en.wikipedia.org/wiki/Image:OIC_Member_Countries.JPG)

APPENDIX 2

The OIC Member Countries and World Bank Classification based on Geographic Regions and Income Groups

Geographic regions	Full members ($N = 57$)	Membership Year	Income groups	OIC countries ($N = 14$) analyzed in this paper
East Asia & Pacific	Indonesia	1969	Lower middle income	Indonesia
	Malaysia	1969	Upper middle income	Malaysia
Latin America & Caribbean	Suriname	1996	Lower middle income	
	Guyana	1998	Lower middle income	
Europe & Central Asia	Turkey	1969	Upper middle income	Turkey
	Azerbaijan	1991	Lower middle income	
	Albania	1992	Lower middle income	
	Kyrgyzstan	1992	Low income	
	Tajikistan	1992	Low income	
	Turkmenistan	1992	Lower middle income	
	Kazakhstan	1995	Lower middle income	
	Uzbekistan	1995	Low income	
Middle East & North Africa	Algeria	1969	Lower middle income	Bahrain
	Egypt	1969	Lower middle income	Egypt
	Iran	1969	Lower middle income	Jordan
	Jordan	1969	Lower middle income	Morocco
	Kuwait	1969	High income: non-OECD	Oman
	Lebanon	1969	Upper middle income	Saudi Arabia
	Libya	1969	Upper middle income	Tunisia
	Morocco	1969	Lower middle income	

APPENDIX 2 (continued)

The OIC Member Countries and World Bank Classification based on Geographic Regions and Income Groups

Geographic regions	Full members (N = 57)	Membership Year	Income groups	OIC countries (N = 14) analyzed in this paper
Middle East & North Africa	Yemen	1969	Low income	Bahrain
	Saudi Arabia	1969	High income: non-OECD	Egypt
	Tunisia	1969	Lower middle income	Jordan
	Bahrain	1970	High income: non-OECD	Morocco
	Oman	1970	Upper middle income	Oman
	Qatar	1970	High income: non-OECD	Saudi Arabia
	Syria	1970	Lower middle income	Tunisia
	United Arab Emirates	1970	High income: non-OECD	
	Iraq	1976	Lower middle income	
	Djibouti	1978	Lower middle income	
South Asia	Afghanistan	1969	Low income	Bangladesh
	Pakistan	1969	Low income	Pakistan
	Bangladesh	1974	Low income	
	Maldives	1976	Lower middle income	
	Brunei	1984	High income: non-OECD	
Sub-Saharan Africa	Chad	1969	Low income	Nigeria
	Guinea	1969	Low income	Côte d'Ivoire
	Mali	1969	Low income	
	Mauritania	1969	Low income	
	Niger	1969	Low income	
	Senegal	1969	Low income	

APPENDIX 2 (continued)

The OIC Member Countries and World Bank Classification based on Geographic Regions and Income Groups

Geographic regions	Full members ($N = 57$)	Membership Year	Income groups	OIC countries ($N = 14$) analyzed in this paper
Sub-Saharan Africa	Sudan	1969	Low income	Nigeria
	Somalia	1969	Low income	Côte d'Ivoire
	Sierra Leone	1972	Low income	
	Gabon	1974	Upper middle income	
	Gambia	1974	Low income	
	Guinea-Bissau	1974	Low income	
	Uganda	1974	Low income	
	Burkina Faso	1975	Low income	
	Cameroon	1975	Low income	
	Comoros	1976	Low income	
	Benin	1982	Low income	
	Nigeria	1986	Low income	
	Mozambique	1994	Low income	
	Togo	1997	Low income	
	Côte d'Ivoire	2001	Low income	

Note: The OIC members are divided among income groups according to 2004 gross national income (GNI) per capita. The groups are: low income, $825 or less; lower middle income, $826 – 3,255; upper middle income, $3,256 – 10,065; and high income, $10,066 or more) based on World Bank classification. Out of the 57 OIC member countries, Palestine is not included in World Bank's *World Development Indicators 2006* (WDI) database.