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Testing the weak-form market efficiency and the day of the week effects of some African countries[♦]

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Abstract

The aims of this work are twofold. On the one hand, it aims to find evidence supporting the presence of the weak form efficiency of several emerging African stock markets by using both parametric as well as non parametric tests. The results indicate that none of the markets are characterised by random walks with the exception of the South African stock market. On the other hand, this study aims to detect the presence of the day of the week effects of these African stock markets. Results show the existence of day of the week effects, that is the typical negative Monday and Friday positive effects in several stock markets.

Keywords: African stock markets, random walk hypothesis, day of the week effects.
JEL classifications: G14, G15.

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1. Introduction

A central issue of the literature of finance is the efficient market hypothesis (EMH). This theory argues that if stock prices reflect all the information available and immediately incorporate all new information then the market can be considered efficient. Fama (1970) defines three types of efficiency that is weak-form efficiency, semi-strong-form efficiency and strong-form efficiency. In a weak-form efficient market, past prices are not relevant in achieving excess returns. Semi-strong-form efficiency implies that prices reflect publicly available information so that no investor can earn excess returns based on any publicly available information. Strong-form efficiency means that stock prices reflect all information so that investors cannot earn excess returns using any information, whether publicly available or not.

Investors take into account whether stock markets are or are not weak form efficient because return predictability may by a source of higher profits. Emerging markets are often characterized by a lower volume and frequency of trading and easiness of manipulation by a few larger traders. If correct information fails to be quickly and fully reflected in the stock prices then stock markets are said to be inefficient because those who are privy to such information can benefit by anticipating the course of such prices. As pointed out by Borges (2007), testing the EMH is relevant for investors as well as regulatory authorities and policymakers. The former are interested in setting up investment strategies in order to diversify their investment portfolios and finding new opportunities for profit. Regulatory authorities and policymakers are also interested in EMH because the lack of efficiency in stock markets does not allow mechanism prices to work correctly. In other words the allocation of capital is not efficient with a negative effect for the overall economy.

The aim of this paper is to find evidence of the EMH in several emerging African stock markets, that is Egypt, Morocco, South Africa and Tunisia. Two reasons explain the increasing attention of practioners and academics in these emerging markets. The first is due to the fact that the increasing globalization of financial markets make emerging markets one of the possible opportunities for investment for international funds seeking new opportunities. The second reason is that relatively few works have been focused on these markets.

Empirical literature on stock markets has focused on developed equity markets, while relatively few studies have focused on the emerging markets. The results of these last few studies have also been conflicting. Ojah and Karemera (1999) using both the multiple variance ratio test as well as the autoregressive fractionally integrated moving-average test, found evidence that the random walk hypothesis[1] (RWH hereafter) was not rejected for the emerging markets of Argentina, Brazil, Chile and Mexico. Whortington and Higgs (2003) had opposite results. By using unit root tests, multivariance test statistics and runs tests, they found that the stock markets of Argentina, Brazil, Columbia, Mexico, Peru and Venezuela are not weak form efficient.

Weak form efficiency characterizes most of the Asian emerging equity markets. Abraham et al. (2002), examined the weak form efficiency of the stock markets of Bahrain, Kuwait and Saudi Arabia by using both the variance ratio test and runs test: these tests show that the random walk hypothesis is rejected when the index levels are used. Correcting this index by the Beveridge and Nelson (1981) decomposition, they found that these markets are weak form efficiency. Marashdeh and Shrestha (2008) by using Augmented-Dickey Fuller and Phillip-Perron tests showed that the United Arab Emirates Securities Market is weak form efficient.

In some cases weak form efficiency of stock markets may be achieved by specific steps taken by national institutions. For instance, Islam and Khaled (2005), found evidence that the Dhaka Stock Market (DHA) returns behaved differently before and after the 1996 stock market crash. Predictability of stock returns seemed to characterize the period after the 1996 crash, while after these events, DHA returns have followed a random walk. In other words this market seems to be weak form efficient. These changes are probably due to several rules introduced by the Bangladeshi Security Commission in order to increase the transparency in the stock market.

A few empirical studies have been conducted relative to African stock markets. For example Olowe (2002) tested the hypothesis of weak form efficiency on the monthly stock prices of 59 Nigerian companies by using the autocorrelation function test. Results showed that security returns were independent, in other words the Nigerian stock market seemed to be efficient in a weak form. Al-Khazali et al (2007), by using rank and sign tests, the runs test, and the conventional VR test showed that MENA[2] emerging equity markets were weak form efficient after having corrected the individual market indices for the statistical biases arising from thin and infrequent trading. EMH was also explored by Jefferis and Smith (2005) through a test of evolving efficiency for six African stock markets (that is Egypt, Kenya, Morocco, Mauritius, Nigeria and South Africa) and weekly closing price indices for the time period covering January 1990 through June 2001. Their results indicate that only the South African stock market was efficient during the full period considered while Egypt, Morocco, and Nigeria became weak form efficient towards the end of the period.

[1] The random walk hypothesis implies that successive stock market prices are random and serially independent. The rejection of the random walk hypothesis implies that investors can earn profits from forecasting future stock prices.

[2] The Middle East and North African (MENA) countries are the following: Bahrain, Egypt, Jordan, Kuwait, Morocco, Oman, Saudi Arabia, and Tunisia.

Appiah-Kusi and Menayah (2002) evaluated EMH hypothesis in 11 African markets. Their results showed that 6 out of 11 stock markets are weak form efficient.

In this work we also want to investigate market anomalies such as the day of the week effect, the monthly and the January effect. The day of the week effect is a phenomenon that constitutes a form of anomaly of the efficient capital market theories. According to this phenomenon, the average daily return of the market is not the same for all days of the week, as we would expect on the basis of the efficient market theory. Most of the studies analysing the days of the week effect have focused on developed stock markets. For the USA, the UK and Canada, most of the studies (Keim and Stambaugh, 1984; Board and Sutcliffe, 1988; Tang and Kwok, 1997) have shown that on Mondays these markets have statistically significant returns while on Fridays statistically positive returns.

Relatively little attention has been given to emerging stock markets. Studies have focused mainly on European emerging markets (Alexakis and Xanthakis, 1995; Balaban, 1995, Coutts et al., 2000; Al-Loughani and David, 2001), Pacific basin (Wong, 1995) and Asian stock markets (Choudhry, 2000). Results obtained are somewhat mixed, some evidence of the presence of the day of the week as well as the weekend effect was shown so indicating the existence of market inefficiency. A smaller amount of studies have focused on African emerging markets. For instance Alagidede (2008) investigated the day of the week effect for several African Stock markets. Some of them are not characterized by the presence of the day of the week while other markets are just characterized by daily seasonality.

Following these results, this paper aims to shed some further light on the existence of calendar anomalies in the emerging African market group as defined previously..

The rest of this paper proceeds as follows. Section 2 provides an overview of the methodologies for assessing the EMH as well as the day of the week effect. Section 3 identifies the data source. Section 4 presents the empirical results. A final section summarizes the conclusions.

2. Methodology

If the stock market is inefficient in the weak form, then it implies that market prices do not follow a random walk. Random walk requires that the time series must contain a unit root. Therefore we started by testing for the presence of a unit root in the stock market equity prices series. We employed the Augmented Dickey Fuller (ADF) test and the Phillips-Perron (PP) test. The ADF test assumes that y series follows an AR(p) process and add p lagged difference terms of the dependent variable y to the right side of the test regression. ADF test uses the following three models.

$$\Delta y_t = c_0 + c_1 t + \delta y_{t-1} + \beta \sum_{i=1}^{p} \Delta y_{t-i} + u_t \tag{1}$$

$$\Delta y_t = c_0 + \delta y_{t-1} + \beta \sum_{i=1}^{p} \Delta y_{t-i} + u_t \tag{2}$$

$$\Delta y_t = \delta y_{t-1} + \beta \sum_{i=1}^{p} \Delta y_{t-i} + u_t \tag{3}$$

Equation 1 is the first model, it includes a constant term c_o, a trend term c_1t,. The second model (equation 2) includes a constant term only, and the third model does not include intercept and trend terms. For all models p is the number of lagged terms in that u_t is white noise and the ADF test for a unit root has the null hypothesis so that $\delta = 0$.

In order to overcome the problem of serial correlation in the error term, Phillips and Perron (1988) developed a non parametric test with the following specification:

$$y_t = c_0 + \rho y_{t-1} + + u_t \tag{4}$$

We also employed the Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) test. The KPSS test (1992) differs from the other unit root tests in that the series y_t is assumed to be stationary under the null. The KPSS test is based on the residuals from the OLS regression of y_t on the exogenous variables x_t, that is:

$$y_t = x_t^{'} \delta + u_t \tag{5}$$

The LM statistic of the KPSS test is defined as $LM = \sum_t = S(t)^2 / (T^2 f_0)$ where f_0 is an estimator of the residual spectrum at frequency zero and *S(t)* is a cumulative residual function. The null hypothesis of the KPSS test is that the process is stationary.

Furthermore to test for the independence of successive price changes we employed either autocorrelation and runs tests, the univariate variance ratio Lo and MacKinlay (1988) as well as the multiple variance ratio test (Chow and Denning, 1993).

The autocorrelation function test (ACF) is a statistical tool that can be used to detect the dependence of successive terms in a given time. This test is often used in order to measure the relationship between the stock return at the current period and its value in the previous period. The specification of the autocorrelation test is the following:

$$\rho_k = \frac{\sum_{t=1-k}^{m}(r_t - \bar{r})(r_{t-k} - \bar{r})}{\sum_{t=1}^{m}(r_t - \bar{r})^2} \tag{6}$$

Where ρ_k is the serial correlation coefficient of returns of lag *k*, *m* is the number of observations, r_t is stock return at time *t*, r_{t-k} is the stock return over period *t-k*; $\bar{r}$ is the sample mean of stock returns, and *k* is the lag of the period. If the stock index returns show a random walk, this means that returns are uncorrelated. To test the joint hypothesis that all serial coefficients ρ_k are simultaneously equal to zero we also applied the Ljung-Box Q-statistics and their p-values. This statistic at lag *k* is a test for the null hypothesis that there is no autocorrelation up to order *k* and is computed as follows:

$$Q_{LB} = m(m+2)\sum_{i=1}^{k}\frac{\delta_j^2}{m-j} \tag{7}$$

Where δ_j is the *j*-th autocorrelation and *m* is the number of observations. We used this test in order to find out whether the serial correlation coefficients are significantly different from zero.

The runs test determines whether successive prices changes are independent. A run is a sequence of successive price changes with the same sign. If the returns series exhibit a greater tendency of change in one direction, the average run will be longer and the number of runs fewer than that generated by a random process. To assign equal weight to each change and to consider only the direction of consecutive changes, each change in return can be classified as positive (+), negative (-), or no change (0). The runs test can also be designed to count the direction of change from any base; for instance, a positive change could be one in which the return is greater than the sample mean, a negative change one in which the return is less than the mean, and zero change representing a change equal to the mean. The actual runs (R) are then counted and compared to the expected number of runs (m) under the assumption of independence as given in the following equation

$$m = \frac{\left[N(N+1) - \sum_{i=1}^{3} n_i^2\right]}{N} \tag{8}$$

where *N* is the total number of observations (price changes or returns) and n_i is the number of price changes (returns) in each category. For a large number of observations $(N > 30)$, the sampling distribution of *m* is approximately normal and the standard error (σ_m) is given by:

$$\sigma_m = \sqrt{\frac{\sum_{i=1}^{3} n_i^2 \left[\sum_{i=1}^{3} n_i^2 + N(N+1)\right] - 2N\sum_{i=1}^{3} n_{i=1}^3 - N^3}{N^2(N-1)}} \tag{9}$$

After computing σ_m, we can obtain the standard normal Z-statistic as follows:

$$Z = (R \pm 0.5 - m)/\sigma_m \tag{10}$$

where *R*, *m* and σ_m are defined as above, while 0.5 is the continuity adjustment in which the sign of the continuity adjustment is negative if $R \geq m$, and positive otherwise. Equation (10) is used to test whether the actual number of runs is consistent with the hypothesis of independence. When actual number of runs exceed (fall below) the expected runs, a positive (negative) *Z* value is obtained. A positive (negative) *Z* value indicates negative (positive) serial correlation in the returns (Abraham et al., 2002)

We further investigated the independence hypothesis by using the Variance Ratio (VR) test (Lo and MacKinlay, 1988). If the index price P_t follows a random walk, then the ratio of the *q-th* difference scaled by *q* to the variance of the first difference tends to equal one, that is:

$$VR(q) = \frac{\sigma^2(q)}{\sigma^2(1)} \tag{11}$$

where $\sigma^2(q)$ is the unbiased estimator of 1/*q* of the variance of the *qth* differences and $\sigma^2(1)$ is the variance of the first differences. Under the null hypothesis VR(q) should be equal to 1. Lo and MacKinlay (1988) produced two statistics. Under the null hypothesis of homoskedasticity, the first test statistic Z(q) is expressed as follows:

$$Z(q) = \frac{VR(q)-1}{\sqrt{v(q)}} \sim N(0,1) \tag{12}$$

where $v(q) = [2(2q-1)(q-1)]/3q(nq)$. While under the null hypothesis of heteroskedasticity, the second test statistic Z*(q) is computed as :

$$Z^*(q) = \frac{VR(q)-1}{\sqrt{v^*(q)}} \sim N(0,1) \tag{13}$$

where $v^*(q) = \sum_{k=1}^{q-1}\left[\frac{2(q-k)}{q}\right]^2 \phi(k)$ and $\phi(k) = \frac{\sum_{t=k+1}^{nq}(x_t - x_{t-1} - \hat{\mu})^2 (x_{t-k} - x_{t-k-1} - \hat{\mu})^2}{[\sum_{t=1}^{nq}(x_t - x_{t-1} - \hat{\mu})^2]^2}$

both the Z(q) and Z*(q) statistics test the null hypothesis that VR(q) approaches one. When the random walk hypothesis is rejected and VR(q)>1, returns are positively serially correlated. When the random walk hypothesis is rejected and VR(q)<1, returns are negatively serially correlated.

As pointed out by Lagoarde-Segot and Lucey (2008) the choice of block length *q* represents one limit to this approach, at the same time the RWH requires that the variance ratios for each block length selected should be equal to one. In order to overcome these drawbacks Chow and Denning (1993) proposed a multiple variance ratio (MVR) test which was based on the Lo and MacKinlay (1988) single variance ratio (VR) test. The Lo and MacKinlay (1988) procedure is implemented in order to test individual variance ratios for a specific aggregation interval, *q*, but the random walk hypothesis requires that VR(q)=1 for all aggregation intervals. In the Chow and Denning's MVR a set of variance ratios is tested against one, that is the null hypothesis $V(q_i)=1$ for i=1,...,n is tested against the alternative that $V(q_i)\neq 1$ for some *i*. Chow and Denning's test statistic is $MV_1=\sqrt{T}\max_{1\le i\le n}|Z(q_i)|$ where $Z(q_i)$ is defined in equation (7). The null hypothesis is rejected at the α level of significance if the MV_1 is greater than the $(1-[\alpha*/2])^{th}$ percentile of the standard normal distribution, where $\alpha*=1-(1-\alpha)^{1/n}$. The heteroskedasticity-robust version can be written as $MV_2=\sqrt{T}\max_{1\le i\le n}|Z*(q_i)|$, where $Z*(q_i)$ is defined in equation (12), and it has the same critical values as MV_1. The Chow and Denning test is based on the following inequality:

$$PR\left\{\max\left(|Z(q_i)|,\ldots,|Z(q_m)|\right)\le SMM(\alpha;m;T)\ge 1-\alpha\right\} \tag{14}$$

where $SMM(\alpha;m;T)\ge 1-\alpha$ is the upper α point of the Standardized Maximum Modulus (SMM) distribution with parameters *m* (number of aggregation intervals) and *T* (sample size) degrees of freedom. Chow and Denning (1993) controlled the size of the MV ratio test by comparing the calculated values of the standardized test statistics, either *Z(q)* or *Z*(q)* with the SMM critical values. If the maximum absolute value of, say *Z(q)* is greater then the SMM critical value than the random walk hypothesis is rejected. Following Chow and Denning (1993), we used the SMM distribution, which has a critical value of 2.491 for the 5 percent level of significance, to test the RWH.

One criticism of the Chow and Denning (1993) multiple variance-ratios test is that their critical values are large by design. In order to overcome this criticism, Geweke and Porter-Hudak (1983) (GPH hereafter) proposed a procedure that can be used to test for random walk. the hypothesis of pure random walk is given by the following equation:

$$P_t = c + P_{t-1} + u_t \tag{15}$$

where P_t is the log of the equity prices series, c is a constant and u is a random error term. The above equation can be rewritten as

$$(I - B)^d R_t = e_t \tag{16}$$

where R_t is a first-differenced stationary stock price series, B is the backshift operator, and d is a fractional integration parameter, e_t is a stationary process, and $(I - B)^d$ is called the fractional integration operator. The parameter d is usually restricted to integer values in the classical time series models, GPH (1983) relaxed that restriction and allow for fractional values of d (fractional integration). Therefore, GPH provided a comparison for the multiple variance-ratio test, with an additional capacity to capture the behaviour of stock prices.

Wright (2000) proposed the use of signs and ranks of differences in place of the differences in the Lo and MacKinlay tests. Wright demonstrated that this nonparametric variance ratio tests based on ranks (R1 and R2) and signs (S1) can be more powerful than the tests suggested by Lo and MacKinlay and are more appropriate when the distribution of returns is not normal.

The test statistics based on ranks (R1 and R2) were computed as follows

$$R_1(k) = \left(\frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(r_{1t} + \ldots + r_{1t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} r_{1t}^2} - 1 \right) \times \phi(k)^{-1/2} \tag{17}$$

and

$$R_2(k) = \left(\frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(r_{2t} + \ldots + r_{2t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} r_{2t}^2} - 1 \right) \times \phi(k)^{-1/2} \tag{18}$$

Where $r_{1t} = \left(r(y_t) - \frac{T+1}{2} \right) / \sqrt{\frac{(T-1)(T+1)}{12}}$ and $r_{2t} = \Phi^{-1}\left(r(y_t)/(T+1)\right)$, T are observations of first differences of a variable, $\{y_1, \ldots, y_T\}$, ϕ_t is the asymptotic variance, $r(y_t)$ is the rank of y_t among $y_1, \ldots, y_T$, and Φ^{-1} is the inverse of the standard normal cumulative distribution function.

The test based on the signs of first differences is given by:

$$S_1(k) = \left(\frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(s_t + ... + s_{t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} s_t^2} - 1 \right) \times \phi(k)^{-1/2} \tag{19}$$

where ϕ_t is the asymptotic variance, $s_t = 2u(y_t, 0)$, $s_t(\overline{u}) = 2u(y_t, \overline{\mu})$ and

$$u(x_t, q) = \begin{cases} 0.5 & if\ x_t > 0 \\ -0.5 & otherwise \end{cases} \tag{20}$$

In this work, the day of the week effect is studied through a Generalized Autoregressive Conditional Heteroskedasticity (GARCH) framework introduced by Bollerslev (1986). The GARCH model provides a flexible framework in order to capture various dynamic structures of conditional variance and it allows simultaneous estimation of several parameters of interest and hypothesis. An important restriction of the GARCH specification is its asymmetry. That is, big negative shocks have the same impact on future volatility as big positive shocks of the same magnitude. An interesting extension is towards asymmetric volatility models, in which good news and bad news have a different impact on future volatility. An asymmetric model allows for the possibility that an unexpected drop in price (bad news) has a larger impact on future volatility than an unexpected increase in price (good news) of similar magnitude. A fruitful approach to capture such asymmetries is provided by Nelson's (1990) exponential GARCH (EGARCH) model. The EGARCH model is given by two equations, that is the mean and the volatility returns equation. Following Karolyi (1995) and Kiymaz and Berument (2003), we modeled the conditional volatility of stock returns by incorporating the day of the week effect into both equations. Our model is given by the following equations:

$$y_t = c + \lambda_1 D_{1t} + \lambda_2 D_{2t} + \lambda_3 D_{3t} + \lambda_4 D_{4t} + \sum_{i=1}^{n} a_i R_{t-i} + \varepsilon_t \tag{21}$$

$$\log\left(\sigma_t^2\right) = c + \delta_1 D_1 + \delta_2 D_2 + \delta_3 D_3 + \delta_4 D_4 + \sum_{j=1}^{q} \beta_j \log\left(\sigma_{t-j}^2\right) + \sum_{i=1}^{p} \alpha_i \left| \frac{\varepsilon_{t-i}}{\sigma_{t-i}} \right| + \sum_{k=1}^{r} \gamma_k \frac{\varepsilon_{t-k}}{\sigma_{t-k}} \tag{22}$$

where R_t represents returns on a selected index, ε_t is an error term, σ_t^2 is the conditional variance. The EGARCH model is asymmetric as long as $\gamma \neq 0$, when $\gamma < 0$, positive shocks generate less volatility than negative shocks. The dummy variables in the mean and variance equations (that is equations 21 and 22) represent four trading days of the week. In other words, D_t is equal to one if the day t is a Monday, and otherwise is zero. We included also Tuesday, Thursday and Friday, while we avoided including also the Wednesday dummy variable in order to avoid the dummy variable trap. The parameters $\lambda_1, \lambda_2, \lambda_3, \lambda_4$ represent the Monday effect, the Tuesday effect, the Thursday effect, and the Friday effect on stock returns respectively. Given the general empirical findings of papers investigating effects of weekdays and weekends on stock markets, the expected sign on the coefficient on the Monday dummy should be negative and significantly different from zero. Some studies indicate that the coefficient on the Friday dummy in equation 21 should be positive (Keim and Stambaugh, 1984; Agrawal and Tandon, 1994). Similarly, based on French and Roll (1986) and Foster and Viswanathan (1990), the expected sign of the significant Monday coefficient in the volatility equation should be positive, and negative for the Friday effect.

In order to check whether our result changed if we added a risk premium variable to our model, we also estimated an Exponential Generalized Autoregressive Condition Heteroskedasticity in mean (M-EGARCH) model: this model which allowed us to incorporate also a risk premium variable. Therefore our second model is the M-EGARCH specification of the following form:

$$y_t = c + \lambda_1 D_{1t} + \lambda_2 D_{2t} + \lambda_3 D_{3t} + \lambda_4 D_{4t} + \sum_{i=1}^{n} a_i R_{t-i} + \phi h_t + \varepsilon_t \quad (23)$$

$$\log\left(\sigma_t^2\right) = c + \delta_1 D_1 + \delta_2 D_2 + \delta_3 D_3 + \delta_4 D_4 + \sum_{j=1}^{q} \beta_j \log\left(\sigma_{t-j}^2\right) + \sum_{i=1}^{p} \alpha_i \left|\frac{\varepsilon_{t-i}}{\sigma_{t-i}}\right| + \sum_{k=1}^{r} \gamma_k \frac{\varepsilon_{t-k}}{\sigma_{t-k}} \quad (24)$$

where ϕ is a measure of the risk premium, as it is possible that the conditional variance, as proxy for risk, can affect market returns. If ϕ is positive, then the risk averse agents must be compensated to accept higher risk.

Another calendar anomaly explored was the so called 'January effect'. This calendar anomaly is characterized by higher stock returns in January than in any other month of the year. Agraswal and Tandon (1994) found that these effects characterize most of the developed stock markets. We wanted to detect whether the January effect also characterizes African stock markets. In order to analyse this issue we followed Coutts et al. (2000) methodology by estimating the following equation by OLS, that is:

$$R_t = \beta_1 + \sum_{k=2}^{12} \beta_k M_{kt} + u_t \quad (25)$$

where R_t is the i-th stock index returns for day t, M_{kt} is a monthly dummy variable (such that M_{2t}=February, M_{3t}=March, M_{4t}=April, M_{5t}=May, M_{6t}=June, M_{7t}=July, M_{8t}=August, M_{9t}=September, M_{10t}=October, M_{11t}=November, M_{12t}=December) and u_t is the disturbance term. The coefficients β_{1t} measure the mean return for January, whereas all remaining coefficients represent the average differences in returns between the month of January and each individual month of the year. If the January effect is present we would expect to find significantly positive mean returns for January, and that positive January return is higher than the return for any other month.

3. Data

The data consist of daily index values for Egypt, Morocco, Nigeria, South Africa, and Tunisia from 4th January 2000 to 26th March 2009. The stock price indices are expressed in local currencies and were extracted from *Thomson Financial Datastream* (see table 1).

Table 1 - Stock index prices (local currency), 1999-2009

Country	Index	Datastream code
Egypt	FTSE W Egypt	WIEGYTL(PI)
Morocco	FTSE W Morocco	WIMORCL(PI)
South Africa	FTSE/JSE All Share	JSEOVER(PI)
Tunisa	Tunindex	TUTUNIN(PI)

Table 1 presents descriptive statistics of the daily returns for the four African markets. The lowest mean returns are in Morocco while the highest mean returns are for Egypt. The standard deviations of returns range from 0.0053 (Tunisia) to 0.017 (Egypt). On this basis, of the four markets the returns for Tunisia and Morocco are the least volatile, with South Africa and Egypt having the most volatile.

Table 2– Summary statistics for daily returns

Index	N obs	Mean	Minimum	Maximum	Std Dev	Skewness	Kurtosis	Jarque-Bera Test	p-value
Egypt	2407	0.000436	-0.196	0.085	0.017	-0.826	15.541	16048.54	0.00
Morocco	2407	0.000312	-0.055	0.073	0.01	0.242	9.060	3706.953	0.00
S.Africa	2407	0.000386	-0.0789	0.0683	0.0134	-0.169	6.729	1406.477	0.00
Tunisia	2407	0.000392	-0.05	0.0461	0.005	0.217	14.720	13796.64	0.00

Notes: The Jarque-Bera statistic tests the null hypothesis of a normal distribution and is distributed as a χ^2 with 2 degrees of freedom.

Fig.1 shows African indices performance during the period considered. It can be seen from the figure that the stock indices saw a slow growth until the first half of the period considered, increased sharply between September 2005 and August 2008, declined sharply from September 2008. Fig.2 showed that the daily returns[3] highly fluctuated between September 2008 and February 2009. For the time period under study all markets experienced positive returns (fig.2), on the other hand we may also see strong non normalities in the unconditional distributions of the returns with

[3] Daily returns are computed as $R_t = \ln(P_t/P_{t-1})$, where P_t is the price of stock index at instant *t*.

either positive and negative skewness and heavy tails for all returns, deviations from normality can be seen also from the reported Jarque-Bera statistic. This leptokurtic behaviour of the returns is clearly shown by the normal quantile graphs presented in fig.3.

Figure 1 – Daily prices for African stock indices

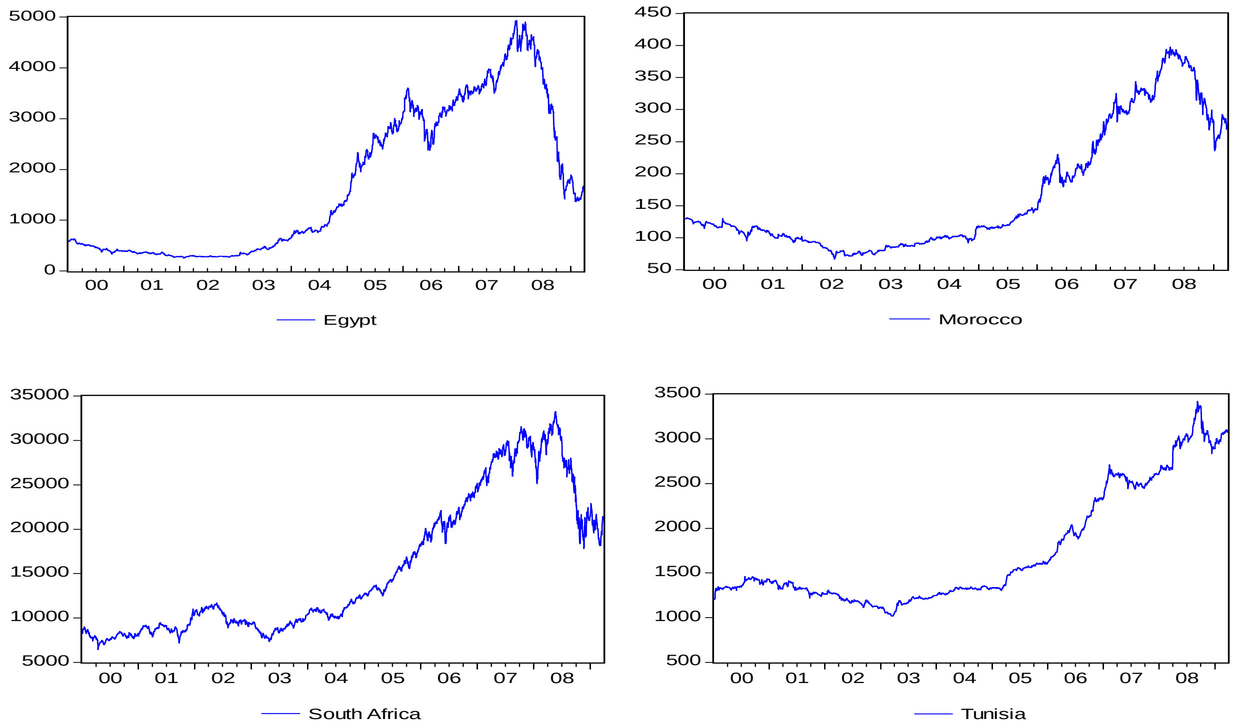


Figure 2 –African stock index returns

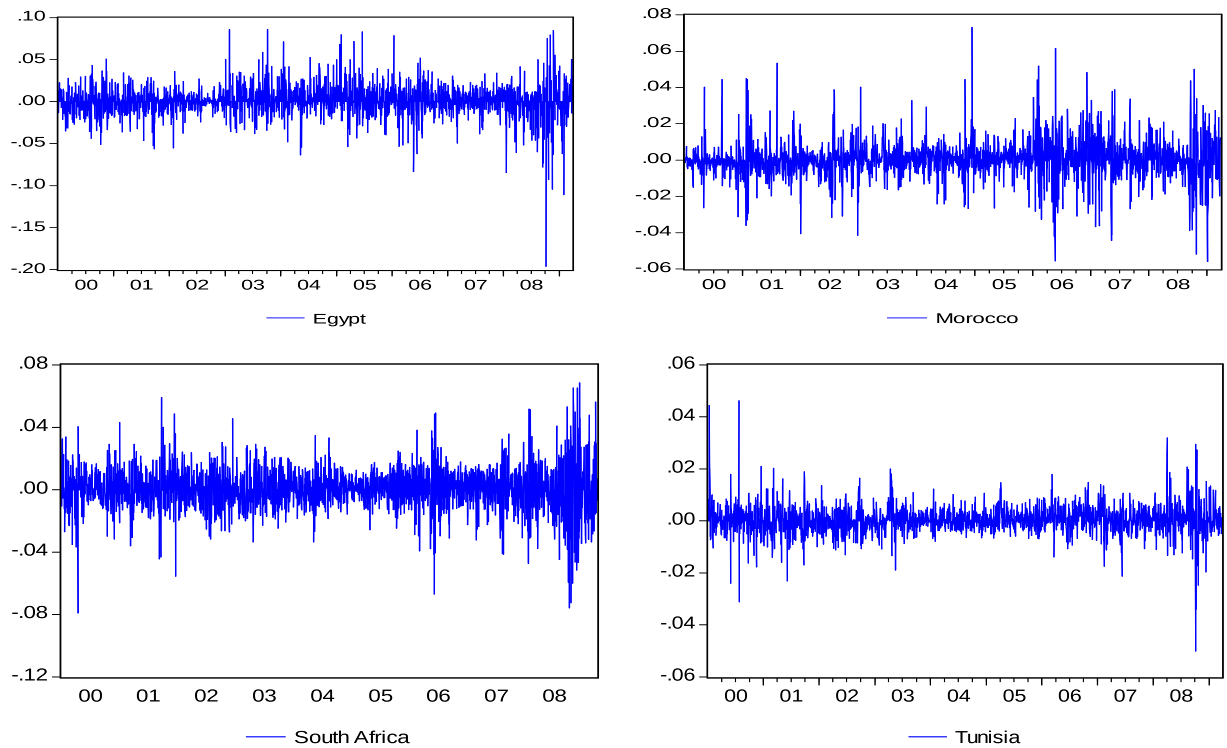
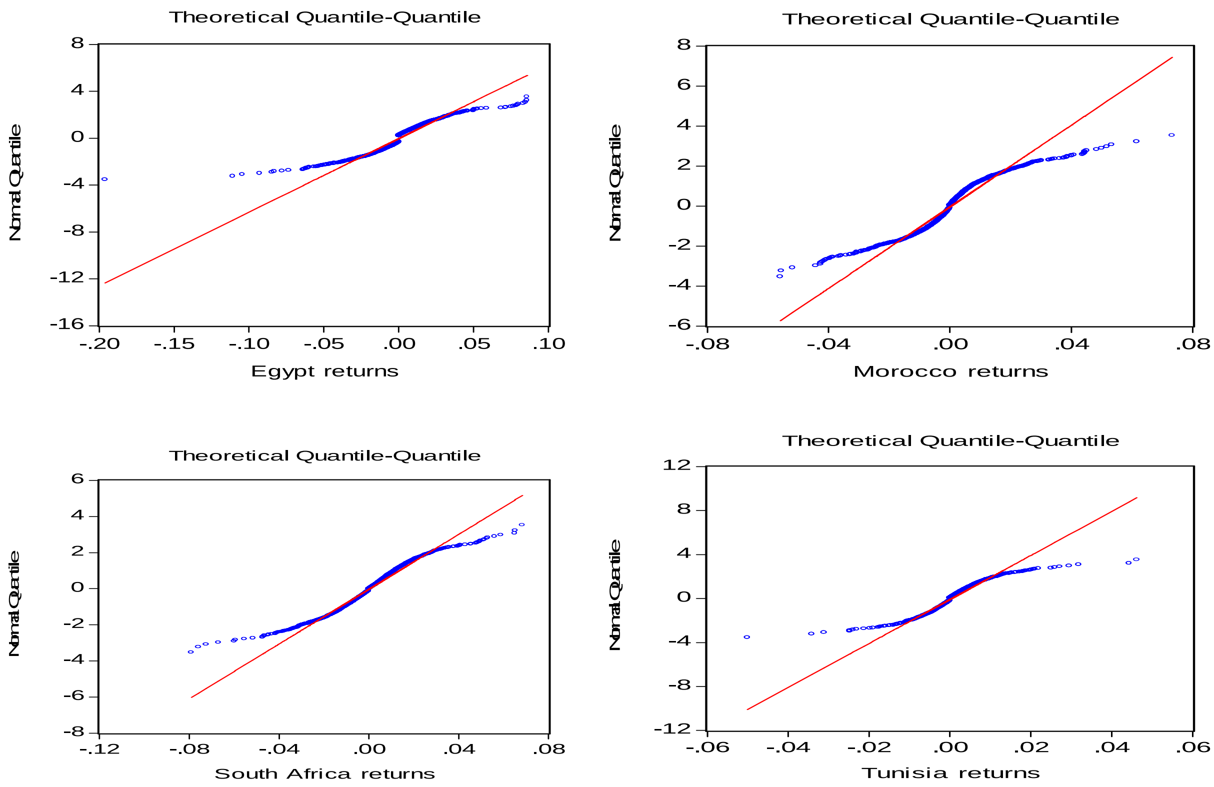


Figure 3 – Q-Q normal plot returns normal distribution of each index



4. Empirical Results

When we find evidence of a unit root in the time series then there is evidence for the RWH, implying market efficiency: in order to verify that hypotheses several unit root tests were carried out. Table 3 illustrates the unit root tests results. At levels, the ADF and PP t-statistics do not reject the null hypothesis of a unit root at the 5% level of significance, thus indicating that all of the price series in log form are non-stationary. We also used the KPSS procedure of Kwiatkowski et al (1992) which has the advantage of being specifically designed to test the null hypothesis of stationarity[4]. For the KPSS tests, the LM-statistics exceed the asymptotic critical values at the 1% level for all markets at the level series, indicating these series are non-stationary. Since the ADF, PP and KPSS tests on the log of prices do not to reject the presence of unit roots, there is no evidence against weak form efficiency for all markets.

[4] The null hypothesis of stationarity is rejected in favour of the unit root alternative if the calculated test statistic exceeds the critical values estimated in Kwiatkowski et al. (1992).

Table 3 - Unit root tests for African stock markets returns

	ADF	PP	KPSS
	Log levels		
Egypt	-0.551 (0.875)	-0.555 (0.877)	5.190
Morocco	0.043 (0.961)	0.1 (0.965)	4.546
South Africa	-0.858 (0.801)	-0.810 (0.815)	5.366
Tunisia	0.916 (0.995)	0.906 (0.995)	4.791
	First log difference		
Egypt	-31.431** (0.00)	-47.398** (0.00)	0.658**
Morocco	-39.978** (0.00)	-39.937** (0.00)	0.602**
South Africa	-46.119** (0.00)	-46.036** (0.00)	0.145
Tunisia	-39.290** (0.00)	-39.415** (0.00)	0.540**

Notes: Test equations for all cases include a constant. The critical value for the ADF and PP tests with intercept are: -3.43(1%); -2.86(5%) and -2.56(10%) while for the KPSS test are: 0.739(1%), 0.463(5%) and 0.347(10%) For ADF test and PP test hypotheses are: H_o: unit root (non stationary), H_1: no unit root (stationary). For The KPSS test hypotheses are H_o: no unit root (stationary), H_1: unit root (non stationary). The asymptotic critical values for the KPSS LM test statistic are are 0.739 (1%), 0.463(5%) and 0.347(10%) for the test including a constant using data on stock returns.

To test RWH for the African stock markets further, autocorrelation tests up to 24 lags were performed for daily stock returns. Results (tab.4) show that the null hypothesis of random walk is rejected for all series: it is worth noting that the positive sign of the autocorrelation coefficients indicates that consecutive daily returns tend to have the same sign, so that positive (negative) return in the current day tends to be followed by an increase (decrease) of return in the next few days.

Table 4 – Autocorrelation tests with *p* lags for African stock markets returns

			Egypt		
p	1	4	8	12	24
ACF	0.041	0.048	-0.022	0.028	-0.005
Q-stat	3.960	(27.947)	(38.581)	(42.215)	(69.490)
p-value	(0.047)	(0.00)	(0.00)	(0.00)	(0.00)
			Morocco		
P	1	4	8	12	24
ACF	0.201	-0.021	0.011	-0.003	0.018
Q-stat	96.887	106.20	107.83	109.14	113.16
p-value	0.00	0.00	0.00	0.00	0.00
			South Africa		
p	1	4	8	12	24
ACF	0.061	-0.005	0.018	-0.025	-0.029
Q-stat	9.057	20.053	29.722	35.510	46.349
p-value	(0.003)	(0.00)	(0.00)	(0.00)	(0.00)
			Tunisia		
p	1	4	8	12	24
ACF	0.217	0.001	0.046	0.039	0.018
Q-stat	110.55	132.49	139.65	159.82	187.71
p-value	0.00	0.00	0.00	0.00	0.00

As pointed out by Abraham et al. (2002) the non-parametric runs test is considered more appropriate than the parametric autocorrelation test since observed returns do not follow the normal distribution. Results of the runs test (tab.5) indicate that the null hypothesis of independence among stock returns is rejected for all stock markets returns with the exception of South Africa.

Table 5 - Runs test for the African stock markets

	Obs	Actual Runs (R)	Expected runs (m)	Z-statistic
Egypt	2407	1454	1068.64	13.083**
Morocco	2407	1050	1198.84	-6.077**
South Africa	2407	1164	1204.04	-1.572
Tunisia	2407	2407	1187.86	-4.604**

Notes. If the Z-statistic is greater than or equal to ± 1.96, then we reject the null hypothesis at 5% level of significance. **Indicates rejection of the null hypothesis that successive price changes are independent.

Table 6 reports the variance ratio estimates and test statistics of random walk hypothesis based on the methodology described in section 2. The variance ratio was computed for multiples of 2, 4, 8, 16 and 32 days, with the one-day return used as a base. Results indicate that almost all of the test statistics for either assuming homoskedasticity or heteroskedasticity-consistent at any number of *q* are significant: this means that stock markets returns show predictability but not South Africa. In other words the random walk hypothesis is rejected for all stock markets with the exception of South Africa.

Table 6 - Variance Ratio estimates VR(q) and variance-ratio test statistics Z(q) and Z*(q) for a one-day base observation period

Returns	2	4	8	16	32
Egypt					
VR(q)	1.041	1.162	1.283	1.325	1.539
Z(q)	(2.033)**	(4.272)**	(4.702)**	(3.624)**	(4.147)**
Z*(q)	[1.487]	[3.208] **	[3.531]**	[2.696]**	[3.057]**
Morocco					
VR(q)	1.2016	1.3601	1.401	1.439	1.528
Z(q)	(9.8921) **	(9.4431) **	(6.6614) **	(4.895)**	(4.061)**
Z*(q)	[5.529] **	[5.4115] **	[4.0389] **	[3.203]**	[2.893]**
South Africa					
VR(q)	1.061	1.08	1.012	0.951	0.966
Z(q)	(3.044)**	(2.086) **	(0.195)	(-0.545)	(-0.256)
Z*(q)	[2.077] **	[1.393]	[0.129]	[-0.361]	[-0.173]
Tunisia					
VR(q)	1.218	1.423	1.503	1.675	1.877
Z(q)	(10.740) **	(11.114) **	(8.355) **	(7.526)**	(6.744)**
Z*(q)	[4.729] **	[4.998] **	[4.215] **	[4.375]**	[4.483]**

Notes. Under the random walk null hypothesis, the value of the variance ratio test is 1 and the test statistic has a standard normal distribution (asymptotically). Test statistics marked with two asterisks indicate that the corresponding variance ratios are statistically different from 1 at the 5% level of significance.

Turning to the Chow and Denning (1993) tests results are reported in tab. 7. We can see that, at the 5% level of significance, the Chow and Denning's (1993) homoskedastic and heteroskedastic nulls reject the random walk hypothesis for all stock markets, given that the maximum value is greater than the 2.49 critical value. These results confirm previous tests' results whereas they are contradictory relative to the South African stock market.

Table 7 – Multiple variance ratio tests

	Egypt	Morocco	South Africa	Tunisia
MV_1	49.113	49.100	49.088	49.034
MV_2	36.233	30.518	33.852	31.264

Note. MV_1 is the homoskedastic and MV_2 is the heteroskedastic-robust version of the Chow-Denning test. ** reject the null hypothesis at the 5% level of significance

Given the improved power properties of Wright's (2000) test, we used this last test in order to check robustness of the Chow and Denning tests results. The ranks and signs based variance ratio statistics test based on Wright's methodology for the entire period is summarized in tab. 8. The rank-based test results show that R_1 and R_2 are significant for all countries with the exception of South Africa for all numbers of *k* above 5. Overall RWH cannot be rejected by ranks and signs based variance ratio tests relative to the South African stock returns for k=10 and k=30.

Table 8 – Wright Non-Parametric Variance Ratio Tests using ranks and Signs

	Number of lags (k)			
	k=2	k=5	k=10	k=30
Egypt				
R_1	1.82*	3.66**	2.72**	1.82*
R_2	1.88*	4.18**	3.44**	2.39**
S_1	1.22	2.21**	3.05**	6.51**
Morocco				
R_1	10.70**	10.77**	9.16**	9.20**
R_2	10.87**	10.28**	8.06**	7.13**
S_1	7.79**	8.51**	7.96**	9.28**
South Africa				
R_1	3.52**	2.24**	0.98	0.84
R_2	3.41**	1.87*	0.41	-0.12
S_1	1.92*	1.06	1.05	1.99**
Tunisia				
R_1	11.33**	11.34**	8.88**	7.99**
R_2	12.15**	11.80**	8.98**	8.04**
S_1	6.69**	7.33**	5.76**	5.10**

*** significant at the 1% level; ** significant at the 5% level, significant at the 10% level.

GPH test results are given in tab.9. We report that the *d* estimate indicates there does not appear to be any consistent convincing evidence supporting the long-memory (biased random-walk) hypothesis for the returns series of any stock indices with the exception of the Tunisian stock market.

Table 9 – GPH fractional integration test of random walk hypothesis for African stock markets indices

	Number of Observations	d(0.50)	d(0.60)	d(0.70)
Egypt	2407	0.103 (0.091)	0.042 (0.059)	0.057 (0.042)
Morocco	2407	0.0153 (0.093)	-0.009 (0.065)	0.042 (0.047)
South Africa	2407	0.023 (0.075)	0.035 (0.057)	0.028 (0.041)
Tunisia	2407	-0.062 (0.097)	0.150** (0.068)	0.031 (0.043)

Notes. d(0.50), d(0.60), d(0.70) give the d estimates corresponding to the spectral regression of sample size $v=T^{0.50}$, $v=T^{0.60}$, $v=T^{0.70}$. respectively. OLS standard error among parenthesis. ** indicates statistical significance for the null hypothesis d=0 at the 5% level. OLS standard error among parenthesis.

Next, we examined the day of the week effect on stock returns and volatility. Panel A of table 10 reports empirical results of the day of the week effects analysis. The coefficient of Monday's dummy variables for the South African index (0.0016) is positive and statistically significant at the 1% level. Friday's returns are positive and significant only for the Tunisian stock market. The estimated coefficient for the Tunisian index (-0.0006) is the lowest and statistically significant at 1% level on Tuesdays. In panel B we also report the estimates of the volatility equation. The day of the week effect is observed relative to Egyptian returns on Mondays and Fridays, while on Tuesdays and Thursdays we observe the day of the week effect on Moroccan returns. Finally we note that conditional volatility equations show a negative and significant value of the γ coefficient only for South African returns, indicating the existence of an asymmetric effect in returns during the sample period. Panel C of table 10 reports the Ljung-Box (Q) statistics for the residuals and Engle's (1982) ARCH-LM test at 4-, 8-, and 20-day lags. From the Q statistics we cannot reject the null hypothesis that residuals are not autocorrelated. Furthermore, there is no significant ARCH effect in any of the EGARCH models estimated.

Table 10 – The day of the week effect in EGARCH (1,1) models

Panel A – Estimates of the mean equation				
Index	Egypt	Morocco	South Africa	Tunisia
Constant	0.0002	-0.0001	-0.0002	0.0002
	(0.0005)	(-0.509)	(0.0004)	(0.0001)
Monday	-0.0002	-0.0003	0.0016***	-0.0002
	(0.0008)	(0.0003)	(0.0006)	(0.0002)
Tuesday	-0.0001	0.0003	0.0002	-0.0006***
	(0.0007)	(0.0003)	(0.0006)	(0.0002)
Thursday	0.002**	0.0004	0.0014**	9.47E-05
	(0.0006)	(0.0003)	(0.0006)	(0.0002)
Friday	-0.0002	0.0002	0.0007	0.0004*
	(0.0005)	(0.0003)	(0.0006)	(0.0002)
$Return_{t-1}$	7.34E-07	0.189**	0.063***	0.203***
	(3.42E-05)	(0.019)	(0.021)	(0.02)
Panel B – Estimates of the volatility equation				
constant	-0.665***	-1.357***	-0.324***	-1.524***
	(0.175)	(0.179)	(0.111)	(0.260)
α	0.0009	0.582***	0.132***	0.398***
	(0.008)	(0.064)	(0.018)	(0.037)
β	0.856***	0.914***	0.982***	0.876***
	(0.042)	(0.013)	(0.0045)	(0.021)
γ	0.004	-0.031	-0.085***	-0.026
	(0.005)	(0.028)	(0.011)	(0.023)
Monday	15.436***	0.187	0.136	-0.072
	(0.753)	(0.150)	(0.329)	(0.142)
Tuesday	-0.103	0.313*	0.058	-0.133
	(0.235)	(0.182)	(0.338)	(0.178)
Thursday	-0.173	0.426**	0.230	-0.143
	(0.231)	(0.184)	(0.172)	(0.179)
Friday	-16.333**	0.107	-0.125	-0.110
	(0.179)	(0.153)	(0.142)	(0.148)

Panel C– Autocorrelation Q statistics and ARCH-LM tests for various lags						
	Q(4)	ARCH(4)	Q(8)	ARCH(8)	Q(12)	ARCH(12)
Egypt	0.0045	0.006	44.617	1.901	58.738	1.379
	[0.998]	[0.999]	[0.00]	[0.055]	[0.00]	[0.167]
Morocco	15.601	0.671	17.061	0.705	20.630	0.595
	[0.001]	[0.611]	[0.017]	[0.687]	[0.037]	[0.847]
South Africa	8.195	0.541	12.500	2.209	14.271	1.643
	[0.042]	(0.705)	[0.085]	[0.024]	[0.218]	[0.073]
Tunisia	13.674	0.201	17.528	0.278	24.576	0.285
	[0.003]	[0.937]	[0.014]	[0.973]	[0.011]	[0.991]

Notes. */**/*** indicate statistical significance at 10%, 5%, and 1% levels. Standard errors are among parentheses, while p-values are among brackets.

Further we employed the M-EGARCH models to determine whether the day of the week effect estimated in the previous EGARCH models, changed by adding an equity risk variable (ϕ). The estimated results are summarized in table 11. Note that if the dummy variables for each day of the trading week are still significant in the mean equation of the M-EGARCH model, it may be concluded that the day of the week effect is not due to the variation in the equity risk. Following this principle, table 11 reveals that equity risk is negative and statistically significant for the

Tunisian stock market returns. In other words the risk premium has a negative impact on returns of the Tunisian equity market. This last result shows that there is a trade-off between return and risk in that market. We also note that mean and volatility equation showed results quite similar to the EGARCH models estimated previously.

Table 11 – The day of the week effect in M-EGARCH (1,1) models

Panel A – Estimates of the mean equation				
Index	Egypt	Morocco	South Africa	Tunisia
Constant	0.0002	-0.0005	0.0001	-0.0005
	(0.003)	(0.0003)	(0.0007)	(0.0003)
Monday	9.02E-05	-0.0003	0.001***	-0.0002
	(0.0016)	(0.0003)	(0.0006)	(0.0002)
Tuesday	0.0001	0.0002	0.0002	-0.0006***
	(0.0009)	(0.0003)	(0.0006)	(0.0002)
Thursday	0.002***	0.0004	0.0014**	0.0001
	(0.0009)	(0.0003)	(0.0006)	(0.0002)
Friday	-0.0002	0.0002	0.0007	0.0004**
	(0.003)	(0.0003)	(0.0006)	(0.0002)
$Return_{t-1}$	-1.57E-06	0.185***	0.062***	0.2***
	(2.98E-05)	(0.019)	(0.021)	(0.021)
ϕ	-0.0016	0.055	-0.043	-0.192**
	(0.01)	(0.036)	(0.064)	(0.07)
Panel B – Estimates of the volatility equation				
constant	-0.633***	-1.375***	-0.320***	-1.618***
	(0.241)	(0.183)	(0.111)	(0.273)
α	0.006	0.602***	0.132***	0.4***
	(0.007)	(0.06)	(0.018)	(0.03)
β	0.838***	0.911***	0.983***	0.867***
	(0.058)	(0.014)	(0.004)	(0.02)
γ	0.002	-0.025	-0.085***	-0.021
	(0.005)	(0.029)	(0.011)	(0.023)
Monday	15.499***	0.180	0.148	-0.056
	(1.061)	(0.150)	(0.138)	(0.142)
Tuesday	-0.086	0.299	0.067	-0.138
	(0.230)	(0.182)	(0.173)	(0.178)
Thursday	-0.304	0.392**	0.249	-0.135
	(0.225)	(0.184)	(0.171)	(0.178)
Friday	-16.693***	0.099	-0.113	-0.108
	(0.180)	(0.153)	(0.142)	(0.145)

Panel C– Autocorrelation Q statistics and ARCH-LM tests for various lags						
	Q(4)	ARCH(4)	Q(8)	ARCH(8)	Q(12)	ARCH(12)
Egypt	0.0046	0.007	50.596	2.514	69.687	1.858
	(0.998)	(0.999)	(0.00)	(0.01)	(0.00)	(0.034)
Morocco	16.556	0.723	18.038	0.728	21.372	0.610
	(0.001)	(0.575)	(0.012)	(0.666)	(0.03)	(0.834)
South Africa	8.139	0.479	12.485	2.170	14.656	1.638
	(0.043)	(0.750)	(0.086)	(0.026)	(0.199)	(0.074)
Tunisia	14.668	0.08	18.617	0.255	26.601	0.277
	(0.002)	(0.986)	(0.009)	(0.979)	(0.005)	(0.992)

Notes. */**/*** indicate statistical significance at 10%, 5%, and 1% level. Standard errors are among parentheses, while p-values are among brackets.

Moving on to check whether there is a January effect, table 12 shows that for the Egyptian stock market, mean returns in January are significantly positive. Further February, March, May, June, October and November have significantly negative returns. The results for the Moroccan stock market returns are almost identical: the only difference with the Egyptian results is that also July and September show negative returns. The Tunisian stock market returns for each month are quite similar to the Egyptian results. In addition, also the December returns are significantly negative. For the South African stock market, we did not find evidence of the January effect. Finally the F-statistics suggest a rejection of the null hypothesis of equal β's for each of the M-GARCH models estimated with the exception of Egyptian and South African models. Overall, our results indicate that seasonality and the January effect are present in Egypt, Morocco and Tunisia, whereas for the South African stock market we did not find either seasonality or the January effect.

Table 12 – Regression analysis for the January effect

	Egypt	Morocco	South Africa	Tunisia
β_1	0.0026**	0.001*	0.0003	0.001***
	(0.001)	(0.0006)	(0.0009)	(0.0003)
β_2	-0.003**	0.0003	-0.0004	-0.001**
	(0.001)	(0.001)	(0.001)	(0.0005)
β_3	-0.002*	-0.0009	-0.0002	-0.0002
	(0.001)	(0.0009)	(0.001)	(0.0005)
β_4	-0.0008	-0.0006	-0.0003	0.0004
	(0.001)	(0.001)	(0.001)	(0.0005)
β_5	-0.003**	-0.002**	0.001	-0.0007
	(0.001)	(0.001)	(0.001)	(0.0005)
β_6	-0.003*	-0.001*	-0.0006	-0.001***
	(0.001)	(0.001)	(0.001)	(0.0005)
β_7	-0.002	-0.002**	-0.001	-0.0007
	(0.001)	(0.001)	(0.001)	(0.0005)
β_8	-0.001	0.0006	0.001	-0.0003
	(0.001)	(0.001)	(0.001)	(0.0005)
β_9	-0.001	-0.002**	0.001	-0.0006
	(0.001)	(0.001)	(0.001)	(0.0005)
β_{10}	-0.003**	-0.002**	-0.0001	-0.001***
	(0.001)	(0.001)	(0.001)	(0.0005)
β_{11}	-0.003*	-0.001	0.0005	-0.001**
	(0.001)	(0.001)	(0.001)	(0.0005)
β_{12}	0.0003	-1.72E-05	0.001	-0.0009*
	(0.001)	(0.001)	(0.001)	(0.0005)
F-statistic	1.359	2.319	0.853	2.276
p-value	0.185	0.007	0.585	0.009

Notes. Standard errors are parentheses. ***/**/* indicate significance at 1%, 5%, and 10% level. Standard errors are among parentheses. *F*-statistics denote test of null hypothesis $\beta_2 = \beta_3 = \beta_4 = \beta_5 = \beta_6 = \beta_7 = \beta_8 = \beta_9 = \beta_{10} = \beta_{11} = \beta_{12}$.

Conclusion

This study examined the random walk hypothesis for the Egypt, Moroccan, South African and Tunisian stock markets. We found that these markets did not follow the random walk hypothesis during the time period considered and therefore they were significantly inefficient with the exception of the South African stock market. The inefficiency of these stock markets imply that the benefits of an efficient stock market are not being realised in these economies. This raises a further consideration. A way for achieving economic development is to raise capital using stock markets. The lack of efficiency of these African financial markets may negatively affect their efforts. Further if stock markets are not efficient and local firms are forced to raise capital locally, then their cost of capital is higher than that of firms with unrestricted access to international capital markets. Further research is necessary to detect causes of stock market inefficiency as well as measures that need to be taken to improve the efficiency of the African stock markets considered here.

Another issue we explored in this study was the day of the week effect. We found the existence of various significant days of the week effects, including the typical negative Monday and Friday positive effects in several stock markets. After adjusting for the equity risks, these effects seem to be present also in M-EGARCH models so estimated. This study also found evidence of asymmetrical markets effect relative to the South African stock markets, whereas similar effects are not present in other markets. One major implication of these findings is that investors in these markets may consider buying shares on Monday and selling them on Friday.

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