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Testing for the Weak-Form Market Efficiency of the Dar es Salaam Stock Exchange

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Abstract

This paper investigates into the weak-form efficiency of the Dar es Salaam stock exchange (DSE), a frontier market, in Tanzania. The study covers the period from January 2007 to December 2014. To establish the consistency and robustness of the obtained conclusions, we employ different tests (i.e., Augmented Dickey-Fuller test, Variance-ratio test, and Ranks and Sign test) to examine the hypothesis that the returns based on the price and return indices follow a random walk process. The results provide convincing evidence that returns series based on price indices indeed follow a random walk. However, when the same tests are performed for the returns based on the return indices, the findings reveal that these series are not weak-form efficient, suggesting that investors might be able to predict future returns based on the current and past data.

Keywords: Weak form market efficiency, ADF, Variance ratio, Ranks and Sign test, Tanzanian stock market

JEL Classification: C14, G14

1. Introduction

As a prominent theory in finance, the efficient market hypothesis (EMH) posits that if prices are determined rationally, then the arrival of new information will cause them to change. This means that the price of a security at any time reflects that investors use optimally all available information. By definition though, information is understood to arise instantly, in an unpredictable manner, and ubiquitously to all investors. It follows that no single individual has a chance to outperform the market as the current price of an asset is the best estimate of its fundamental value (Fama, 1970).

There are three types of efficiency: allocational, operational, and informational. Operational efficiency is concerned with the cost to buyers and sellers of securities on the exchange markets. Allocational (economic) efficiency refers to the supply of scarce resources to the most productive sectors of the economy. And finally, informational (pricing) efficiency is about the extent to which market prices accurately and instantly reflects all available and relevant information, hence implying a true representation of the fundamental value of the underlying asset. It is what the term 'efficient' market hypothesis implies.

The EMH is normally defined and tested in three different forms, namely the weak-form, semi-strong form, and the strong-form (Fama, 1970, 1991). The purpose of this paper is to investigate the weak-form efficiency of Dar-es-salaam stock exchange. A stock market is described as a weak-form efficient if market participants cannot predict the prices of securities and hence generate abnormal returns (except by chance) by analyzing movements of past information e.g. prices and trading volumes. This means that various mechanical investment strategies such as technical analysis or fundamental analysis are of no value, since all investors already know any signal conveyed by the historical information about future performance. Therefore, security prices will always be at equilibrium.

Most of the research on market efficiency over the past decades has centred on the world's developed and emerging equity markets. Little attention is given to the African frontier markets despite them offering relatively higher rates of return than mature markets and attracting significantly the interests of foreign investors. More specifically, in spite of their economic potential for efficient allocation of scarce resources, the weak-form efficiency studies of the East African stock exchanges are sparse. The evidence from the relatively few known studies is too old to explain the prevailing states of market developments that can be attributed to the benefits of globalization, industrialization and technology. Moreover, the results offer contradicting conclusions (see e.g., Dickinson and Muragu, 1994; Smith et al., 2002).

To our best knowledge, we offer for the first time an extensive examination of the behavior of stock prices and returns of the Dar-es-Salaam stock exchange (DSE) since its inception. The market was established in 1996 and started listing companies in 1998 where only one company was listed. Since then the total number of listed companies has increased to 21 by the end of 2014. Of these 14 are domestic while seven are cross-listed; six from Nairobi Stock Exchange and one from London Stock Exchange. The DSE was also declared as the best bourse in Africa in the same year in terms of market capitalisation.

In this paper, we extend the evidence on the weak-form efficiency in frontier markets by examining the data from the DSE in Tanzania. We use the all share index (DSEI) including domestic and foreign firms and Tanzania share index (TSI) covering only domestic firms as well as the indices covering banks, finance, and investment companies (BI), commercial services (CS), and industrial and allied sectors (IA). We employ the Augmented Dickey-Fuller test, variance-ratio test, and rank and sign-test to determine if the results are consistent. We find that returns series based on price indices for DSEI, BI, and CS indeed follow a random walk, whereas the returns series for TSI and

IA do not follow this pattern. However, findings on the same tests for the returns based on the return indices show that they are all not efficient.

The remainder of this paper is organized as follows. In section 2, we provide an overview of the random walk theory and the EMH. In section 3, we present the empirical literature review on the random walk. Furthermore, sections 4 and 5 describe the data and research methodology, respectively. Section 6 presents empirical findings. Section 7 provides concludes the paper and provides some recommendations.

2. The random walk theory and the efficient market hypothesis

The random walk (RW) theory is a version of the EMH. The EMH is built on the assumption that an efficient market is made up of a large number of active, rational, and profit-maximizer individuals who compete to predict the future market values of securities on the basis of freely available information (Fama, 1965, 1970, 1995). The author further points out that the competition among the many participants, and the instantaneous incorporation of both current and expected information, makes the actual prices good estimates of the intrinsic value of the security.

Fama (1970) posits that the RW model should be regarded as an extension of the general expected return or fair game efficient market model. The theory states that given all past information about the stock, the price P_{it} of a firm i today ($t = 0$), is expected to be equal to tomorrow's price P_{it+1} . The term "expected" means that the chance of the stock price to raise is the same as it is to fall. Hence, the change is equal to zero. The relationship can be summarized as follows:

$$P_{it} = P_{it-1} + \varepsilon_{it} \quad (1)$$

For $i = 1, 2, \dots, N$ traded stocks; periods $t = 1, 2, \dots, T$; and where P_{it-1} is the lagged price of the stock; and ε_{it} is an identically and independently distributed random variable with zero mean and unit variance ($\varepsilon_{it} \sim i.i.d, N(0, \sigma^2)$). The RW model can econometrically be written as:

$$P_{it} = \beta P_{it-1} + \varepsilon_{it} \quad (2)$$

where β is the slope coefficient. This is the most restrictive definition of the RW than the martingale version because it requires the price P_{it+1} to be statistically independent. The martingale difference sequence (MDS), on the other hand, is less restrictive. It relaxes the independence restriction on the memory of the sequence (Gaussian-random variable assumption). That is, while agreeing the successive price changes are not predictable, it allows for the predictability of the conditional variances of price changes from past variances (Ntim et al., 2011). In other words, an asset's price series at time t is an MDS if it satisfies the following conditions; $E[P_t] < \infty$, and $E[P_{t+1}-P_t \mid P_t, P_{t-1}, \dots]=0$.

If β is equal to 1, then equation (2) can be written as $P_{it} - P_{it-1} = \Delta P_{it} = \varepsilon_{it}$. The model in this case is said to have a unit root problem or to be at a non-stationary situation. That is, ΔP_{it} is a RW or non-stationary, which means, the changes in price are explained by the error term. The term unit root is derived from the fact that the slope coefficient $\beta = 1$. This is a necessary condition for the random walk model to hold.

Since the establishment of the RW property of the stock return generating process there has been a long-standing interest by researchers for testing the behavior of stock price changes. This is because the theory has obvious investment strategy and economic implications. For example, if a stock market does not follow a RW, it may be considered inefficient, implying that the stock prices are not appropriately priced. Thus, intelligent market participants may use the mispricing to predict the path by which the actual prices would move and make profit out of it. The presence of the random walk is, therefore, essential if the stock markets are to generate the expected benefits such as improving availability and allocation of capital to a real economy.

3. Empirical literature review on weak-form market efficiency

The early tests of the EMH dealt with testing for the presence of the random walk on stock prices (Fama, 1965, 1970, 1991, 1995). Until the early 1970s, for example, there was little evidence against

the EMH. Since mid 1980s, however, contradictory evidence began to emerge regarding the applicability of the EMH in explaining the functioning of the financial markets and their participants (De Bondt and Thaler, 1985, 1987; Pesaran and Timmermann, 1995). These scholars found that stock returns could be predicted by means of the available information. To many, this contradiction seemed to be the evidence of market inefficiency which in turn was interpreted as an indication of the invalidity of the EMH.

In the wake of this, scholars like Fama and French (1988a), Pesaran and Timmermann (1995) and Lander et al. (1997) have examined the performance of various investment strategies in predicting returns or timing the market. Their findings indicate that various financial variables such as price to earnings (P/E) ratio, dividends to price ratio, and short-term treasury bills, can be used to forecast future movements of stock returns.

The discovery of the mean-reversion property in stock prices provides another evidence of returns' predictability. The mean-reversion hypothesis states that stock prices contain a predictable temporary component that is mean-reverting. That is, actual stock prices temporarily swing away from their fundamental values from time to time, but will then tend to go back to their mean. Using autocorrelation based tests, Fama and French (1988b) show a slowly mean-reverting component of stock prices. They assert that their findings indicate that the returns are positively and negatively correlated for short- and long-horizons, respectively. They further indicate that the predictable variations of 3-to-5-year stock returns' variances range from 25% to 40% for large- and small-firm portfolios, respectively. Jegadeesh (1991), however, disputes these observations. He points out that although the equally weighted index of stocks traded on the New York Stock Exchange (NYSE) exhibited mean-reversion over the 1926–1988 period, the evidence showed both the equally weighted and the value weighted indices exhibited significant serial correlation after the World War

II. More specifically, the long-term mean-reversions concentrated in January, and no evidence of the same was found when the entire sample period was included in the test.

Other studies on mean-reversion and the stock price predictability use the variance-estimator testing methodology. Poterba and Summers (1988) and Lo and MacKinlay (1989) contend that variance-ratio tests are more powerful tests of the null hypothesis of market efficiency since they give more precise estimates of the relative predictability of returns over different horizons. In their study, Poterba and Summers (1988) report the existence of a transitory component in stock prices, with returns showing positive autocorrelation over a period of less than one year, and negative autocorrelation over longer periods of up to eight years. Using the variance-ratio test to analyze the importance stationary components in stock prices, they report negative autocorrelations for both real and excess returns at long horizons. They further caution that non-trading effects can affect the interpretation of the positive serial correlation at a short horizon.

Lo and MacKinlay (1988) test the random-walk hypothesis for weekly market returns from data sampled at different sampling frequencies. In contrast to Fama (1998), they find both statistically and economically significant positive serial correlation for the weekly and monthly stock returns. They further show that, during the 1962-1985, the weekly first-order autocorrelation coefficient of the equal-weighted return index was 30% while that of the value-weighted average was 8%. Their results provide evidence against the random walk model, though do not support the mean-reversion hypothesis. On the other hand, Mukherji (2011) uses a powerful nonparametric block bootstrap method, and finds evidence of existence of weak mean-reversion that persists for small-company stocks.

There are also numerous empirical studies on weak-form efficiency from emerging stock markets, but like those conducted in developed markets, they provide conflicting evidence. Buguk and Brorsen (2003) examine the informational efficiency of the Istanbul Stock Exchange using its

composite, industrial and financial index weekly closing prices. Their findings indicate that all three series follow a random-walk process. Urrutia (1995) tests the random walk and market efficiency for Latin American emerging equity markets. The findings from the runs test fail to reject the null hypothesis that the markets are weak-form efficient. However, consistent with those obtained by Lo and MacKinlay (1988), variance ratio test rejects the null hypothesis, and that patterns of the test's rejection do not support the mean-reverting process.

Investigations on the weak-form market efficiency on African markets have also attracted the interests of many researchers, but with inconclusive findings. One of the earliest studies on the efficiency of African stock markets was carried by Dickinson and Muragu (1994). The results obtained from using the serial correlation and runs tests suggest that small markets may conform to the weak-form efficiency notion. Magnusson and Wydick (2002) carried out weak-form efficiency tests using 1989 to 1998 data on eight African markets (including South Africa) listed in the International Finance Corporation (IFC) global composite index. The findings reveal that six out of the eight analyzed stock market indices are weak-form efficient. On the contrary, Smith et al. (2002) used multiple variance-ratios to examine the informational efficiency in stock indices of a group of eight largest African markets covering 1990 to 1998. Except for South Africa, the results in seven of these markets reject the random walk hypothesis due to autocorrelation. Despite using a different approach to a group of seven African markets, Jefferis and Smith (2005) obtained similar results to those of Smith et al. (2002).

In another study, Ntim et al. (2011) tested the efficiency of a set of 24 African continent-wide stock price indices and eight individual African national stock price indices using ranks and signs tests. The results showed that in comparison to the national indices, the African continent-wide stock price indices exhibited a reasonable normal distribution and had superior weak-form efficiency.

Another multi-country study was carried by Mlambo and Biekpe (2007). They found sufficient evidence to conclude that the markets for Namibia, Kenya and Zimbabwe were generally weak-form efficient. In all other markets, majority of stocks rejected the random walk. In a recent study, Smith and Dyakova (2014) examined the weak-form efficiency of eight African markets. The results of variance ratio test showed Kenya and Zambia markets are the most predictable and least informationally efficient. They further reveal that least predictable stock market was Egypt, followed jointly by South Africa and Tunisia.

In summary, the empirical evidence from the African stock markets is not only small but also still offers contradictory conclusions. Many of these studies, however, indicate that the South African and Egyptian stock markets are weak-form efficient (Jefferis and Smith, 2005; Smith and Dyakova, 2014). The Kenyan market is the only most studied from the Eastern-Africa region (Dickinson and Muragu, 1994; Ntim et al., 2011; Smith and Dyakova, 2014). The main reason could be that it is among the oldest markets in Africa, established in 1954. The Tanzanian market, the DSE, is relatively new and small. The conflicting results from past studies in other markets makes it appealing to undertake an empirical work for the DSE.

4. Data source

The data used in this study are daily closing price of the five DSE indices (i.e., DSEI, TSI, BI, CS, and IA) in local currency, the Tanzanian shilling (TZS). All indices are value-weighted market capitalization. Given the size and newness of the market, it was thought that daily stock prices would yield sufficient observations for meaningful statistical analyses and representative of the true distribution characteristic of the market. The data set for testing the random-walk hypothesis covered the period from January 2007 to December 2014. This choice of the starting time was dictated by the moment when the DSE began to compile electronically and maintain a computerized

database. By December 2006, there were ten listed companies. Therefore, this was when the market was more active. All data were provided by the DSE's market research and development department.

5. Methodology

To our best knowledge, this is the first and the only comprehensive study to test the weak-form market efficiency of the DSE. It was considered therefore, worthwhile to employ a battery of tests to determine whether the conclusions obtained confirm the random-walk hypothesis or are fragile. The following sections present the tests that were performed.

5.1. Augmented Dickey–Fuller test

The Augmented Dickey-Fuller (ADF) test is a popular approach used for testing the unit root null hypothesis. The tests were performed on raw price indices and logarithm-transformed data in both levels and first differences. The ADF test employs the following regression model:

$$\Delta Y_t = \beta_1 + \beta_2 t + \delta Y_{t-1} + \sum_{i=1}^k \alpha_i \Delta Y_{t-i} + \varepsilon_t \quad (3)$$

where Δ = the first difference operator; ΔY_{t-i} = lagged values of the dependent variable, for example, $\Delta Y_{t-1} = (Y_{t-1} - Y_{t-2})$, $\Delta Y_{t-2} = (Y_{t-2} - Y_{t-3})$, etc.; ε_t is a white noise error term; β_1 is a constant; β_2 is a slope coefficient on time trend t ; δ is a coefficient of lagged Y_{t-1} ; and Y_t is the logarithm of the stock price or market price index. Recall under (2), it was asserted that there is a unit root if $\beta = 1$. Econometrically, however, it is argued that this regression equation cannot be estimated using the ordinary least squares (OLS) method. In addition, the hypothesis $\beta = 1$ cannot be tested using the standard t-distribution since the test is based on the residual terms which may be highly autocorrelated, and thus can lead to biased estimate of δ (Gujarati and Porter, 2009). Instead, the ADF test was used. The model in equation (3) was used to examine the returns (ΔY_t) in order to take into account the autocorrelation problem regarding the residual terms. In this specification, the

tested unit root hypothesis was $\delta = 0$ (where $\delta = \beta - 1$). As the literature suggests, in order to attain the white-noise structure in ε_t and the unbiased estimate of δ , it is important to select the appropriate lag length by including enough terms. The choice of the lag length was based on the Schwarz Information Criterion (SIC). Following the above discussion, it is hereby hypothesized that:

H₀: Stock price/returns indices at the DSE follow a random walk process (i.e., $\delta = 0$).

5.2. The variance-ratio test

The major criticism against the ADF tests is that they have low power for testing the unit root null hypothesis. In addressing this problem, Lo and MacKinlay (1988) developed a variance ratio test for evaluating the RW properties of asset prices. One of such properties is the indefinite linear increase of the variance (σ^2) of a series (Y_t) over time (t) [i.e. $\text{Var}(Y_t) = t\sigma^2$], which violates the condition of stationarity. Lo and MacKinlay (1988) model exploited this property that variances of differences of time series data computed over different time intervals are linearly related. Put it differently, they assumed that, if a natural logarithm of a series of stock prices follows a RW (with drift), then the variance of say, monthly-period differences should be four times the variance of the weekly-period differences. Let $nq + 1$ denote observations consisting $X_1, X_2, \dots, X_{nq}$ of a time series. Then, the variance-ratio of q^{th} difference, $\text{VR}(q)$, is defined as:

$$\text{VR}(q) = \frac{\sigma^2(q)}{\sigma^2(1)} \quad (4)$$

where $\sigma^2(q) = 1/q$ times the variance of the q -differences and $\sigma^2(1)$ is the variance of the first differences. Lo and MacKinlay (1988) further provide the formula for computing the values of the estimator of the variance of q -period difference and of the first difference, $\sigma^2(q)$ and $\sigma^2(1)$ respectively, as follows:

$$\sigma^2(q) = \frac{1}{n} \sum_{k=q}^{nq} (X_k - X_{k-q} - q\hat{\mu})^2 \quad (5)$$

where:

$$m = q(nq - q + 1) \left(1 - \frac{q}{nq} \right)$$

and

$$\sigma^2(1) = \frac{1}{nq-1} \sum_{k=1}^{nq} (X_k - X_{k-1} - \hat{\mu})^2 \quad (6)$$

where:

$$\hat{\mu} = \frac{1}{nq} (X_{nq} - X_0)$$

The null hypothesis is $VR(q) = 1$. In principle, two reasons may lead to the rejection of the null of the RW. It could be either due to heteroscedasticity or autocorrelation in the asset's series (Lo and MacKinlay, 1988). Additionally, the authors point out that since volatilities do change over time, a rejection of the RW due to heteroscedasticity would have marginal interest. The $VR(q)$ ratio is formulated with two test statistics to examine the null hypotheses of random walk under the homoscedastic and heteroskedastic assumptions about the error term. The homoscedastic-consistent test allow for the strict RW hypothesis, that is, $\varepsilon_{it} \sim i.i.d; N(0, \sigma^2)$. The heteroskedastic-consistent test, on the other hand, relaxes the strict Gaussian assumption by considering that time-varying volatilities that might exist in a series. Put differently, the test statistic allows for some forms of conditional heteroscedasticity and dependence. It follows that Lo and MacKinlay (1988) variance-ratio is widely employed to test the RW and MDS hypotheses of the weak-form market efficiency (Urrutia, 1995; Buguk and Brorsen, 2003; Ntim et al., 2011). The homoscedasticity test statistic $Z(q)$ is expressed as:

$$Z(q) = \frac{VR(q)-1}{[\phi(q)]^{1/2}} \sim N(0,1) \quad (7)$$

where:

$$\phi(q) = \frac{2(2q-1)(q-1)}{3q(nq)} \quad (8)$$

If we assume heteroskedastic increments, then the test static $Z^\wedge(q)$, that uses overlapping intervals, is more robust. This is expressed as:

$$Z^{\wedge}(q) = \frac{VR(q)-1}{[\phi^{\wedge}(q)]^{1/2}} \sim N(0,1) \quad (9)$$

where:

$$\phi^{\wedge}(q) = \sum_{j=1}^{q-1} \left[\frac{2(q-1)}{q} \right]^2 \hat{\delta}(j) \quad (10)$$

and

$$\hat{\delta} = \frac{\sum_{k=j+1}^{nq} (X_k - X_{k-1} - \hat{\mu})^2 (X_{k-j} - X_{k-j-1} - \hat{\mu})^2}{\sum_{k=1}^{nq} [(X_k - X_{k-1} - \hat{\mu})^2]^2} \quad (11)$$

It should be noted, however, that the two test statistics derived by Lo and MacKinlay (1988) focus on testing individual variance-ratios for a specific aggregation interval, q , and not for all periods (Smith et al., 2002). A procedure that Chow and Denning (1993) developed provides a means of comparing multiple sets of variance-ratio estimates with unit. The decision criterion to reject the null hypothesis against the alternative hypothesis is based on whether any of the estimated variance-ratios is significantly different from unity.

5.3. Rank and sign-based variance-ratio tests

Lo and MacKinlay (1988) variance-ratio test, like other parametric tests, is criticized for lacking power, particularly when the series of data are non-normal. Wright (2000) argues against the assumption that a time series is independently and identically distributed (*i.i.d*). The author contends that the hypothesis may not always hold when the data are serially correlated. He further points out that although Lo and MacKinlay showed that the test can be robustified against conditional heteroskedasticity, the finite-sample null distribution of the test statistic is asymmetric and non-normal. As a remedy, Wright (2000) proposed an alternative non-parametric test that uses standardized ranks and signs instead of the underlying asset's returns differences. One advantage of the non-parametric statistical tests is that they are exact, avoid making asymptotic approximation, and have low size distortion. The other is that the tests are more powerful in the presence of highly

non-normal data compared to the conventional tests. The evidence in table 1 (based on JB statistics) justifies the application of this method to analyze our data.

Wright's (2000) alternative model is described as hereunder. Letting $r(y_t)$ be the rank of y_t among $y_1, y_2, y_3, \dots, y_T$, he defined:

$$r_{1t} = \left(r(y_t) - \frac{T+1}{2} \right) / \sqrt{\frac{(T-1)(T+1)}{12}} \quad (12)$$

$$r_{2t} = \Phi^{-1}(r(y_t)/(T+1)) \quad (13)$$

where Φ is the standard normal cumulative distribution. Wright (2000) defines the series r_{1t} as a simple linear transformation of the ranks, standardized to have a sample mean 0 and sample variance 1. The same applies to series r_{2t} (the inverse normal or van der Waerden scores), it has a sample mean 0 except that it has a sample variance of approximately equal to 1.

The rank-based ratio tests are calculated by substituting r_{1t} and r_{2t} for y_t in Lo and MacKinlay (1988) definition of variance-ratio test statistic. Wright (2000) therefore proposed the following test statistics:

$$R_1 = \left(\frac{\frac{1}{Tk} \sum_{t=k+1}^T (r_{1t} + r_{1t-1} + \dots + r_{1t-k})^2}{\frac{1}{T} \sum_{t=1}^T r_{1t}^2} - 1 \right) \left(\frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2} \quad (14)$$

$$R_2 = \left(\frac{\frac{1}{Tk} \sum_{t=k+1}^T (r_{2t} + r_{2t-1} + \dots + r_{2t-k})^2}{\frac{1}{T} \sum_{t=1}^T r_{2t}^2} - 1 \right) \left(\frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2} \quad (15)$$

According to Wright (2000), the tests based on ranks are exact under *i.i.d* assumption. The rank $r(y_t)$ is a random permutation of the numbers 1, 2, ..., T, each with equal probability. The proposed modified test statistics based on the signs are as described hereafter. Wright (2000) shows that, for any series y_t , if $u(x_t, q) = 1(x_t > q) - 0.5$, then $u(x_t, 0)$ is 0.5 if x_t is positive and -0.5, otherwise. Also, if $S_t = 2u(y_t, 0) = 2u(\varepsilon_t, 0)$; and S_t is an *i.i.d* series with mean 0 and variance 1, then each S_t is equal to 1 with probability $1/2$ and is equal to -1, otherwise. The test statistic (S_t) is defined as:

$$S_1 = \left(\frac{\frac{1}{Tk} \sum_{t=k+1}^T (s_t + s_{t-1} + \dots + s_{t-k})^2}{\frac{1}{T} \sum_{t=1}^T s_t^2} - 1 \right) \left(\frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2} \quad (16)$$

As with the ranks, Wright (2000) points out that exact sampling distribution of S_1 can be simulated for different levels of T and k . Moreover, the author also demonstrates in a Monte-Carlo experiment that the tests based on ranks are consistent with the homoscedasticity assumption, while the signs are robust under the conditional heteroscedasticity assumption.

6. Analyses and presentation of the results

The following sub-sections discuss the test results of the study.

6.1. Descriptive statistics and diagnostics

The analyses in the subsequent sections were performed on the raw and the logarithm transformed daily price indices. The daily returns (R_t) are computed from the logarithm transformation of the daily price data as follows:

$$R_t = \ln(P_t) - \ln(P_{t-1}) \quad (17)$$

where P_t is the value of the index at time t ; P_{t-1} is the one day lagged daily value; and $\ln$ is the natural logarithm. Table 1 presents the descriptive statistics and diagnostics of the daily data for price indices, log of price indices, and returns of the five studied indices. For the period under consideration, all indices experienced positive mean returns, which are comparable to what Ntim et al. (2011) reports.

[Insert table 1 here]

The lowest daily return in table 1 is 0.05% for the DSEI and the highest one is 0.09% for the IA series. The examination of the standard deviations, a measure of market risk, shows generally that volatility for the percentage returns is considerable for all indices. The lowest deviation, which implies lower risk, is observed in DSEI and TSI indices. Furthermore, the results reveal that the CS has the highest standard deviation. This is followed closely by BI and IA indices.

The measures of skewness and kurtosis for the perfectly normal distribution are supposed to be zero and three, respectively. Considering the results for log price index in table 1, it is found that CS has extremely high negative skewness. Likewise, the returns for BI and CS are negatively skewed. This kind of returns' distribution implies that the returns are characterized by regular little gains and few excessive losses. On the contrary, the returns for the remaining indices are positively skewed suggesting that they are experiencing frequent small losses with lesser chances of extreme gains. The results for kurtosis tests show that the values for the price indices and percentage returns for all indices are positive and greater than three. This means that the distributions are leptokurtic relative to normal distribution. In other words, there is strong evidence that the data are highly concentrated around the mean because the variation within the observations is low. This observation signifies that the market exhibits unexpected moments of very low and high returns. It is therefore plausible to argue that these indices are more likely to attract risk-averse investors who avoid large return surprises that cause high variation in their stocks holdings. This is supported by the fact that the annualized¹ daily returns appear to be reasonably high. They range from a low of 12.5% in the case of DSEI to a high of 22.5% in the case of IA. Correspondingly, the Jarque-Bera test statistic strongly rejects the normality hypothesis for all indices. This observation is consistent with the evidence in prior studies that non-normality is a common phenomenon in African stock market price and return series (Jefferis and Smith, 2005; Ntim et al., 2011).

6.2. *Augmented Dickey-Fuller test results*

We employ equation (3) to test the null hypothesis that returns based on price indices and those based on returns indices at the DSE follow a random walk. The decision criterion is to reject the null hypothesis if the test statistic is greater than the critical values in absolute terms. Table 2 show

¹ We calculate the average daily return for each index and then multiply it by the number of trading days in a year (i.e., 250 in our case) to obtain the annualized return.

the results of the ADF test for the five indices. Under the column headed returns based on price indices, we find that, at 5% significance level, we do not have sufficient evidence to reject the null hypothesis regarding levels in Panel A for all indices. This suggests that all series are non-stationary.

Gujarati and Porter (2009) point out that the first differences of a random walk time series are stationary. Accordingly, we performed the ADF test with the first differences to examine the indices for the presence of a second unit root. The results are presented in panel B of table 2 under the returns based on price indices column. It is observed that the null hypothesis of a second unit root is rejected except with TSI. The findings reveal absence of a second unit root and confirm the contention in Gujarati and Porter (2009) for the four series. Hence, we can verify that all series under consideration satisfy the necessary condition for the random walk to hold.

[Insert table 2 here]

Panel A, in table 2 also presents the results of the ADF test for the returns based on return indices. The examination of the findings provides strong evidence against the null of a unit root. Hence, we can confirm that all series under consideration are stationary in levels.

6.3. Variance-ratio test results

We employ Lo and MacKinlay (1988) variance-ratio test to test the random walk for each of the five DSE indices. We perform the tests using daily sampling intervals (q) of 2, 4, 8, and 16 days. As there is more than one specified period for each time series, Eviews (the data processing package that we use) reports two sets of results. The “joint tests” output presents the results of the joint null hypothesis for all periods, whereas the output of the “individual tests” relates to the tests for individual periods. Table 3 reports the variance-ratio estimations.

In panel A of the column labelled returns based on price indices, we present the results assuming homoscedasticity in the error term. The findings provide strong evidence to reject the joint null

hypothesis at a 5% significance level for all indices, except the DSEI series. A further analysis of the individual tests shows that the null of a RW is strongly rejected for the indices BI, IA, and TSI. Moreover, we find that the null is rejected for the first two intervals of the CS, whereas it is not rejected with DSEI.

We present the heteroscedastic-consistent results in panel B of the column labelled returns based on price indices. The findings reveal that the null of joint hypotheses cannot be rejected with all indices except with TSI. Similarly, the individual variance ratio tests fail to reject the null hypothesis that the stock price indices in the DSE follow a random walk process, except for period 8 and 16, respectively of the TSI. We fail to reject the random walk hypothesis in both panels for the DSEI series for all lags. The findings under returns based on price indices in table 3 provide strong evidence against the null hypothesis in panel A but not in panel B with BI, CS, and IA series. Accordingly, these rejections may be due to heteroscedasticity in the residuals. It, therefore, means that the respective indices meet some of the strict conditions of the random walk process. Furthermore, the results suggest strong rejection of the RW with TSI regardless of the type of the increment assumption; this implies that the rejection is due to autocorrelation of the daily increments.

[Insert table 3 here]

Under the null hypothesis suggesting a RW, our interest is whether the value of the variance-ratio equals one, and the test statistics have asymptotically standard normal distribution. The findings under returns based on price indices in table 3 reveal that the majority of the variance-ratios (except for TSI) are less than unity, and the $Z^{\wedge}$ scores are not significant (panel B). Moreover, only TSI series has its joint test statistics significantly different from one at the 1% level at lag 16. The individual tests are significant at all lags at the 5% level in panel A. The significance of the results, however, disappears at lag 2 and 4 when heteroscedasticity-consistent estimators are used.

Generally, this observation indicates a mean-reversion across the lags of TSI (see also Lo and MacKinlay, 1988; Grieb and Reyes, 1999). Furthermore, notice also that the variance ratios' results obtained with the DSEI are consistent with the implication of ADF results.

Next, we use the variance ratio test to examine the behavior of the returns of stock price indices for all five indices. The findings under returns based on return indices in table 3 presents the results. The RW hypothesis is strongly rejected for all five indices under both joint and individual tests (panel A).

6.4. Rank and sign-based variance ratio tests

In this section, we report the results of Wright's (2000) rank and sign-based variance-ratio tests (table 4). The tests are based on daily sampling intervals (q) equal to 2, 4, 8, and 16 days as before. EViews perform Wright's (2000) variance-ratio test with ties replaced by the average of the tied ranks for the tied data (using the 'tie handling' option). It uses the permutation bootstrap to obtain the probabilities for the joint and individual test statistics. The standard error estimates for the ranks (R_1 , R_2) and signs assume no heteroskedasticity.

[Insert table 4 here]

The findings under returns based on price indices in panel A reveal that the null hypothesis of an RW cannot be rejected for the DSEI, BI, and CS indices for all intervals of q . In the case of the IA index, the RW is rejected only for $q = 2$. For the TSI index, on the other hand, we fail to reject the null at periods 2 and 4. In contrast, the results obtained using the sign-based variance-ratio test strongly rejects the heteroscedasticity-consistent random walk for all indices for all intervals except the CS (panel B).

We also assess whether Lo and MacKinlay (1988) and Wright (2000) variance-ratio tests produce consistent results. The findings reveal that the two versions reach to similar conclusions under the *i.i.d* for all indices except the BI and CS. For these data series, Lo and MacKinlay

variance ratio rejects the null under the homoscedasticity assumption. Wright's ranks and rank scores, however, do not reject the same hypothesis. The opposite is true in the presence of heteroscedasticity. All indices, except the TSI, fail to reject the null of the random walk hypothesis under Lo and MacKinlay test. However, only the CS and TSI indices produce consistent results when the Wright's sign test is used. That is, all indices except the CS rejected the hypothesis that the series are random walk, which is also implied in table 3 when we reject the null hypothesis that variance-ratio equals one.

Table 4 presents the results of Wright's tests for the returns of the five stock price indices. The findings in Panel A of the column labelled returns based on return indices indicate that R_1 and R_2 strongly reject the null hypothesis that the series follow the random walk process for all indices at 1% significance level. The sign-based tests slightly offer mixed results (panel B). For the DSEI index, the results are in agreement with those in the rank-based tests. For the BI and TSI indices, the null hypothesis is rejected only for $q = 2$ with the sign test. For the IA index, the sign test rejects the null hypothesis of a random walk for $q > 4$, but fails to support the null hypothesis with $q = 2$ and 4 intervals at the 5% level. Prior studies have also observed that individual tests tend to provide contradictory results (see Wright, 2000; Buguk and Brorsen, 2003; Ntim et al., 2011). That is, tests reject the null hypothesis for some periods of q but fail to reject for others. In the case of CS, the sign test rejects the random walk hypothesis at period 4 and 8 at the conventional significance level. Except for the CS index, the joint tests for all other cases provide strong evidence against the random walk hypothesis.

The individual interval results obtained from Wright's rank-based tests for the returns of the stock price indices are all consistent with Lo and MacKinlay (1988) results under the homoscedasticity assumption. Under the heteroscedasticity-consistent tests, however, only DSEI and TSI indices' results are in agreement with each other. Specifically, we fail to reject the null hypothesis of the

random walk for BI, CS and IA under Lo and MacKinlay variance-ratio test in both cases; that is, the joint and individual tests. By contrast, the sign-based variance-ratio test results are ambiguous. The joint test is strongly against the null hypothesis for the BI and IA series. The individual test results for the same index rejects the hypothesis for $q > 2$. For the CS, we fail to reject the null hypothesis under the joint test at the 5% level of significance. The individual tests, however, reject the random walk hypothesis for $q = 2$ and 16.

6.5. Summary of the findings

Wright (2000) shows that ranks and sign tests can be exact, and have better power properties than the conventional variance-ratio tests. Moreover, they demonstrate that rank-based variance-ratio tests have, generally, more power than the sign-based variance-ratio tests. Considering Lo and MacKinlay (1988) variance-ratio tests, Wright (2000) demonstrates that the non-robust $Z_{(q)}$ is more powerful than the heteroscedastic robust $Z^{\wedge}_{(q)}$. The summary presented in table 5 is based on these rejection rules. This is because the data in this study are highly non-normal (see table 1).

Panel A indicates that the results obtained from the three different tests are largely mixed. The ADF test shows that all five indices fail to be in line with the random walk hypothesis. However, the method is widely criticized for lacking power, particularly, when data exhibit non-normality. For the DSEI, BI and CS indices, nevertheless, the results are supported by the powerful rank-based variance-ratio test results. In contrast, the ranks and sign tests do not support the findings of the ADF test for the IA and TSI indices. Specifically, rank and sign-based variance-ratio test results are consistent with the homoscedastic random walk findings based on variance ratio tests.

[Insert table 5 here]

The results in panel B seem to be more consistent under the three different tests. The ADF, $Z_{(q)}$, R_1 , R_2 , and S tests largely reject the random walk hypothesis. Three indices under $Z^{\wedge}_{(q)}$, on the other hand, fails to reject this hypothesis.

7. Conclusions and recommendations

Tanzania, like many other African countries, established her stock exchange in 1996, as part of implementation of economic reforms that were sponsored by the IMF and the World Bank. The DSE became operational in 1998. Since then, however, to the best knowledge of the researchers, there has been no study that has vigorously examined the behavior of stock price and returns movement in the market from the efficiency point of view. The aim of this study is, therefore, to examine the price and return behavior of the DSE stock price indices using the ADF, conventional variance-ratio, and rank and sign-based nonparametric variance-ratio tests.

Our analysis shows that the distribution of the data used in the study are highly non-normal. According to Wright (2000), with such characteristics the conventional tests yield ambiguous results. He further points out that the ranks- and sign-based tests have better power properties than the former mentioned tests. The conclusions derived from the latter may therefore be the most appropriate ones as they represent the true underlying values of the series. Hence, we can conclude that the DSE's all share index, banking, finance and investment index, and the commercial services index tend to have patterns that are in line with the random walk hypothesis. In contrast, the industrial and allied sector index and the Tanzania share index are conclusively not weak-form efficient. Moreover, the results show that the returns based on return indices do not follow a random walk process for all five indices at the DSE, implying the absence of the weak-form efficient markets.

The literature suggests a number of possible explanations for the rejection of weak-form efficiency in infant stock markets that can squarely be applicable to the DSE (Smith et al., 2002; Ntim et al., 2011). The first reason is the absence of derivatives trading. This is associated with lower speed, poor quality and increased costs related to information processing. As a result, any profit opportunity arising from price disagreement from traders may take long time to be brought back to

equilibrium levels. In other words, absence of derivative trading negatively impacts the DSE's objective of facilitating a price discovery process in the market.

Second, DSE is considered to be a very small stock exchange in terms of market capitalization, and number and size of individual stocks. The market has listed only 21 companies to date. The DSE recorded an impressive growth in market capitalization since its establishment. For instance, its market capitalization grew by 249% from US\$ 3.669 billion in October 2011 to US\$ 12.800 billion in December 2014, (DSE, Quarterly Reports). Again, the market topped the list in 2014 in terms of market capitalization among the African stock exchanges.

Low liquidity and non-trading at the DSE may also explain the rejection of the weak-form efficiency. According to Demirguc-Kunt and Maksimovic (1996), liquid secondary capital markets lower informational asymmetry and transaction costs. Despite the fact that the market has grown rapidly, the DSE suffers from low liquidity. Its average stock turnover ratio in 2011 and 2012 were 2.5% and 1.6%, respectively (World Bank, 2015). A ratio of such magnitude demonstrates inability of the market to facilitate active price formation process (Smith et al., 2002; Ntim et al., 2011). Related to liquidity are the behavior of the market participants. Many investors at the DSE may be classified as short-term investors. They are characterized by targeting on dividends as the main source of investment return instead of trading on stocks price fluctuations.

In addition, the majority of Tanzanians have not yet appreciated the importance of the stock exchange as an option for diversifying their investments. This is evidenced by the low involvement of individuals in the market, which is estimated to be less than one percent of the population. Indeed, with such a small number of investors, it is not expected to have a vibrant capital market. The government has, however, lifted the restrictive 60% limit for foreigners' ownership in order to attract a critical mass of investors to enhance liquidity (CMSA, 2014).

Our results reveal that returns based on price indices for the IA and TSI series are not weak-form efficient. These indices are composed by domestic companies only. Moreover, due to the effects of home bias, the majority of their floating shares are locally owned. As mentioned earlier, since many individual investors tend to buy and hold on these stocks, they cause thin-trading in the market. This behavior causes a mismatch between the supply of and demand for a particular stock. Consequently, higher demand for (supply of) these stocks push (pull) their prices up (down). It is, therefore, potentially easy to speculate the direction of prices for these stocks.

According to the EMH, stock pricing inefficiencies may be exploited by sophisticated investors (Fama, 1965). Our findings do not in any way provide evidence indicating that exploitation of these inefficiencies is profitable or not. As a recommendation for further study, though, it appears that there is a need to examine whether investors at the DSE can systematically make abnormal returns. The main policy implication from this evidence is that the market is still not informationally efficient. The regulatory authorities have to put more efforts to promote and develop the functioning of the capital markets sufficiently to produce significant informational efficiency. Successful implementation of this role is a vital incentive for attracting foreigners. The economic implication is that stocks at the DSE are not appropriately priced at their equilibrium values. This distortion in pricing may have repercussions on the ability of the market to play the crucial role of capital allocation and risk pricing.

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Table 1: Descriptive statistics and diagnostics

		Mean	Std. Dev.	Skewness	Kurtosis	Jarque-Bera	N
DSEI	Price Index	1,373.053	388.374	1.709	5.390	1439.677**	1,986
	Returns index	7.192	0.243	1.203	3.809	533.388**	1,986
	Returns (%)	0.050	0.572	8.189	170.412	2,247,083**	1,906
BI	Price Index	1,418.895	898.351	2.247	8.467	4145.724**	1,987
	Returns index	7.126	0.465	1.333	4.266	720.748**	1,987
	Returns (%)	0.065	9.253	-5.450	604.589	28,781,265**	1,908
CS	Price Index	1,339.551	526.858	0.847	4.334	384.292**	1,982
	Returns index	7.082	0.646	-4.115	25.231	46,408.67**	1,982
	Returns (%)	0.057	10.196	-15.128	929.067	68,037,409**	1,902
IA	Price Index	1,700.499	1,319.610	2.488	8.961	4990.162**	1,986
	Returns index	7.255	0.550	1.031	4.747	604.257**	1,986
	Returns (%)	0.090	6.656	15.051	875.410	60,515,964**	1,906
TSI	Price Index	1,487.536	1,014.086	2.385	8.129	4061.442**	1,987
	Returns index	7.164	0.4702	1.652	4.611	1,118.730**	1,987
	Returns (%)	0.079	0.585	2.369	28.379	52,988.63**	1,908

Notes: The asterisk ** represents statistical significant at 1% level.

Table 2: ADF unit root test results

Model	Returns based on price indices					Returns based on return indices				
	DSEI	BI	CS	IA	TSI [#]	DSEI	BI	CS	IA	TSI
<i>Panel A: Level</i>										
Intercept	2.837	1.514	0.353	3.097	5.999	-27.866**	-25.488**	-7.793**	-22.022**	-18.687**
Trend and intercept	0.548	2.140	-1.745	1.670	5.452	-27.997**	-25.482**	-7.794**	-22.296**	-25.489**
<i>Panel B: First difference</i>										
Intercept	-28.335**	-29.696**	-56.506**	-16.185**	3.497					
Trend and intercept	-28.561**	-29.681**	-56.546**	-16.443**	2.789					

Notes: **, * means significant at the 1% and 5% level, respectively. Critical values for the ADF test for 1%, 5%, and 10% significance levels with the constant model are -3.43, -2.87, and -2.57, respectively. For the trend and constant model, the critical values are -3.97, -3.42, and -3.13, respectively. [#] Stationary in second differences. DSE all share index (DSEI) consisting of home and foreign firms, and Tanzania shares index (TSI) covering only domestic firms. It also includes the indices covering banks, finance, and investment companies (BI), commercial services (CS), and industrial and allied sectors (IA).

Table 3: Lo and MacKinlay variance ratio test results

		Sampling interval (q-days)							
Series		2	4	8	16	2	4	8	16
Returns based on price indices					Returns based on return indices				
<i>Panel A: Variance-ratios assuming homoscedasticity</i>									
DSEI	VR _(q)	0.9683	0.9991	1.0052	0.9863	0.4499	0.2373	0.1140	0.0490
	Z _(q)	-1.3857 ^a	-0.0202	0.0770	-0.1364	-23.5185 ^{**a}	-17.4313 ^{**}	-12.8066 ^{**}	-9.2369 ^{**}
BI	VR _(q)	0.5638	0.3807	0.2889	0.2168	0.3321	0.1560	0.0942	0.0436
	Z _(q)	-19.0518 ^{**a}	-14.4589 ^{**}	-10.5009 ^{**}	-7.7717 ^{**}	-28.5776 ^{**a}	-19.3041 ^{**}	-13.1029 ^{**}	-9.2976 ^{**}
CS	VR _(q)	0.8324	0.9085	0.9392	0.9975	0.3419	0.4120	0.2395	0.0852
	Z _(q)	-7.3099 ^{**a}	-2.13409 [*]	-0.89729	-0.0253	-28.1066 ^{**a}	-13.4236 ^{**}	-10.9804 ^{**}	-8.8756 ^{**}
IA	VR _(q)	0.5644	0.3873	0.3057	0.2734	0.3292	0.1338	0.0980	0.0456
	Z _(q)	-19.0173 ^{**a}	-14.2984 ^{**}	-10.2468 ^{**}	-7.2067 ^{**}	-28.6783 ^{**a}	-19.7951 ^{**}	-13.0372 ^{**}	-9.2707 ^{**}
TSI	VR _(q)	0.9543	1.1368	1.5088	2.1505	0.3936	0.2182	0.1131	0.0549
	Z _(q)	-1.9947 [*]	3.1945 ^{**}	7.5136 ^{**}	11.4172 ^{**d}	-25.9491 ^{**a}	-17.8806 ^{**}	-12.8297 ^{**}	-9.1876 ^{**}
<i>Panel B: Heteroscedasticity-consistent variance-ratios</i>									
DSEI	VR _(q)	0.9683	0.9991	1.0052	0.9863	0.4499	0.2373	0.1140	0.0490
	Z [^] _(q)	-0.6471 ^a	-0.0104	0.0469	-0.0970	-3.6069 ^{**a}	-3.2777 ^{**}	-3.1914 ^{**}	-3.1038 [*]
BI	VR _(q)	0.5638	0.3807	0.2889	0.2168	0.3321	0.1560	0.0942	0.0436
	Z [^] _(q)	-1.1604 ^a	-1.0984	-1.0811	-1.1077	-1.6250 ^a	-1.3332	-1.2062	-1.1790
CS	VR _(q)	0.8324	0.9085	0.9392	0.9975	0.3419	0.4120	0.2395	0.0852
	Z [^] _(q)	-1.0972 ^a	-0.3995	-0.2276	-0.0089	-5.23440 ^a	-2.4726	-2.4590	-2.6416
IA	VR _(q)	0.5644	0.3873	0.3057	0.2734	0.3292	0.1338	0.0980	0.0456
	Z [^] _(q)	-1.2245 ^a	-1.1482	-1.1152	-1.0893	-1.5157 ^a	-1.2770	-1.1245	-1.1036
TSI	VR _(q)	0.9543	1.1368	1.5088	2.1505	0.3936	0.2182	0.1131	0.0549
	Z [^] _(q)	-0.7608	1.3171	3.4315 [*]	5.7070 ^{**d}	-6.2750 ^{**a}	-4.9937 ^{**}	-4.3668 ^{**}	-3.7685 ^{**}

Notes: A test statistic with **, and * indicate a 1% and 5% level of significance, respectively. H₀: VR (q) =1; i.e., the series follows a random walk process. Joint tests Max |z| : at period 2^a, 4^b, 8^c, and 16^d. In other terms, the notations “a, b, c and d” imply the joint significance at 2, 4, 6 or 8 intervals, respectively. DSE all share index (DSEI) consists of home and foreign firms, and Tanzania shares index (TSI) covers only domestic firms. It also includes the indices covering banks, finance, and investment companies (BI), commercial services (CS), and industrial and allied sectors (IA).

Table 4: Rank and sign-based variance ratio test results

Series		Sampling interval (q-days)							
		2	4	8	16	2	4	8	16
		Returns based on price indices				Returns based on return indices			
<i>Panel A: Ranks</i>									
DSEI	R ₁	0.0683	1.8678 ^b	0.7736	0.1822	-20.5618 ^{**a}	-15.2059 ^{**}	-11.4662 ^{**}	-8.3621 ^{**}
	R ₂	-0.7535	0.8851	-0.6153	-1.4370 ^d	-21.7738 ^{**a}	-16.0182 ^{**}	-12.0034 ^{**}	-8.6529 ^{**}
BI	R ₁	0.7253	3.1001	4.2282 ^c	3.8826	-19.4633 ^{**a}	-15.1128 ^{**}	-11.4231 ^{**}	-8.5815 ^{**}
	R ₂	0.7217	3.2266	4.3248 ^c	3.8455	-20.8604 ^{**a}	-16.0052 ^{**}	-11.9907 ^{**}	-8.8800 ^{**}
CS	R ₁	-2.3395 ^a	-1.6912	-1.1391	-0.0348	-21.4935 ^{**a}	-16.4489 ^{**}	-12.2150 ^{**}	-8.7777 ^{**}
	R ₂	-2.4778 ^a	-1.8932	-1.7035	-0.5575	-21.7865 ^{**a}	-16.7749 ^{**}	-12.4560 ^{**}	-8.9354 ^{**}
IA	R ₁	-3.9440 ^{*a}	-1.2447	0.4351	0.1501	-22.4674 ^{**a}	-16.1964 ^{**}	-11.6697 ^{**}	-8.3594 ^{**}
	R ₂	-4.9501 ^{**a}	-2.1724	-0.1851	-0.0859	-23.5148 ^{**a}	-16.8926 ^{**}	-12.1644 ^{**}	-8.6929 ^{**}
TSI	R ₁	1.1691	3.8551	5.7375 ^{**}	6.7461 ^{**d}	-19.7498 ^{**a}	-15.1220 ^{**}	-11.6126 ^{**}	-8.6359 ^{**}
	R ₂	0.0214	3.1628	5.2309 [*]	5.9696 ^{*d}	-21.8065 ^{**a}	-16.2755 ^{**}	-12.2515 ^{**}	-8.9441 ^{**}
<i>Panel B: Signs</i>									
DSEI	S	6.0928 ^{**}	9.8193 ^{**}	10.6589 ^{**c}	8.9687 ^{**}	-9.4024 ^{**a}	-5.2383 ^{**}	-2.6014 ^{**}	-1.9421 ^{**}
BI	S	20.8788 ^{**}	31.4615 ^{**}	39.6025 ^{**}	41.6005 ^{**d}	9.6985	18.9999 ^{**}	26.3125 ^{**}	28.0263 ^{**d}
CS	S	28.4785	42.2599	52.2302	52.4674 ^d	22.7590	35.1189 [*]	44.2007 [*]	45.0331 ^d
IA	S	9.2538 ^{**}	15.9165 ^{**}	19.3160 ^{**c}	16.5376 ^{**}	-1.4033	5.5634	9.9548 ^{**c}	9.5938 [*]
TSI	S	15.6591 ^{**}	23.6665 ^{**}	29.2549 ^{**}	30.9033 ^{**d}	3.9495	11.6298 ^{**}	17.7681 ^{**}	20.4940 ^{**d}

Notes: A test statistic with **, and * indicate a 1% and 5% level of significance, respectively. H₀: The series follow a random walk process. Joint tests Max |z|: at period 2^a, 4^b, 8^c, and 16^d. In other terms, the notations “a, b, c and d” imply the joint significance at 2, 4, 6 or 8 intervals, respectively. DSE all share index (DSEI) consists of home and foreign firms, and Tanzania shares index (TSI) covers only domestic firms. It also includes the indices covering banks, finance, and investment companies (BI), commercial services (CS), and industrial and allied sectors (IA).

Table 5: Summary of the findings

Series	ADF	Z(q)	Z [^] (q)	R ₁	R ₂	S
<i>Panel A:</i> Returns based on price indices						
DSEI	Fail	Fail	Fail	Fail	Fail	Reject
BI	Fail	Reject	Fail	Fail	Fail	Reject
CS	Fail	Reject	Fail	Fail	Fail	Fail
IA	Fail	Reject	Fail	Reject	Reject	Reject
TSI	Fail	Reject	Reject	Reject	Reject	Reject
<i>Panel B:</i> Returns based on return indices						
DSEI	Reject	Reject	Reject	Reject	Reject	Reject
BI	Reject	Reject	Fail	Reject	Reject	Reject
CS	Reject	Reject	Fail	Reject	Reject	Fail
IA	Reject	Reject	Fail	Reject	Reject	Reject
TSI	Reject	Reject	Reject	Reject	Reject	Reject

Notes: ‘Fail’ means the null hypothesis proposing the presence of random walk is not rejected and ‘Reject’ means the null hypothesis is rejected. *ADF* is the augmented Dickey-Fuller test, *Z(q)* is the Z-test based on the variance ratio test assuming homoscedasticity; *Z[^](q)* is the Z-test based on the variance ratio test that is robust to heteroscedasticity; *R₁* is the rank test based on the simple linear transformation of ranks; *R₂* is the rank test based on the inverse normal or van der Waerden scores; and *S* is the sign test.