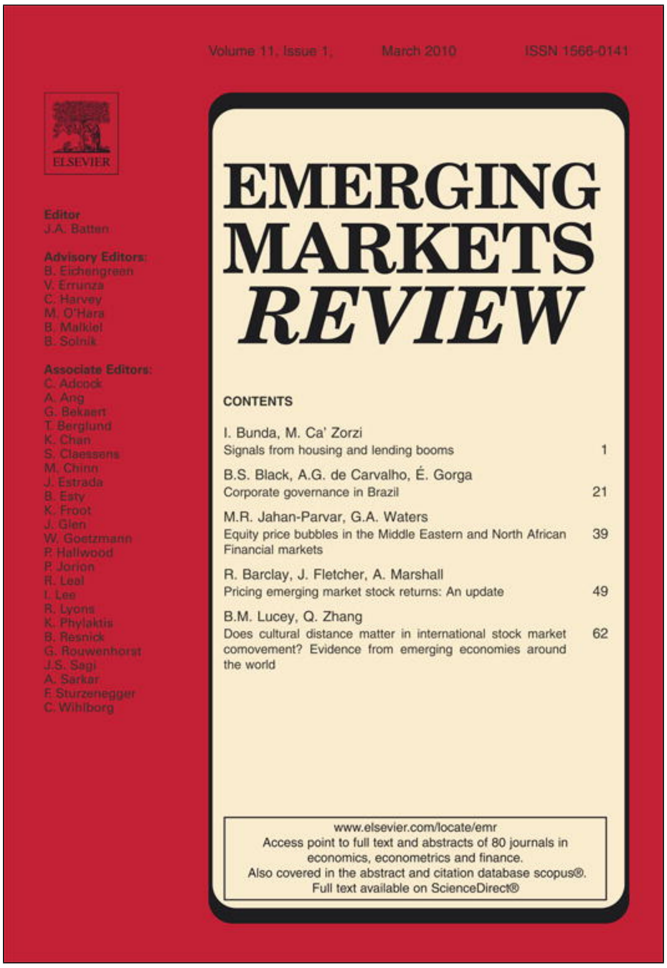
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Equity price bubbles in the Middle Eastern and North African Financial markets

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ABSTRACT

We empirically investigate the existence of periodically collapsing bubbles in seven Middle East and North African (MENA) financial markets for the period ending in May 2009. We use the Taylor and Peel (1998) residual augmented least square Dickey and Fuller test (RALS DF) to detect the bubbles. We find that the hypothesis of a bubble formation cannot be rejected for all seven markets investigated in our study, leading us to believe that in fact there has been a break down in the cointegration relationship between real equity prices and real dividends and also between real market capitalizations and real dividends.

1. Introduction

Between the turn of the century to mid-2008, Middle Eastern and North African (MENA) stock markets experienced astonishing performance. While returns on the Thomson Reuters' *Datastream World Index* had a monthly average equal to −0.18% between January 2000 and October 2008, MENA markets indices in our study posted monthly average returns ranging between a low of 0.64% for Turkey to a high of 1.56% for Oman. Similarly, market value of *Datstream World Index* for the same period had an average monthly growth rate equal to −0.19%, while market values for MENA country indices experienced monthly growth rates varying between a low of 0.23% for Tunisia to a high of 1.17% for Israel. These markets outperform

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world market indices by a larger margin if we focus on 2001–2007 period. Whether the behavior of these indices represents a bubble or is indicative of the expected future performance of fundamentals is an open question. In this paper, we formally address this question and test for formation of speculative bubbles in seven MENA equity markets in the period ending in May 2009. We present both statistical and descriptive evidence in support of our assertion that a speculative bubble formed in MENA equity markets studied in this paper.

The importance of the issue of the proper policy response to asset bubbles has been highlighted by the recent financial crisis. Chan et al. (2003) believe that in absence of rational bubbles, monitoring the market fundamentals in conducting monetary policy is sufficient. Otherwise, to divert expectations from the bubble path, positive policy action is needed. However, targeting financial bubbles as a reasonable policy for central banks is controversial.[2] In part, this controversy stems from the difficulty of detecting bubbles. This is what we have done for seven MENA equity markets. While bursting of financial bubbles in emerging and frontier markets may have a smaller global impact than the subprime crisis in the US, Asian financial crisis and Russian default episodes in the late 1990s warn us not to dismiss emerging markets in a globalized financial system. In particular, Parke and Waters (2007) demonstrate that the uncertainty about fundamentals is a major contributing factor to the formation of bubbles. Such uncertainty would tend to be present in maturing asset markets such as the MENA stock markets.

Hypothetically, it is possible to detect formation of bubbles by monitoring deviations from the "correct" price, based on fundamentals. One way of obtaining fundamentals-based prices is to use CAPM-type pricing models, and then compare the market outcomes and the model predictions, using a suitable measure for divergence. In the context of international markets, such an approach requires significant conditional correlations between the local and world index returns. One particular problem with the majority of MENA markets is weak correlations between MENA and world index returns. Cheng et al. (forthcoming) provide detailed documentation of asset pricing characteristics of nine MENA markets in CAPM setting (static, constant parameter intertemporal, and Markov switching variants), but do not study the possibility of bubble formation. They conclude that there is very strong evidence of segmentation in MENA markets from international financial system except for Israel, Turkey, and to a lesser extent, Bahrain. They confirm and document weak conditional correlations between MENA and world index returns.[3]

A crucial issue from our point of view is that Cheng et al., (forthcoming) findings suggest that since MENA markets are segmented from the world financial system, we cannot use international CAPM or its extensions to price returns from these markets. The majority of the MENA markets seem to price assets based on local information alone, as in Merton (1973). Hence, we cannot detect formation of bubbles based on CAPM-based pricing. Thus, formal testing for bubbles is required.

We introduce formal cointegration tests between price and dividends to detect equity price bubbles in seven MENA financial markets. Diba and Grossman (1988) argue that if bubbles are not present, prices and dividends should be cointegrated. Evans (1991) constructs a class of periodically collapsing bubbles that may not be detected by simple cointegration tests. Taylor and Peel (1998) introduce a test for cointegration that is robust to the skewness and excess kurtosis, and, hence, is able to detect such bubbles. Our conclusion is based on results from both types of tests.

In general, there are not many studies focused on MENA financial markets. In addition to Cheng et al. (forthcoming) discussed earlier, we briefly review some recent examples. Errunza (2001) focuses on the liberalization and integration of financial markets in Egypt, Israel, Jordan, Morocco, and Turkey. Ghysels and Cherkaoui (2003) study trading costs in Morocco. Lagoarde-Segot and Lucey (2008) study information efficiency in seven MENA markets and find heterogeneous levels of efficiency. Billmeier and Massa (2009) study the role of oil reserves, remittances, and institutions besides the traditional factors, and find that they appear to play a role in determination of market capitalization in MENA and Central Asian financial markets. Alsubaie and Najand (2009) investigate the informational role of trading volume in predicting the direction of short-term returns for the Saudi Stock Exchange.

[2] See Reuters, September 28, 2009. Bank of Canada Governor Mark Carney opined on the debate among central bankers in August 2009 in Jackson Hole, Wyoming, on the use of monetary policy against credit growth and asset bubbles, and whether such a course of action is compatible with inflation targeting.

[3] They follow Bekaert and Harvey (1995) in their definition of segmentation and integration and apply a very similar estimation method to measure the degree of integration of the MENA markets in the global financial system.

Still fewer studies investigate the possibility of bubble formation in MENA markets. In a paper related to our work, Billmeier and Massa (2008) study the possibility of non-cointegration between Egyptian stock market index and the underlying fundamentals. They find that this possibility cannot be ruled out. Their work is focused on a single market. We, on the other hand, study a more diverse set of markets and are formally looking for evidence in favor of bubble formation. We are not aware of any other recently published paper on MENA equity markets that directly addresses speculative bubbles.

The rest of the article proceeds as follows. In Section 2 we briefly discuss the models used in detection of asset price bubbles, review the Taylor and Peel (1998) methodology, and discuss the estimation equations and variables. In Section 3 we introduce the data. Section 4 contains presentation and discussion of our main empirical findings. Section 5 concludes.

2. The model

The standard present value model of the stock prices is often presented as

$$P_t = \frac{1}{1+r}\mathbb{E}_t(P_{t+1} + D_{t+1}), \tag{1}$$

where P_t is the real stock price at time t, D_{t+1} is the real dividend paid between t and $t+1$, and E_t denotes the expectation operator for information at time t, as in Campbell et al. (1997). In this formulation discount factor, $0<(1+r)^{-1}<1$, is assumed to be constant. If we impose the transversality condition $\lim_{n\to\infty}(1+r)^{-n}\mathbb{E}_t P_{t+1}=0$, then Eq. (1) has a unique solution of the form:

$$F_t = \sum_{j=1}^{\infty} \frac{1}{(1+r)^j}\mathbb{E}_t D_{t+j} \tag{2}$$

where $P_t = F_t$.

Together, these equations imply that

$$P_t - \frac{1}{r}D_t = \frac{1+r}{r}\sum_{j=1}^{\infty}\frac{1}{(1+r)^j}\mathbb{E}_t \Delta D_{t+j}. \tag{3}$$

This equation implies that if both P_t and D_t are generated by $I(1)$ processes, then $P_t - r^{-1}D_t$ is cointegrated and the parameter of cointegration is equal to r^{-1}.

If the above mentioned transversality condition fails to hold, then $P_t = F_t$ instead of being the unique solution to Eq. (1), is just one of potentially infinite solutions which belong to the class given by

$$P_t = F_t + B_t, \tag{4}$$

see Taylor and Peel (1998). In this class, B_t represents a rational bubble term, which must satisfy

$$B_t = \frac{1}{1+r}\mathbb{E}_t B_{t+1}. \tag{5}$$

If these bubbles are non-zero, then Eq. (3) must be augmented by B_t. This rules out cointegration between P_t and D_t since in general, B_t are not stationary and lead to explosive conditional expectations for the $P_t - r^{-1}D_t$ process.

Based on this observation, Diba and Grossman (1988) propose that testing for non-cointegration between real stock prices and dividends, combined with unit root tests for real stock prices and dividends and their first differences, can be interpreted as a test for detection of bubbles.

Evans (1991) introduces a class of periodically collapsing bubbles which cannot be detected using Diba and Grossman (1988) methodology. This class can be formalized as

$$B_{t+1} = (1+r)B_t \nu_{t+1}, \text{ if } B_t < \alpha \tag{6}$$

$$B_{t+1} = \left[\delta + \frac{(1+r)}{\pi}\theta_{t+1}\left(B_t - \frac{\theta}{(1+r)}\right)\right]\nu_{t+1}, \text{ if } B_t > \alpha. \tag{7}$$

In these equations, α and δ are positive parameters where $(1+r)\alpha>\delta>0$, θ_t is an *iid* Bernoulli process which takes the value 1 with probability π and 0 with probability $1-\pi$ where $1\geq\pi>0$ and is viewed as the probability of the continuation of the bubble, and ν_t is an *iid* positive random variable independent of θ_t such that $\mathbb{E}_t\nu_{t+1}=1$. This class of bubbles admit partial collapses with probability one, are strictly positive, and do not vanish. Hence they satisfy the stylized requirements of stock price bubbles. Most importantly, Evans (1991) by using Monte Carlo simulations shows that application of standard cointegration tests often leads to failure to reject the stationarity of periodically collapsing bubble processes, since standard tests 'mistake' sudden collapse with mean reversion.

The estimation equation follows the simple linear form of

$$P_t = \beta_0 + \beta_1 D_t + \epsilon_t, \tag{8}$$

and the important issue is the stationarity of the residuals. As is well known, to have stationarity, one needs $|\beta_1|<1$. Waters (2008) argues that the proper test for periodically collapsing bubbles uses log prices and dividends. Furthermore, that paper demonstrates that simple cointegration tests using logs are able to detect the class of bubbles introduced by Charemza and Deadman (1995). We present results using both levels and logs.

We briefly describe the Taylor–Peel estimator here. One salient point of this method is incorporation of skewness and excess kurtosis in the construction of the estimator. Most cointegration-based tests for rational bubbles rely on testing on the residuals of Perron (1989) regression, as

$$\Delta\hat{\epsilon}_t = \psi\hat{\epsilon}_{t-1} + u_t \tag{9}$$

where the null hypothesis of no cointegration implies $\psi=0$ and the alternative of a stationary residual requires $\psi<0$. Taylor and Peel (1998) correct the least squares estimate in Eq. (9) for skewness and excess kurtosis to first obtain a more efficient estimator of ψ, and second, to increase the power of the test to correctly reject a mean-reverting error as a bubble, in comparison to the standard cointegration tests. Their method is a two-step estimation procedure. First, regress the first difference of the residuals of the cointegrating equation on their lagged levels, as in Eq. (9). Use the new residuals, $\hat{u}_t$, and the estimated variance, $\hat{\sigma}^2$, to construct the vector $\hat{w}_t=[(\hat{u}_t^3-3\hat{\sigma}^2\hat{u}_t)(\hat{u}_t^2-\hat{\sigma}^2)]'$. Notice that the first element of this vector is the skewness and the second element is the excess kurtosis of the residual. In the second step, re-estimate Eq. (9) with the addition of vector $\hat{w}_t$, which corrects for skewness and excess kurtosis of the residuals following

$$\Delta\tilde{\epsilon}_t = \psi\tilde{\epsilon}_{t-1} + \phi\hat{w}_t + \nu_t. \tag{10}$$

In this equation, ν_t follows a white noise process. This method delivers a residual-augmented least squares Dickey–Fuller (RALS DF) test of no cointegration. The key test statistic here is $CR\tau_A = \hat{\psi}/\sqrt{\mathrm{Var}(\hat{\psi})}$. Here, $\hat{\psi}$ is the estimator in Eq. (10) and $\mathrm{Var}(\hat{\psi})$ which is the variance–covariance matrix of $\hat{\psi}$, is given in pages 223 and 224 of Taylor and Peel (1998). Taylor and Peel (1998) denote standard cointegrating Dickey–Fuller statistic by $CR\tau$.

Almost all studies of rationally collapsing bubbles look at cointegration between real asset prices and real dividend payments. Diba and Grossman (1988), Evans (1991), Charemza and Deadman (1995), Taylor and Peel (1998), Bohl (2003), Doffou (2008), among many others use price index levels as the proxy for P_t.

Hence the estimation equation is of the form introduced in Eq. (8). In this formulation, we rely on the relationship between market activity, captured by the level of the real price index, P_t, and real dividends, D_t. Following Waters, (2008), this relationship needs to be expressed in logarithmic values for testing stochastic explosive unit root models such as Evans (1991). We substitute P_t and D_t in Eq. (8) by their natural logarithmic values, p_t and d_t respectively.

However, an alternative formulation exists, based on the market value of an index. Market value reflects both the fluctuations in the price level and the volume of tradeable shares. Hence, it also acts as a measure for the market size or market capitalization. If we are interested in the possibility of a bubble in prices, we believe that it is reasonable to study the behavior of the aggregate market as well as price behavior alone, since price increases may be caused by a decrease in the quantity of equity available for trade due to, for example, a share buy back program. Hence we propose to consider the following relationship as well as the familiar Eq. (8). In this context, we substitute P_t by MV_t, which is the real market value at time t. In the logarithmic relationship, we use mv_t which is the natural log of MV_t.

3. Data

We use real monthly data in 2005 US dollars from seven MENA financial markets obtained from Thomson Reuters' *Datastream*: Egypt, Israel, Lebanon, Morocco, Oman, Tunisia, and Turkey. The source for the data is Standard and Poor's/International Finance Corporation (S&P/IFCG). We look at price index (P_t), market value (MV_t), and dividends (D_t) series from these markets. We use US dollar denominated values to maintain uniformity of results. While it would have been optimal to include more countries, we are severely restrained by data availability. For example, short length of available data from the majority of (Persian) Gulf Cooperation Council (GCC) countries, such as Saudi Arabia, Kuwait, and United Arab Emirates in S&P/IFCG data bank, excludes them from our study. Arab countries in our study, with the exception of Oman which is a minor oil producer, can be categorized as "Mediterranean" following Rauch and Kostyshak (2009) example. These economies are not dependent on hydrocarbon exports as their main source of income. Many of them rely on remittances (for example, Egypt) or are active trading countries (for example, Lebanon).

Price indices are value weighted indices of traded equities in the respective market. Market values are the product of the price of constituent index stocks times the number of stocks available for trading, and thus is a measure of market capitalization of the index. Dividend variables reflect the aggregate paid dividend of constituent stocks of each index. Lebanon's dividend data contains significant number of zero entries. Some, but not all, of these entries pertain to the summer of 2006 war. Due to this reason, we exclude Lebanon from analysis of logarithmic values of variables.

Table 1 reports summary statistics for the data in this study. The length of series is not equal across countries. It ranges between December 1987 and October 2008 for Turkey which yields 250 observations

Table 1
Summary statistics of the data.
Source: Thomson Reuters' *Datastream*.

	Dates	No. obs.	Variables	Mean	Std. dev.	Skewness	Kurtosis	Min	Max
Egypt	02/25/97 to	141	P_t	183.87	147.66	1.13	0.31	37.60	605.54
	11/25/2008		MV_t	12,116.67	10,755.69	1.13	0.15	1,901	42,915
			D_t	344.07	187.98	0.69	−0.15	21.58	926.99
Israel	2/25/1997 to	141	P_t	180.40	62.60	0.71	−0.52	94.75	336.28
	11/25/2008		MV_t	39,957.78	23,892.80	0.93	−0.31	5021	99,919
			D_t	867.44	706.45	1.24	0.62	102.93	3,057.19
Lebanon	2/25/2000 to	112	P_t	108.10	65.17	0.98	0.23	44.03	318.67
	5/25/2009		MV_t	2,592.61	1,843.83	1.13	0.52	926	8619
			D_t	23.96	49.13	2.57	6.65	0.00	238.75
Morocco	2/25/1997 to	141	P_t	282.83	181.37	1.71	1.76	125.11	825.55
	10/25/2008		MV_t	9,445.50	6,876.28	1.84	2.50	652	31172
			D_t	232.16	117.90	2.19	7.58	27.38	828.95
Oman	1/25/2000 to	113	P_t	183.49	110.27	1.19	1.02	65.69	513.42
	5/25/2009		MV_t	4,232.11	2,744.28	1.09	0.44	1,229.83	12,001
			D_t	183.10	127.42	1.35	1.48	7.30	596.03
Tunisia	1/25/1997 to	149	P_t	58.27	18.24	1.08	−0.01	36.77	107.40
	5/25/2009		MV_t	1,322.62	339.25	0.76	0.02	779	2,341
			D_t	45.82	13.63	0.01	0.92	8.73	86.21
Turkey	12/25/1987 to	250	P_t	564.00	380.03	1.26	0.95	117.71	1,888.78
	10/25/2008		MV_t	22,697.97	17,011.97	1.02	0.63	535.70	78,464
			D_t	542.56	398.32	1.34	2.24	18.37	2,052.82
World	1/25/1986 to	281	P_t	778.16	351.72	0.68	−0.34	216.00	1696.17
	5/25/2009		MV_t	19,244,659	12,278,417	0.64	−0.54	2,758,000	49,846,690
			D_t	399,317.96	289,444.94	1.30	1.14	78,327.20	1,545,522.37

Notes: Variables P_t, MV_t, and D_t represent deflated price index, real market value in millions of US dollars, and real paid dividends in millions of US dollars. Prices are deflated using GDP deflator in 2005 base.

per series on one hand, to 112 observations per series in the case of Lebanese data which spans February 2000 to May 2009 period. All reported data are end of the month recorded values. The following properties of the data are worth noting. First, unconditional standard deviations of P_t are either in the same order of magnitude or an order of magnitude smaller than unconditional means. For MV_t and D_t, both unconditional means and standard deviations are of the same order of magnitude. Second, all variables demonstrate negligible unconditional skewness. Third, all variables show negligible excess kurtosis at the level.[4]

4. Empirical findings

We examine the stochastic properties of the price index, market value, and dividend series from each country separately. In the first step, we test for stationarity and the order of integration using Augmented Dickey Fuller (ADF) method introduced in Dickey and Fuller (1979) and expanded in Said and Dickey (1984).

The results are reported in Table 2. As expected, the null hypothesis of the existence of a unit root is not rejected for price index and market value data across all countries. The null hypothesis of a unit root is rejected for dividend series for Lebanon and Morocco, leading us to believe that dividend series are stationary in the Lebanese and Moroccan data. As expected, the null hypothesis of a unit root is rejected for log difference values for all three variables across all countries, which is evidence for stationarity at the first difference. These results are not reported but are available upon request.

The fact that in Lebanon and Morocco series asset prices are of the order $I(1)$ and dividend payments are of the order $I(0)$, is indicative of the existence of speculative bubbles in the aforementioned markets. Due to difference in orders of integration, cointegration tests are misleading on the data from these two markets.

For testing the presence of cointegration between data series, we perform Johansen and Juselius (1990) trace-based test. These results are reported in Table 3. As is seen in Table 3, the null hypothesis of no cointegrating vector between P_t and D_t or MV_t and D_t variables is rejected for the majority of the markets studied.

The exception is Tunisia. We fail to reject the null hypothesis of no cointegration between P_t and D_t variables, but we reject this null hypothesis for MV_t and D_t. Also, as mentioned earlier, due to different orders of integration between dividend and equity price proxies, ordinary cointegration tests are not to be trusted for Lebanon and Morocco.

Testing for the existence of one cointegrating vector in natural log specification, yields similar results. We reject the null hypothesis of no cointegrating vector between p_t and d_t in all markets except Tunisia. The same testing procedure is carried out for mv_t and d_t, and in all markets we reject the null hypothesis of no cointegrating vector.

In sum, Johansen and Juselius cointegration tests indicate that price index and dividends or market value and dividends are cointegrated in the majority of MENA markets studied here. This procedure, along with stationarity results reported in Table 2, indicates that following Diba and Grossman (1988), we cannot rule out the formation of a rational bubble between market values and dividends in Tunisian data. To a lesser extent, we are unable to rule out a rational bubble in Lebanese and Moroccan data since dividends seem to be stationary while price measures seem to be non-stationary, ruling out cointegration.

This may be interpreted as absence of rationally collapsing bubbles in the rest of the markets in our study. But as noted earlier, conventional cointegration tests are often unable to detect periodically collapsing bubbles found in Evans (1991). Hence we need to carry out further testing to rule out formation of bubbles in MENA markets in the period under study.

As discussed earlier, we use Taylor and Peel (1998) method in our study for detection of rationally collapsing bubbles. These results are reported in Table 4. The null hypothesis pertaining to test statistics reported in the first and the fourth columns of this table is no cointegration between dividends and price index/market values. The null hypothesis for student t-statistics reported in columns two, three, five, and six is a simple H_o: $\phi_i = 0$, where $i = 1,2$. This hypothesis means that we are testing whether incorporation of skewness and kurtosis in Eq. (3), which yields Eq. (10), is statistically significant.

[4] Price index returns (log differences) and percentage changes in market value and dividends demonstrate significant excess kurtosis, as is expected in financial markets.

Table 2
Stationarity test results.
Source: Thomson Reuters' *Datastream*.

Country	Variable	ρ	$Pr<\rho$	τ	$Pr<\tau$
Egypt	P_t	−5.08	0.8106	−1.83	0.6848
	MV_t	−5.46	0.7807	−1.75	0.7243
	D_t	−9.86	0.4325	−2.21	0.4774
	p_t	−2.56	0.9524	−1.50	0.8259
	mv_t	−3.51	0.9115	−1.44	0.8456
	d_t	−8.60	0.5256	−2.37	0.3930
Israel	P_t	−10.74	0.3730	−2.12	0.5286
	MV_t	−12.49	0.2739	−2.49	0.3304
	D_t	−3.08	0.9321	−1.11	0.9228
	p_t	−11.30	0.3388	−2.23	0.4702
	mv_t	−53.36*	0.0005	−5.07*	0.0003
	d_t	−8.99	0.4955	−2.11	0.5361
Lebanon	P_t	−16.15	0.1307	−2.86	0.1781
	MV_t	−15.45	0.1508	−2.82	0.1924
	D_t	−25.25*	0.0168	−3.47*	0.0473
	p_t	−7.26	0.6311	−2.09	0.5471
	mv_t	−7.13	0.6421	−2.12	0.5292
Morocco	P_t	−3.58	0.9078	−1.37	0.8648
	MV_t	−1.03	0.9873	−0.51	0.9819
	D_t	−46.51*	0.0005	−4.68*	0.0012
	p_t	−1.67	0.9768	−0.88	0.9549
	mv_t	−6.84	0.6688	−1.70	0.7475
	d_t	−48.71*	0.0005	−4.91*	0.0005
Oman	P_t	−13.69	0.2145	−2.42	0.3662
	MV_t	−7.01	0.6520	−1.62	0.7796
	D_t	−20.14	0.0553	−3.12	0.1059
	p_t	−8.56	0.5255	−1.92	0.6369
	mv_t	−6.03	0.7340	−1.68	0.7520
	d_t	−35.99*	0.0011	−4.20*	0.0062
Tunisia	P_t	−6.15	0.7265	−2.09	0.5485
	MV_t	−12.52	0.2730	−2.89	0.1674
	D_t	−17.19	0.1096	−2.90	0.1649
	p_t	−6.08	0.7320	−2.14	0.5180
	mv_t	−10.50	0.3893	−2.56	0.2983
	d_t	−19.79	0.0631	−3.12	0.1067
Turkey	P_t	−12.25	0.2937	−2.45	0.3513
	MV_t	−16.74	0.1259	−2.85	0.1801
	D_t	−11.79	0.3181	−2.00	0.5972
	p_t	−14.92	0.1795	−2.72	0.2275
	mv_t	−12.53	0.2792	−2.59	0.2843
	d_t	−15.7	0.1524	−2.91	0.1605

Notes: This table reports stationarity results for augmented Dickey–Fuller (ADF) tests for the levels and logarithmic levels of price index, market values, and dividend payments. * Denotes failure to reject the null hypothesis of a unit root at 5% or better confidence level. Since Lebanese dividend payment data includes zero entries at the level, stationarity tests for this data series cannot be carried out. The null hypothesis of a unit root at 5% or better confidence level is soundly rejected for first differences of all series. These results are not reported, but are available upon request.

The left hand side panel (Panel A) of Table 4 reports the estimated RALS DF statistics ($CR\tau_A$) for Eq. (8) using P_t and D_t as variables to be tested, along with values of student t-statistics associated with estimated $\hat{\phi}_1$ and $\hat{\phi}_2$, from estimation of relevant Eq. (10) for the price index and dividend relationship. We report 5% critical values for RALS DF $CR\tau_A$ and ordinary cointegrating Dickey and Fuller statistics from Taylor and Peel (1998). Their sample size is 116 observations, which is slightly smaller than our sample. On the other hand, the estimated values of $CR\tau_A$s in our sample are so small that we reasonably believe that failure to reject the null would not be affected at reasonable statistical confidence levels.

The right hand side panel (Panel B) of Table 4 reports the same three sets of estimated statistics for dividend and market value series. Similar to the previous discussion, we obtain extremely small $CR\tau_A$

Table 3
Cointegration test results.
Source: Thomson Reuters' *Datastream*.

Country	Vars	H_o	Eigenvalue	Trace
Egypt	P_t, D_t	$r=0$	0.1857	29.465*
	MV_t, D_t	$r=0$	0.1493	22.990*
	p_t, d_t	$r=0$	0.3945	70.185*
	mv_t, d_t	$r=0$	0.4652	87.518*
Israel	P_t, D_t	$r=0$	0.1837	30.476*
	MV_t, D_t	$r=0$	0.2927	50.629*
	p_t, d_t	$r=0$	0.2084	35.138*
	mv_t, d_t	$r=0$	0.3731	69.705*
Morocco	P_t, D_t	$r=0$	0.2272	40.181*
	MV_t, D_t	$r=0$	0.2229	37.483*
	p_t, d_t	$r=0$	0.2428	43.418*
	mv_t, d_t	$r=0$	0.3807	70.995*
Oman	P_t, D_t	$r=0$	0.2405	32.570*
	MV_t, D_t	$r=0$	0.2657	36.067*
	p_t, d_t	$r=0$	0.1584	20.761*
	mv_t, d_t	$r=0$	0.1676	21.333*
Tunisia	P_t, D_t	$r=0$	0.1098	18.441
	MV_t, D_t	$r=0$	0.1150	34.199*
	p_t, d_t	$r=0$	0.1084	18.273
	mv_t, d_t	$r=0$	0.1282	32.627*
Turkey	P_t, D_t	$r=0$	0.1490	42.000*
	MV_t, D_t	$r=0$	0.1832	52.942*
	p_t, d_t	$r=0$	0.1334	38.038*
	mv_t, d_t	$r=0$	0.1900	58.372*
Lebanon	P_t, D_t	$r=0$	0.2586	42.466*
	MV_t, D_t	$r=0$	0.2616	41.977*

Notes: This table reports (Johansen and Juselius, 1990) trace-based cointegration test results. Imposed restriction is a constant drift in the process. A constant drift in the Error Correction Model (ECM) is assumed. The null hypothesis, H_o: $r=0$ assumes no cointegrating vector. The null hypothesis, H_A: $r>0$ assumes at least one cointegrating vector. The critical value for the test at $\alpha=5\%$ is 19.99. * Denotes rejection of the null at the 5% level, which indicates existence of one cointegrating vector.

Table 4
Results of Taylor and Peel (1998) test for detection of rationally collapsing bubbles.
Source: Thomson Reuters' *Datastream*.

Country	$CR\tau_A$	Skewness *t*-stat	Kurtosis *t*-stat	$CR\tau_A$	Skewness *t*-stat	Kurtosis *t*-stat
Panel A: P_t and D_t				*Panel B: MV_t and D_t*		
Egypt	−2.17E−06	1.6285	1.7975†	−4.50E−10	1.7499†	1.9910*
Israel	−4.67E−05	6.1954*	−1.8299†	−2.97E−10	10.8426*	−2.7721*
Lebanon	−1.07E−06	6.3545*	−2.5449*	−4.68E−09	8.6831*	1.9151†
Morocco	−1.90E−06	9.0661*	−3.4457*	−1.14E−09	8.9734*	−3.3338*
Oman	−4.72E−06	8.4474*	−1.8403†	−1.15E−08	7.6571*	−0.5020
Tunisia	3.23E−05	1.8412†	−1.9406†	−1.69E−07	13.7115*	−2.9687*
Turkey	−2.57E−07	7.7429*	−0.1850	−4.19E−11	7.4288*	−1.0009
Panel C: p_t and d_t				*Panel D: mv_t and d_t*		
Egypt	−0.0315	3.2710*	0.8711	−0.0310	6.7255*	0.8116
Israel	−1.0210	7.8438*	−3.5686*	−0.1424	13.2959*	1.6625†
Lebanon						
Morocco	0.0907	14.7490*	−2.6447*	−0.0781	9.9493*	−2.0707*
Oman	−0.0571	12.7421*	0.2907	−0.0626	12.9317*	−1.0048
Tunisia	−0.0024	5.1567*	−0.1478	−0.2760	12.0845*	−2.2507*
Turkey	−0.0554	1.3082	−0.2415	−0.0043	1.2586	2.2651*

Notes: This table reports test results from applying Taylor and Peel (1998) test procedure to the data. Five percent critical value for RALS DF, $CR\tau_A$, is −3.790 and for standard cointegrating DF, $CR\tau$, is −3.242. Skewness and kurtosis *t*-statistics pertain to values associated with $\hat{\phi}_1$ and $\hat{\phi}_2$ estimated parameters in Eq. (10). † and * pertain to failure to reject the null hypothesis of H_o: $\phi_i=0$ where $i=1,2$ from Eq. (10) at $\alpha=10\%$ and $\alpha=5\%$ respectively.

values. These values are in fact considerably smaller even in comparison with what is reported in Panel A. Again, the null of no cointegration cannot be rejected.

In both panels, it can be seen that the majority of reported t-statistics are statistically significant at the conventional $\alpha = 5\%$. Thus, we can conclude that inclusion of skewness and kurtosis in Eq. (10) is warranted.

The remaining two panels in Table 4, namely Panels C and D, report the Taylor and Peel (1998) test results for Eq. (8) when variables p_t and d_t, and mv_t and d_t are used. Again, inclusion of skewness and kurtosis in Eq. (10) is warranted. Moreover, estimated values for $CR\tau_A$ statistic are very small, leading to failure to reject the null hypothesis of no cointegrating vector.

It is clearly seen from this table, we cannot reject the null of no cointegration, given the extremely small values of estimated $CR\tau_A$s. This leads use not to rule out the existence of bubbles in equity prices in MENA stock markets studied for the late 1990s to 2008 period. We acknowledge that failure to find a cointegrating vector, hence no cointegration, does not provide a final answer to the existence of rational bubbles in equity markets.

But this is a very strong indication, which is borne by the fact that the collapse of equity prices in these markets in post-2008 period was not accompanied by a similar collapse in dividend payments. As an example, consider the behavior of S&P/IFCG Israel index and the aggregate dividend payments associated with this index. Between end of the May of 2008 and the end of May 2009, the index fell by 45.78%, from 318.84 to 218.72. In the same period, dividend payments fell only by 29%. Lebanon presents a more dramatic example. In the same time period discussed for Israel, paid dividends of S&P/IFCG Lebanon index rose from 23.80 to 107.56 million USD. Meanwhile, Lebanon index fell from 229.62 to 166.03, or a decrease of 27.70%.

Based on the econometric evidence and descriptive evidence presented up to this stage, we feel comfortable to conclude that based on Taylor and Peel (1998) method, we cannot rule out a financial market bubble in the seven MENA markets studied.

5. Conclusion

In this paper, we formally address an open question in emerging markets finance literature. We investigate whether rationally collapsing bubbles can be viewed as an explanatory factor for the unusually bullish performance of the MENA financial markets in the period ending in the first decade of the 21st century. We conclude that based on our statistical findings and descriptive evidence presented, such a hypothesis cannot be ruled out.

We believe that based on the work of Cheng et al. (forthcoming), it is hard or even impossible to assess the performance of MENA markets based on their static or dynamic relationship with composite world financial market price indices, since these markets are generally segmented from the global financial system. Hence, detection of statistically significant divergences from CAPM-based return predictions is hard. As a result, we believe that to evaluate the performance of these markets, formal testing for rationally collapsing bubbles is needed. We carry out this task by following the methodology of both Diba and Grossman (1988) and Taylor and Peel (1998). Based on Diba and Grossman methodology, four out of the seven MENA financial markets studied have price series which seem to be cointegrated with dividend series. The hypothesis of the absence of a rational bubble cannot be rejected except for Tunisia, and to a lesser extent for Lebanon and Morocco.

Since Evans (1991) shows that conventional cointegration methodology fails in the face of periodically collapsing bubbles, we also test for this class of bubbles. Using Taylor and Peel (1998) methodology to test for periodically collapsing bubbles, we find that the null hypothesis of non-cointegration between prices and dividends, which is evidence of a bubble, cannot be rejected at any reasonable statistical level for all markets in our sample. Along with the descriptive evidence of market performance since October 2008, we find this outcome to be supportive of bubbles in MENA financial markets. Our results are of interest to financial scholars conducting research on emerging and frontier markets, investors seeking global opportunities, and international and national policy makers with an interest in detection or taking action against financial bubbles.

The dramatic collapse of stock markets in many Western economies in 2008, particularly in the U.S., have led to serious consequences for the world economy in 2008–09 period. These events highlight the need for testing the vulnerability to bubble formation in any asset market that is integrated with the broader economy. Our results demonstrate that MENA stock markets are not immune to such concerns. Our findings are *ex post* in nature and pertain to markets with relatively short trading histories. But we believe that given the political and economic sensitivity of the MENA region, it is very important to monitor for the formation of financial asset bubbles in these countries in order to avoid potential political or economic instability which may spill over to other markets, either via financial or economic contagion, or through political unrest and conflict.

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