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On the linkages between Africa's emerging equity markets and global markets: Evidence from fractional integration and cointegration☆

Luis Gil-Alana[a,*], Hector Carcel[b], Emmanuel Joel Aikins Abakah[c]

[a] *University of Navarra, Faculty of Economics and NCID, Edificio Amigos, E-31009 Pamplona, Spain*
[b] *Bank of Lithuania, Vilnius, Lithuania*
[c] *The University of Adelaide Business School, Adelaide, Australia*

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Abstract

This paper uses fractional integration and cointegration for the period of January 2000–June 2018 to investigate the stochastic properties of the bilateral linkages between stock markets in Africa and selected international markets to establish if markets in Africa co-move with the rest of the world. Results from the univariate analysis show that there exists a high degree of persistence with orders of integration about 1 or higher than 1, implying that shocks to these stock markets have significant permanent effects. Concerning bivariate results and testing for cointegration, evidence of cointegration is found for Egypt and Kenya against the UK and the Europe Zone. There are some other cases where partial evidence of cointegration is found, though in general, in all cases, we observe that the degree of cointegration is very low, implying very long periods of convergence.

JEL classification: C32; G15

Keywords: African stock markets; Linkages; Fractional integration; Fractional cointegration

1. Introduction

Integration of international equity markets in recent years has attracted a lot of attention in the finance and economic literature. Phylaktis and Ravazzolo (2002) attribute the increase to deregulation, globalization and advances in information technology in trading systems, leading to the exploration of various aspects of stock markets integration such as volatility across stock market correlation structures, and financial crises contagion (e.g. Claessens and Forbes, 2001; Koedijk et al., 2002; Ng, 2000). The surge in interest in global stock markets integration, which is now central in finance and economic research, is the result of its crucial implications in portfolio allocation, asset pricing and policy formulation. First, Marashdeh and Shrestha (2010) argue that it facilitates risk sharing across integrated markets. Second, Driessen and Laeven (2007) document that diversification of assets help financial investors to benefit by investing in two different stock markets due to the markets cointegration. In addition, it contributes to financial stability by enhancing competition and efficiency in the allocation of resources, as well as reducing the cost of capital and price volatility between integrated markets (Tai, 2007). The integration of stock markets also plays an important role in promoting domestic savings, investment and could positively affect total factor productivity and economic growth (Levine, 2001).

Following the seminal works of Engle and Granger (1987) and Johansen (1988) on cointegration, several studies have explored market integration in developed equity markets, the results of which support the financial integration hypothesis in the developed equity markets (Dumas and Solnik, 1995; De Santis and Gerard, 1997; Pukthuanthong and Roll, 2009; Donadelli, 2013). A large section of research in the literature has focused on regional market integration (e.g. Corhay et al., 1993; Fratzscher, 2001; Yang et al., 2003a, 2003b); on Asian, and Pacific stock markets (e.g. Dekker et al., 2001; Sharma and

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* Corresponding author.
E-mail address: alana@unav.es (L. Gil-Alana).
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Wangbangpo, 2002; Yang et al., 2003a,b); on Latin American stock markets (e.g. Chen et al., 2002); and on cross-regional stock market linkages (e.g. Bessler and Yang, 2003; Chaudhuri and Wu, 2003; Masih and Masih, 1997; Ratanapakorn and Sharma, 2007). Though a plethora of studies have statistically supported the hypothesis of cointegration between the markets, many others documented a lack of it, including for example, Gerrits and Yuce (1999), Huang et al. (2000) and Kwan et al. (1995).

There is a considerable gap in the analysis of the relationship between the stock markets in Africa and in the advanced economies due to the under-development and illiquidity of these markets. In fact, there are only very few papers that have focused on markets in the sub-Saharan region, investigating the relation between African stock markets and the world markets (Adjasi and Biekpe, 2006; Alagidede, 2009; Alagidede et al., 2011). However, recent years have witnessed a considerable improvement and liberalisation in a number of stock markets in Africa as documented by Allen et al. (2010). Due to this and to the growing importance of Africa in the global economy, we re-examine in this paper the linkages between stock markets in Africa and the main stock markets in the world.

Prior studies that focused on the linkages between Africa's equity markets and advanced markets found no empirical evidence on cointegration across the markets. Thus, for example, Alagidede (2009) and Alagidede et al. (2011), using Johansen's (1988, 1991 maximum likelihood-based cointegration tests, did not find any cointegration relationship between the African markets and the developed ones, and similar results were obtained in Adjasi and Biekpe (2006), based on a dynamic Vector AutoRegressive regression (VAR) model. In another recent study, Mensah and Alagidede (2017) investigated four African markets (South Africa, Nigeria, Egypt and Kenya) using copulas, and documented that dependence is time varying and weak for most African markets except South Africa. Thus with the exception of this country, the rest of the markets examined in Africa (Kenya, Nigeria and Egypt) seem to be immune to risk spillovers from advanced markets. On the other hand, using linear dependence measures is at odds with the widely acknowledged fact that return distributions are non-normal. It is therefore essential to assess the dependence between African stock markets and other international markets with more accurate measures of dependence. In view of that, this paper investigates linkages between nine selected African markets: South Africa, Egypt, Morocco, Kenya, Nigeria, Tunisia, Ghana, Mauritius, and the Ivory Coast and six selected global markets (US, UK, Japan, China, Australia and Europe) using updated techniques based on the concepts of fractional integration and cointegration. These methods might be more appropriate than the standard ones (unit roots and cointegration tests in the classical sense) given that they allow for a much greater degree of flexibility in their dynamic specification, allowing for fractional degrees of differentiation.

2. Equity market integration: a methodological note

Financial integration defined as the extent to which financial markets are connected (Kenen, 1976). The empirical literature on stock market integration is very comprehensive and from a methodological point of view, several studies have employed different econometric approaches, ranging from but not limited to structural models and time series approaches such as volatility models, cointegration and principal components, with the models yielding mixed results. Taylor and Tonks (1989), employing the two-step Engle and Granger's (1987) approach, found that the UK stock market was cointegrated with those in the US, Japan, the Netherlands and West Germany. Jeon and Von Furstenberg (1990) employed a VAR approach, and established the existence of an increase in cross-border cointegration from 1987 onwards. Bekaert and Harvey (1995) used a conditional regime-switching model, and found that the degree of integration in emerging countries may be independent of the extent of investment restrictions. Antoniou et al. (2007) established that the UK stock market is more integrated with Europe in terms of both the aggregate and stock markets sectors after extracting the time varying conditional correlations between the US, the UK and European equity markets. Owing to the drawbacks of linear correlation models, multivariate GARCH models emerged as an approach to modelling time-varying stock dependence. Several studies adopted this approach to examine the stock markets interdependence including, among others, Syllignakis and Kouretas (2011), Gjika and Horvath (2013), Baumöhl and Lyócsa (2014) and Kundu and Sarkar (2016). However, one major limitation of the multivariate GARCH approach is the assumption that return innovations are characterized by a symmetric multivariate normal or Student-t distribution (Patton, 2006; Garcia and Tsafack, 2011). Evidently, this assumption seems to be at odds with the empirical evidence; the distribution of financial returns possesses heavier tails than those of the normal distribution and dependence between stocks returns are usually nonlinear and asymmetric (Embrechts et al., 2002). Against this background, researchers have resorted to a relatively new approach, copula, to model the dependence between stock returns. Hence, Mensah and Alagidede (2017) employed copulas to examine the relationship between Africa's emerging markets and the advanced markets. They found that with the exception of South Africa, the African markets examined were immune to risk spillover from the advanced markets. In another study, Alagidede (2009), using maximum likelihood and based on cointegration, showed that there exists weak stochastic trends between African markets and world markets. Still from a methodological standpoint and with reference to studies on markets integration, Caporale et al. (2016) adopted a fractional cointegration approach to examine the linkages between US and Europe stock markets. Wong et al. (2004) used the fractional cointegration approach to investigate the long run relationship between Indian stock markets and the major developed markets of which they concluded that the stock market of India is integrated with mature markets and is sensitive to the dynamics in these markets in the long run. Yi et al. (2009), using fractionally integrated VECM, also investigated the long run cointegration relations binding China's stock market and its neighbouring Hong Kong market and the US market from which they documented strong cointegration between the Chinese and Hong Kong stock markets. In addition, Assaf (2007) documented significant long memory in the

Middle East and North African (MENA) equity markets using a fractional integration approach. In this paper, we adopt a similar approach, though with some slight modifications with respect to the methodology used. Though fractional integration has been widely used in the analysis of stock markets, to the best of our knowledge there are no studies examining the linkage among markets with fractional cointegration techniques, especially in the context of the African continent.

3. Fractional integration and cointegration

Given a covariance stationary process $\{u_t, t=0, \pm 1, \ldots\}$ we say that it displays the property of short memory or I(0) behaviour if the infinite sum of its autocovariances is finite. That is, assuming that $\gamma_j = \mathrm{Cov}(u_t, u_{t+j})$,

$$\lim_{T\to\infty} \sum_{j=-T}^{T} |\gamma_j| < \infty. \tag{1}$$

Now, assuming that x_t has an absolutely continuous spectral distribution function, with a spectral density function given by:

$$f(\lambda) = \frac{1}{2\pi} \sum_{j=-\infty}^{\infty} \gamma_j \cos \lambda_j, \quad -\pi < \varepsilon \le \pi, \tag{2}$$

another definition of short memory states that the spectral density must be positive and bounded at all frequencies, i.e.,

$$0 < f(\lambda) < \infty. \quad -\pi < \lambda \le \pi. \tag{3}$$

Within this class of I(0) models, we can include the white noise and the stationary ARMA-type of models. These are characterized by the dependence between the observations far away in time being short and disappearing relatively quickly as the distance between the observations is increased.[1]

On the other hand, a process is said to be long memory if the infinite sum of the autocovariances is infinite, i.e.,

$$\lim_{T\to\infty} \sum_{j=-T}^{T} |\gamma_j| = \infty \tag{4}$$

or, alternatively, in the frequency domain if the spectral density function is unbounded at some frequency λ in the interval $[0, \pi]$, i.e.,

$$f(\lambda) \to \infty, \quad \text{as} \quad \lambda \to 0^+. \tag{5}$$

(see McLeod and Hipel, 1978). In this context, one of the most popular models satisfying the above two properties is the one based on fractional integration. Thus, a process x_t is said to be (fractionally) integrated of order d, and denoted by I(d) if it can be represented as:

$$(1 - L)^d x_t = u_t, \quad t = 0, \pm 1, \ldots, \tag{6}$$

with $x_t = 0$ for $t \le 0$, and $d > 0$, where L is the lag-operator ($Lx_t = x_{t-1}$) and u_t is I(0) as defined above. Note that for any real value d, the polynomial in the left hand side of Eq. (6) can be expressed as

$$(1 - L)^d = \sum_{j=0}^{\infty} \psi_j L^j = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j$$
$$= 1 - dL + \frac{d(d-1)}{2} L^2 - \ldots, \tag{7}$$

and thus

$$(1 - L)^d x_t = x_t - d\, x_{t-1} + \frac{d(d-1)}{2} x_{t-2} - \cdots. \tag{8}$$

In this context, d plays a crucial role as an indicator of the degree of dependence in the series. Thus, the higher the value of d is, the higher the level of association is between the observations. Also, if $d > 0$, x_t belongs to the "long memory" category mentioned above. These processes, originally proposed in the 80s by Granger (1980, 1981), Granger and Joyeux (1980) and Hosking (1981), gained popularity fifteen years later (Crato and Rothman, 1994; Baillie, 1996; Gil-Alana and Robinson, 1997, 2001; Michelacci and Zaffaroni, 2000; etc.) and their use has been generalized in the analysis of macroeconomic and financial data during the last ten years (see, e.g., Mayoral, 2006; Baillie et al., 2007; Christensen et al., 2010; Martins and Rodrigues, 2012; Gil-Alana and Moreno, 2012; Hassler et al., 2014; Cavaliere et al., 2015; Abbritti et al., 2016; etc.). Note that by estimating d rather than imposing it as a fixed integer value, we allow for a wide range of alternatives to be tested: stationarity and short memory, i.e. $d = 0$; stationarity with long memory ($0 < d < 0.5$); nonstationarity though mean reversion ($0.5 \le d < 1$); unit roots ($d = 1$), and lack of mean reversion behaviour ($d \le 1$).[2]

The natural generalization to the multivariate case of the concept of fractional integration is the one of fractional cointegration. In a bivariate context, given two real numbers d, b, the components of the vector z_t are said to be cointegrated of order d, b, denoted $z_t \sim CI(d, b)$ if:

(i) all the components of z_t are I(d), and
(ii) there exists a vector $\alpha \neq 0$ such that $s_t = \alpha' z_t \sim I(\gamma) = I(d - b)$, $b > 0$.

Here, α and s_t are called the cointegrating vector and error respectively. In our empirical application we conduct the following strategy: we first estimate individually the orders of integration of the series using both Whittle and maximum likelihood parametric methods (Dahlhaus, 1989; Sowell, 1992; Robinson, 1994; Beran, 1995) along with semiparametric approaches based on "local" Whittle methods (Robinson, 1995a; Velasco, 1999a; Abadir et al., 2007) and log-periodogram-type of estimators (Robinson, 1995b; Kim and Phillips, 1999, 2006;

[1] In the frequency domain, these processes are characterized because the spectral density function is positive and bounded at all frequencies.

[2] Gil-Alana and Hualde (2009) present an updated review of fractional integration (and cointegration) and its applications in economic time series.

Velasco, 1999b).[3] Next we test the homogeneity of the orders of integration in the bivariate representation of the system (i.e., H_o: $d_x = d_y$), where d_x and d_y refer now to the orders of integration of the two individual series. Here we use an adaptation of Robinson and Yajima's (2002) statistic to log-periodogram estimation.[4] In the final step, we perform a Hausman test for no cointegration of Marinucci and Robinson (2001) comparing the estimate $\hat{d}_x$ of d_x with the more efficient bivariate one of Robinson (1995b), which uses the information that $d_x = d_y = d_*$. Marinucci and Robinson (2001) show that:

$$H_{ix} = 8m\left(\hat{d}_* - \hat{d}_i\right)^2 \rightarrow_d \chi_1^2 \quad \text{as} \quad \frac{1}{m} + \frac{m}{T} \rightarrow 0, \qquad (9)$$

where T refers to the sample size and $m < [T/2]$ is the bandwidth number going to infinity, albeit at a slower rate than T; $\hat{d}_i$, i = 1 and 2 are the univariate estimates of the parent series, and $\hat{d}_*$ is a restricted estimate obtained in the bivariate context under the assumption that $d_x = d_y$. In particular,

$$\hat{d}_* = -\frac{\sum_{j=1}^{s^*} 1_2' \hat{\Omega}^{-1} Y_j v_j}{2\, 1_2' \hat{\Omega}^{-1} 1_2 \sum_{j=1}^{s^*} v_j^2}, \qquad (10)$$

where 1_2 indicates a (2×1) vector of 1s, $\hat{\Omega}$ is the estimated variance covariance matrix, and with $Y_j = [\log I_{xx}(\lambda_j), \log I_{yy}(\lambda_j)]^T$, with $\lambda_j = 2\pi j/T$, j = 1,2, ...T, as the Fourier frequencies, and where for arbitrary sequences ξ_t and ζ_t we define the discrete Fourier transform and the (cross)-periodogram:

$$w_\xi(\lambda) = \frac{1}{\sqrt{2\pi T}} \sum_{t=1}^{T} \xi_t e^{i\lambda t}, \quad I_{\xi\zeta}(\lambda) = w_\xi(\lambda) w_\zeta^T(-\lambda), \quad I_\xi(\lambda) = I_{\xi\xi}(\lambda)$$

and $v_j = \log j - \frac{1}{s}\sum_{j=1}^{s} \log j$. The limiting distribution above is presented heuristically, but the authors argue that it seems sufficiently convincing for the test to warrant serious consideration.

4. Data and empirical results

4.1. The data

Our dataset contains time series of equity price indices, with a weekly frequency, for a sample of nine countries with stock markets in Africa and six countries with developed stock markets. Our sample dataset from Africa includes; South Africa, Nigeria, Egypt, Kenya, Ghana, Tunisia, Morocco, Mauritius and the Ivory Coast, while the selected global developed markets include the United States, the United Kingdom, Japan, China and Europe. The Morgan Stanley Capital International (MSCI) stock market indices was obtained from DataStream International for the following countries; South Africa, Nigeria, Egypt, Kenya, Ghana, Tunisia, Morocco, Mauritius, the Ivory Coast, the United Kingdom and Australia. For the United States, Japan, China and Europe we obtained S&P 500, NIKKEI 225, SHANGHAI SE and EURO STOXX 50/STOXX 200 respectively. All returns series employed in the study are in USD. Our study period ranges from 1 January 1990 to 23 June 2018.

4.2. Univariate work: fractional integration

To begin with, we employed the conventional unit root testing methods of ADF (Dickey and Fuller, 1979); PP (Phillips and Perron, 1988); KPSS (Kwiatkowski et al., 1992); ERS (Elliot et al., 1996) and NP (Ng and Perron, 2001) to discriminate between I(0) and I(1) behaviour. We obtained strong evidence of unit roots, following the above standard unit roots tests, which have been found to have very low power under certain types of alternatives such as structural breaks, non-linearities and fractional integration. In particular, for the latter case, Diebold and Rudebusch (1991), Hassler and Wolters (1994) and Lee and Schmidt (1996) showed that if a series is integrated of order d and the value of d is different from 0 or 1, then the standard methods such as those applied above may not be suitable.

In view of the above, we begin this empirical section by estimating the fractional differencing parameter d in the model outlined below;

$$y_t = \beta_0 + \beta_1 t + x_t, \; (1-L)^d x_t = u_t, \; t = 1, 2, ..., \qquad (11)$$

where y_t is the observed series, β_0 and β_1 are the coefficients corresponding to an intercept and a linear time trend and $x_t$is assumed to be $I(d)$, where d can take any real value. Therefore, the error term, u_t, is $I(0)$ and is assumed in turn to be white noise and autocorrelated.

We test the null hypothesis:

$$H_o : d = d_o \qquad (12)$$

for given real values $d_o$in the model given by Eq. (11) above. Table 1 presents the results obtained for the fractional differencing parameter d under no autocorrelation for the log-transformed data, along with their corresponding 95% confidence intervals corresponding to the non-rejection values of d_o testing (11) in (12), in the three cases of no deterministic terms (i.e., $\beta_0 = \beta_1 = 0$ a priori in the undifferenced Eq. (11)), including only a constant (β_0 unknown, and $\beta_1 = 0$ a priori), and finally the case of a constant with a linear time trend (β_0 and β_1 unknown). Following Eq. (11), if u_t is assumed to be uncorrelated, then the estimates of d are all about 1. From Section A in Table 1, which reports findings for the selected African markets, an intercept seems to be sufficient to describe the deterministic components of the series except for the Ivory Coast where an

[3] We do not report the results based on all these methods since, within its class, the results were very similar across all them.

[4] The functional form of the test statistic can be found in Gil-Alana and Hualde (2009).

Table 1
Estimates of d (and 95% confidence bands) under no autocorrelation.

Country	No regressors	An intercept	A linear trend
Section A: African markets			
Egypt	1.00 (0.97, 1.05)	1.05 (1.02, 1.09)	1.00 (0.97, 1.05)
Ghana	1.01 (0.98, 1.06)	**1.15 (1.12, 1.19)**	1.15 (1.12, 1.19)
Ivory Coast	0.99 (0.96, 1.05)	1.02 (0.99, 1.06)	**1.02 (0.99, 1.06)**
Kenya	0.99 (0.96, 1.04)	**1.06 (1.02, 1.10)**	1.06 (1.02, 1.10)
Mauritius	1.00 (0.96, 1.04)	**1.08 (1.05, 1.12)**	1.08 (1.05, 1.12)
Morocco	1.00 (0.96, 1.04)	**1.04 (1.01, 1.08)**	1.04 (1.01, 1.08)
Nigeria	1.01 (0.97, 1.06)	**1.12 (1.08, 1.16)**	1.12 (1.08, 1.16)
South Africa	1.00 (0.96, 1.04)	**0.94 (0.90, 0.98)**	0.94 (0.90, 0.98)
Tunisia	1.00 (0.96, 1.04)	**1.04 (1.00, 1.08)**	1.04 (1.00, 1.08)
Section B: selected global markets			
UK	1.00 (0.96, 1.04)	**0.92 (0.89, 0.95)**	0.92 (0.89, 0.95)
Australia	1.00 (0.96, 1.05)	**0.95 (0.92, 0.99)**	0.95 (0.92, 0.99)
Japan	1.00 (0.96, 1.04)	**0.96 (0.93, 1.00)**	0.96 (0.93, 1.00)
China	1.01 (0.97, 1.05)	**1.03 (0.99, 1.08)**	1.03 (0.99, 1.08)
USA	1.00 (0.96, 1.04)	0.93 (0.89, 0.96)	**0.93 (0.89, 0.96)**
Europe (STOXX 200	1.00 (0.96, 1.04)	**0.95 (0.91, 0.98)**	0.95 (0.91, 0.98)
Europe (STOXX 50)	1.00 (0.96, 1.04)	**0.96 (0.92, 1.00)**	0.96 (0.93, 1.00)

In bold, the significant models according to the deterministic terms.

intercept and a linear time trend seem to be required. For Section B, which documents findings for the selected global markets, with the exception of the US, again an intercept seems to be sufficient to explain the deterministic component of the series (UK, Japan, Australia, China and Europe). If we focus on the estimated values of d we see that for the African countries most of the estimates are slightly above 1. Only for South Africa do we get an estimate statistically smaller than 1 (d = 0.94)[5]; for the Ivory Coast and Tunisia the unit root null hypothesis (i.e. d = 1) cannot be rejected, while this hypothesis is rejected in favour of d > 1 in the remaining cases. However, for the developed economies, the estimates of d are somewhat smaller, being around 1 (in the case of Japan, 0.96; China, 1.03) and Europe (Stoxx 50, 0.96) or even smaller than 1 (UK, 0.92; Australia, 0.95; USA, 0.93; and Europe Stoxx 300, 0.95).

The results, allowing for autocorrelated errors, are reported in Table 2. Starting with the African countries, we obtain evidence of I(d) behaviour with d > 1 for Egypt (1.09), Ghana (1.19), Kenya (1.08), Mauritius (1.17), Morocco (1.06), Nigeria (1.13) and Tunisia (1.07), while the unit root null hypothesis cannot be rejected in the cases of the Ivory Coast (1.03) and South Africa (0.96). For the developed markets, the values of d range between 0.97 (Australia) and 1.05 (China), and the unit root null cannot be rejected in any single case.

We also employ a semi-parametric approach based on a "local" Whittle estimate that degenerates to zero (Robinson, 1995a).[6] In Table 3 we report the values of d for a selected number of bandwidth parameters from m = 25, . . . (1), . . ., 34. Taking into consideration where m is approximately $(T)^{0.5}$, i.e., 29 and 30, following results in Table 3, we reject the null hypothesis of unit root for Egypt, Ghana, Morocco, Nigeria and the United States. However, we fail to reject this null hypothesis of a unit root in favour of d > 1 for the Ivory Coast, Kenya, Mauritius, South Africa, Tunisia, UK, Australia, Japan, China and Europe zone.

The results obtained so far with both parametric and semi-parametric techniques, are consistent with prior studies that document evidence of unit roots in stock indices in most developed stock markets (e.g., Narayan 2005, 2006; Narayan and Smyth, 2004; Lean and Smyth, 2007; Murthy et al., 2011). On the whole, with reference to the univariate results presented in Tables 1–3, we can conclude by saying that there exists a high degree of persistence with orders of integration about 1 or higher than 1, which implies that shocks to these stock markets have significant permanent effects.

4.3. *Multivariate work: fractional cointegration*

In order to establish the linkages between African emerging equity markets and the world, we carried out bivariate tests to investigate first the hypothesis of equality in the orders of integration between vis-a-vis markets. The results are reported in Table 4. We find evidence against the null hypothesis of equal orders of integration in a number of cases, including Egypt and Australia; Ghana and Mauritius with respect to all the developed markets; Kenya with Australia; and Nigeria versus the UK, Australia, Japan and the US. On the other hand, for the Ivory Coast, Morocco, South Africa and Tunisia, we cannot reject the null of equal orders of integration in any single case.

For all these bivariate cases where we cannot reject the null of equal orders of integration, we test for cointegration, testing the null hypothesis of no cointegration against the alternative of fractional integration (see, Gil-Alana, 2003). First, we run OLS regressions on each of the African stock markets

[5] Note that for this series, the confidence interval excludes the value of 1, and the interval is strictly below the unity.

[6] The results here were based on the first differenced data, adding then the value of 1 to get the proper estimates of d.

Table 2
Estimates of d (and 95% confidence bands) under autocorrelation.

Country	No regressors	An intercept	A linear trend
Section A: African markets			
Egypt	1.01 (0.93, 1.06)	1.09 (1.04, 1.06)	1.09 (1.04, 1.16)
Ghana	1.03 (0.97, 1.10)	**1.19 (1.13, 1.26)**	1.19 (1.13, 1.26)
Ivory Coast	0.98 (0.93, 1.06)	1.03 (0.97, 1.10)	**1.03 (0.97, 1.10)**
Kenya	1.00 (0.94 1.06)	**1.08 (1.02, 1.13)**	1.07 (1.02, 1.13)
Mauritius	0.99 (0.94, 1.07)	**1.17 (1.11, 1.24)**	1.17 (1.11, 1.24)
Morocco	0.99 (0.94, 1.06)	**1.06 (1.01, 1.12)**	1.06 (1.01, 1.12)
Nigeria	1.00 (0.95, 1.07)	**1.13 (1.07, 1.20)**	1.13 (1.07, 1.20)
South Africa	0.97 (0.92, 1.05)	**0.96 (0.90, 1.03)**	0.96 (0.90, 1.07)
Tunisia	1.00 (0.94, 1.08)	**1.07 (1.02, 1.14)**	1.07(1.02, 1.14)
Section B: selected global markets			
UK (MSCI)	0.99 (0.93, 1.07)	**1.01 (0.95, 1.08)**	1.01 (0.95, 1.08)
Australia (MSCI)	0.99 (0.93 1.06)	**0.97 (0.92, 1.04)**	0.97 (0.92, 1.04)
Japan (NIKKEI 225)	1.00 (0.94, 1.06)	**1.02 (0.95, 1.08)**	1.02 (0.95, 1.08)
China (SHANGAI SE)	1.00 (0.94, 1.07)	**1.05 (0.98, 1.11)**	1.05 (0.98, 1.11)
USA (S&P 500)	0.99 (0.93, 1.06)	1.00 (0.94, 1.07)	**1.00 (0.94, 1.07)**
Europe (STOXX 200)	0.99 (0.93, 1.06)	**1.03 (0.96, 1.10)**	1.03 (0.96, 1.10)
Europe (STOXX 50)	0.99 (0.93, 1.07)	**1.03 (0.96, 1.09)**	1.03 (0.96, 1.09)

In bold, the significant models according to the deterministic terms.

Table 3
Estimates of d based on a semiparametric “local” Whittle method.

Country	25	26	27	28	29	30	31	32	33	34
Section A: African markets										
Egypt	1.153	1.159	1.174[a]	1.194[a]	1.194[a]	1.215[a]	1.215[a]	1.232[a]	1.214[a]	1.215[a]
Ghana	1.461[a]	1.342[a]	1.317[a]	1.350[a]	1.352[a]	1.374[a]	1.348[a]	1.363[a]	1.378[a]	1.292[a]
Ivory Coast	1.108	1.098	1.115	1.113	1.128	1.110	1.109	1.118	1.105	1.117
Kenya	1.131	1.126	1.145	1.148	1.157	1.113	1.135	1.128	1.105	1.164[a]
Mauritius	0.921	0.953	0.981	1.013	1.028	1.061	1.086	1.117	1.142	1.164[a]
Morocco	1.196[a]	1.192[a]	1.202[a]	1.210[a]	1.227[a]	1.230[a]	1.237[a]	1.257[a]	1.126	1.250[a]
Nigeria	1.263[a]	1.207[a]	1.192[a]	1.129	1.221[a]	1.117	1.093	1.106	1.271[a]	1.121
South Africa	0.942	0.922	0.919	0.893	0.883	0.892	0.890	0.907	0.909	0.924
Tunisia	1.099	1.119	1.137	1.138	1.128	1.103	1.113	1.119	1.092	1.091
Section B: selected global markets										
UK (MSCI)	1.053	1.030	1.062	1.088	1.096	1.109	1.086	1.113	1.133	1.125
Australia (MSCI)	0.921	0.907	0.937	0.968	0.974	0.967	0.952	0.977	0.987	0.982
Japan (NIKKEI 225)	1.027	1.012	1.032	1.024	1.021	1.031	1.032	1.043	1.051	1.058
China (SHANGAI SE)	1.052	1.058	1.083	1.073	1.091	1.030	1.055	1.075	1.105	1.094
USA (S&P 500)	1.145	1.126	1.153	1.139	1.146	1.151[a]	1.116	1.136	1.148	1.103
Europe (STOXX 200)	1.053	1.040	1.068	1.093	1.091	1.108	1.077	1.102	1.123	1.107
Europe (STOXX 50)	1.034	1.027	1.053	1.077	1.071	1.089	1.059	1.082	1.103	1.108
Lower I(1)	0.835	0.838	0.841	0.844	0.847	0.849	0.852	0.854	0.856	0.859
Upper I(1)	1.164	1.161	1.158	1.155	1.152	1.150	1.147	1.145	1.143	1.141

[a] Evidence of I(d) with $d > 1$ at the 5% level.

Table 4
Homogenity tests for the orders of integration.

Country	UK MSCI	Australia MSCI	Japan NIKKEI225	China SHANG.SE	US SP500	Europe STOXX200	Europe STOXX50
Egypt	1.34	2.02[a]	1.17	0.67	1.51	1.01	1.01
Ghana	3.03[a]	3.70[a]	2.86[a]	2.35[a]	3.20[a]	2.69[a]	2.69[a]
Ivory Coast	0.33	1.01	0.16	−0.33	0.50	0.01	0.00
Kenya	1.18	1.85[a]	1.01	0.50	1.34	0.84	0.85
Mauritius	2.35[a]	3.27[a]	2.52[a]	2.02[a]	2.86[a]	2.35[a]	2.35[a]
Morocco	0.84	1.51	0.67	0.16	1.01	0.50	0.51
Nigeria	2.02[a]	2.69[a]	1.85[a]	1.34	2.19[a]	1.67	1.68
South Africa	−0.84	−0.16	−1.01	−1.51	−0.67	−1.18	−1.18
Tunisia	1.01	1.68	0.84	0.33	1.18	0.67	0.68

[a] Evidence against the null hypothesis of equal orders of integration at the 5% level.

Table 5
Estimated values of d on the estimated co-integrating residuals.

Country	Stock market	No autocorrelation	Autocorrelation
Egypt	UK (MSCI)	**0.87 (0.84, 0.90)**	0.94 (0.87, 1.00)
	Australia (MSCI) AUS	**0.90 (0.87, 0.94)**	0.94 (0.88, 1.00)
	Japan (NIKKEI 225)	1.03 (1.00, 1.07)	1.07 (1.01, 1.14)
	China (SHANGHAI SE)	1.04 (1.01, 1.08)	1.09 (1.04, 1.15)
	US (S&P 500)	0.95 (0.91, 0.98)	1.03 (0.98, 1.09)
	Europe zone (STOXX 200)	**0.89 (0.86, 0.91)**	0.96 (0.90, 1.00)
	Europe zone (STOXX 50)	**0.90 (0.86, 0.93)**	0.95 (0.88, 1.01)
Ghana	UK (MSCI)	1.09 (1.06, 1.12)	1.16 (1.10, 1.19)
	Australia (MSCI) AUS	1.10 (1.07, 1.13)	1.15 (1.09, 1.18)
	Japan (NIKKEI 225)	1.08 (1.05, 1.11)	1.14 (1.09, 1.17)
	China (SHANGHAI SE)	1.11 (1.08, 1.14)	1.16 (1.10, 1.19)
	US (S&P 500)	1.11 (1.08, 1.14)	1.18 (1.12, 1.22)
	Europe zone (STOXX 200)	1.10 (1.07, 1.14)	1.17 (1.12, 1.21)
	Europe zone (STOXX 50)	1.10 (1.07, 1.14)	1.15 (1.12, 1.19)
Ivory Coast	UK (MSCI)	**0.91 (0.88, 0.95)**	0.98 (0.92, 1.03)
	Australia (MSCI) AUS	**0.91 (0.88, 0.95)**	**0.92 (0.86, 0.98)**
	Japan (NIKKEI 225)	1.01 (0.98, 1.05)	1.01 (0.97, 1.04)
	China (SHANGHAI SE)	0.99 (0.96, 1.03)	1.00 (0.96, 1.04)
	US (S&P 500)	**0.93 (0.90, 0.97)**	0.97 (0.93, 1.00)
	Europe zone (STOXX 200)	**0.92 (0.89, 0.95)**	0.98 (0.92, 1.03)
	Europe zone (STOXX 50)	**0.93 (0.90, 0.96)**	0.96 (0.91, 1.02)
Kenya	UK (MSCI)	**0.88 (0.85, 0.92)**	**0.89 (0.86, 0.92)**
	Australia (MSCI) AUS	**0.90 (0.86, 0.94)**	**0.90 (0.86, 0.93)**
	Japan (NIKKEI 225)	1.05 (1.01, 1.09)	1.05 (1.00, 1.09)
	China (SHANGHAI SE)	1.03 (1.00, 1.07)	1.03 (1.00, 1.07)
	US (S&P 500)	**0.92 (0.88, 0.96)**	**0.93 (0.90, 0.97)**
	Europe zone (STOXX 200)	**0.90 (0.85, 0.94)**	**0.91 (0.88, 0.95)**
	Europe zone (STOXX 50)	**0.92 (0.88, 0.96)**	**0.92 (0.89, 0.96)**
Mauritius	UK (MSCI)	**0.88 (0.85, 0.90)**	0.95 (0.90, 1.00)
	Australia (MSCI) AUS	**0.89 (0.85, 0.93)**	**0.89 (0.85, 0.94)**
	Japan (NIKKEI 225)	1.06 (1.03, 1.10)	1.06 (1.03, 1.10)
	China (SHANGHAI SE)	1.03 (1.00, 1.08)	1.03 (1.00, 1.07)
	US (S&P 500)	**0.93 (0.91, 0.96)**	**0.93 (0.90, 0.96)**
	Europe zone (STOXX 200)	**0.89 (0.85, 0.92)**	**0.89 (0.86, 0.92)**
	Europe zone (STOXX 50)	**0.90 (0.86, 0.93)**	**0.90 (0.86, 0.93)**
Morocco	UK (MSCI)	**0.93 (0.90, 0.97)**	0.95 (0.97, 0.99)
	Australia (MSCI) AUS	0.99 (0.95, 1.04)	0.96 (0.99, 1.02)
	Japan (NIKKEI 225)	1.02 (0.99, 1.06)	1.04 (0.98, 1.10)
	China (SHANGHAI SE)	1.02 (0.98, 1.05)	1.04 (0.99, 1.11)
	US (S&P 500)	1.01 (0.98, 1.04)	1.01 (0.96, 1.05)
	Europe zone (STOXX 200)	**0.95 (0.91, 0.99)**	0.95 (0.90, 1.01)
	Europe zone (STOXX 50)	**0.95 (0.92, 0.98)**	0.97 (0.91, 1.02)
Nigeria	UK (MSCI)	**0.93 (0.90, 0.97)**	**1.02 (0.96, 1.07)**
	Australia (MSCI) AUS	0.97 (0.93, 1.01)	0.99 (0.93, 1.04)
	Japan (NIKKEI 225)	1.08 (1.04, 1.12)	1.10 (1.03, 1.16)
	China (SHANGHAI SE)	1.03 (1.00, 1.07)	1.07 (1.01, 1.13)
	US (S&P 500)	1.05 (1.02, 1.09)	1.07 (1.04, 1.12)
	Europe zone (STOXX 200)	**0.96 (0.92, 0.99)**	1.04 (0.98, 1.10)
	Europe zone (STOXX 50)	**0.97 (0.93, 1.00)**	1.04 (0.98, 1.11)
South Africa	UK (MSCI)	0.97 (0.93, 1.00)	0.98 (0.94, 1.02)
	Australia (MSCI) AUS	**0.90 (0.83, 0.96)**	0.92 (0.85, 1.01)
	Japan (NIKKEI 225)	**0.94 (0.90, 0.98)**	0.96 (0.91, 1.04)
	China (SHANGHAI SE)	**0.94 (0.90, 0.97)**	0.95 (0.90, 1.01)
	US (S&P 500)	**0.95 (0.93, 0.98)**	0.96 (0.91, 1.02)
	Europe zone (STOXX 200)	1.00 (0.96, 1.04)	1.00 (0.94, 1.04)
	Europe zone (STOXX 50)	0.99 (0.95, 1.03)	1.00 (0.94, 1.04)
Tunisia	UK (MSCI)	1.02 (0.98, 1.08)	1.03 (0.99, 1.10)
	Australia (MSCI) AUS	0.98 (0.93, 1.02)	1.01 (0.94, 1.08)
	Japan (NIKKEI 225)	1.04 (1.00, 1.08)	1.07 (1.00, 1.13)

Table 5 (*Continued*)

Country	Stock market	No autocorrelation	Autocorrelation
	China (SHANGHAI SE)	1.03 (0.98, 1.07)	1.04 (0.99, 1.10)
	US (S&P 500)	1.04 (0.99, 1.08)	1.04 (0.98, 1.09)
	Europe zone (STOXX 200)	1.01 (0.98, 1.05)	1.04 (0.97, 1.10)
	Europe zone (STOXX 50)	1.01 (0.97, 1.05)	1.04 (0.97, 1.10)

Table 6
Hausman tests of no cointegration against fractional cointegration.

	UK MSCI	Australia MSCI	Japan NIKKEI225	China SHANG.	US SP500	Europe STOXX200	Europe STOXX5
Egypt	H_1: 6.52	–	H_1: 0.48	H_1: 0.33	H_1: 2.64	H_1: 5.34	H_1: 4.86
	H_2: 2.64		H_2: 0.01	H_2: 0.01	H_2: 0.33	H_2: 2.64	H_2: 2.27
	d = 0.87		d = 1.03	d = 1.04	d = 0.95	d = 0.89	d = 0.90
Ghana	–	–	–	–	–	–	–
Ivory Coast	H_1: 1.94	H_1: 1.94	H_1: 0.12	H_1: 0.21	H_1: 1.34	H_1: 1.63	H_1: 1.34
	H_2: 1.34	H_2: 0.48	H_2: 0.05	H_2: 0.48	H_2: 0.66	H_2: 1.63	H_2: 1.34
	d = 0.91	d = 0.91	d = 1.00	d = 0.99	d = 0.93	d = 0.92	d = 0.93
Kenya	H_1: 5.39	–	H_1: 0.12	H_1: 0.33	H_1: 3.45	H_1: 4.36	H_1: 3.45
	H_2: 2.27		H_2: 0.12	H_2: 0.05	H_2: 0.86	H_2: 2.27	H_2: 1.63
	d = 0.88		d = 1.05	d = 1.03	d = 0.92	d = 0.90	d = 0.92
Mauritius	–	–	–	–	–	–	–
Morocco	H_1: 2.27	H_1: 1.34	H_1: 0.21	H_1: 0.21	H_1: 0.33	H_1: 1.63	H_1: 1.63
	H_2: 0.86	H_2: 0.01	H_2: 0.00	H_2: 0.12	H_2: 0.01	H_2: 0.86	H_2: 0.86
	d = 0.93	d = 0.96	d = 1.02	d = 1.02	d = 1.01	d = 0.95	d = 0.95
Nigeria	–	–	–		–		
South Africa	H_1: 0.01	H_1: 0.48	H_1: 0.05	H_1: 0.05	H_1: 0.01	H_1: 0.21	H_1: 0.21
	H_2: 0.21	H_2: 0.66	H_2: 0.86	H_2: 1.63	H_2: 0.33	H_2: 0.12	H_2: 0.12
	d = 0.97	d = 0.90	d = 0.94	d = 0.94	d = 0.95	d = 1.00	d = 1.00
Tunisia	H_1: 1.09	H_1: 0.12	H_1: 0.21	H_1: 0.21	H_1: 0.12	H_1: 0.48	H_1: 0.48
	H_2: 0.12	H_2: 0.66	H_2: 0.01	H_2: 0.05	H_2: 0.21	H_2: 0.05	H_2: 0.05
	d = 1.02	d = 0.98	d = 1.04	d = 1.03	d = 1.04	d = 1.01	d = 1.01

Evidence against the null hypothesis of equal orders of integration at the 5% level.

against the developed economics and the results in terms of the estimated values of d, under both uncorrelated and autocorrelated errors, are reported in Table 5. The values where evidence of mean reversion (i.e., $d < 1$) is found are marked in the table in bold (and with an asterisk). If the errors are uncorrelated, we observe that mean reversion takes place in the cases of Egypt against the UK, Australia and the Europe Zone; for Ivory Coast, Kenya and Mauritius, it is found against the same stocks along with those of the USA; for Morocco and Nigeria mean reversion is found in the residuals against the UK, and Europe Zone; and for South Africa in case of Australia, Japan, China and the USA. With autocorrelation, there are fewer cases. In particular, for the Ivory Coast against Australia; for Kenya against the UK, Australia, the USA and the Europe Zone; for Mauritius against Australia, the USA and the Europe Zone, and finally for Morocco against the UK.

However, this evidence of mean reversion does not necessarily guarantee the existence of cointegration. The results of the proper tests, based on Marinucci and Robinson (2001), are reported in Table 6. Evidence of cointegration is found for Egypt and Kenya against the UK and the Europe Zone. The findings for Egypt and Kenya differs from the conclusion obtained by Mensah and Alagidede (2017). There are some other cases where partial evidence of cointegration is found, i.e., it is found only against one of the two hypothesis, though in general, in all cases, we observe that the degree of cointegration is very low, implying very long periods of convergence in the long run.

5. Summary and concluding comments

In this paper we have examined the stochastic properties of several stock markets throughout the world, and the bilateral linkages between African and selected global markets using weekly stock prices obtained from DataStream, for the time period 1 January 1990–23 June 2018. For this purpose, we have used a long memory approach based on fractionally integrated and cointegrated techniques. We adopted this approach to investigate the dependence between the selected markets since it allows for much richer dynamics than the classical models employed in prior studies that rely on integer degrees of differentiation and thus based on the I(0)/I(1) dichotomy.

The empirical results from the univariate analyses show that there exists a high degree of persistence in the series with orders of integration about 1 or higher than 1 which implies that shocks to these stock markets have significant permanent effects. Thus, for example, in the context of autocorrelated disturbances, the unit root null hypothesis is rejected in favour of d > 1 for Egypt (1.09), Ghana (1.19), Kenya (1.08), Mauritius (1.17), Morocco (1.06), Nigeria (1.13) and Tunisia (1.07), while the unit root null hypothesis cannot be rejected in the cases of the Ivory Coast (1.03) and South Africa (0.96). For the developed markets, the values of d range between 0.97 (Australia) and 1.05 (China), and the unit root null cannot be rejected in any single case.

Performing OLS regressions between the African and the developed markets we found residuals which are mean reverting in a number of cases, including the regressions of the Ivory Coast against Australia; for Kenya against the UK, Australia, the USA and the Europe Zone; for Mauritius against Australia, the USA and the Europe Zone, and finally for Morocco against the UK. However, testing properly the null of no cointegration against fractional cointegration by means of Robinson and Marinucci's (2001) approach, evidence of cointegration is only found for Egypt against the UK, the Europe Zone, and for Kenya against the UK and the Europe Zone and, though there are some other cases where partial evidence of cointegration might exist, generally we observe that the degree of cointegration is very low, implying very long periods of convergence in the long run.

The documented evidence has important implications for policy makers and markets participants. The presence of a very low degree of cointegration between some African emerging equity markets and the global stock markets shows that international investors holding African stocks in their portfolio will have potential gains. Morever, the low cointegration between these markets could serve as a conduit for policy makers to draw more portfolio investments to the continent in general. Apparently, they are still isolated from the core of major international markets, though many markets in Africa have undergone structural changes and developments. We can conclude that African markets need to deepen structural relations both regionally and with the rest of the world.

References

Abadir, K.M., Distaso, W., Giraitis, L., 2007. Nonstationarity-extended local Whittle estimation. J. Econom. 141, 1353–1384.

Abbritti, M., Gil-Alana, L.A., Lovcha, Y., Moreno, A., 2016. Term structure persistence. J. Financ. Econom. 14 (2), 331–352.

Adjasi, C.K.D., Biekpe, N.B., 2006. Stock market development and economic growth: the case of selected African countries. Afr. Dev. Rev. 18 (1), 144–161.

Alagidede, P., 2009. Are African stock markets integrated with the rest of the world? Afr. Finance J. 11 (1), 37–53.

Alagidede, P., Panagiotidis, T., Zhang, X., 2011. Causal relationship between stock prices and exchange rates. J. Int. Trade Econ. Dev. 20 (1), 67–86.

Allen, F., Otchere, I., Senbet, L.W., 2010. African financial systems: a review. Rev. Dev. Finance 1 (2), 79–113.

Antoniou, A., Pescetto, G., Stevens, I., 2007. Market-wide and sectoral integration: evidence from the UK, US and Europe. Managerial Finance 33 (3), 173–194.

Assaf, A., 2007. Fractional integration in the equity markets of MENA region. Appl. Financ. Econ. 17 (9), 709–723.

Baillie, R.T., 1996. Long memory processes and fractional integration in econometrics. J. Econom. 73, 5–59.

Baillie, R.T., Han, Y.W., Myers, R.J., Song, J., 2007. Long memory models for daily and high frequency commodity future returns. J. Future Mark. 27, 643–668.

Baumöhl, E., Lyócsa, S., 2014. Volatility and dynamic conditional correlations of worldwide emerging and frontier markets. Econ. Modell. 38, 175–183.

Bekaert, G., Harvey, C., 1995. Time varying world market integration. J. Finance 50 (2), 403–444.

Beran, J., 1995. Maximum likelihood estimation of the differencing parameter for invertible short and long memory ARIMA models. J. R. Stat. Soc. Ser. B 57, 659–672.

Bessler, D.A., Yang, J., 2003. The structure of interdependence in international stock markets. J. Int. Money Finance 22 (2), 261–287.

Caporale, G.M., Gil-Alana, L.A., Orlando, J.C., 2016. Linkages between the US and European stock markets: a fractional cointegration approach. Int. J. Finance Econ. 21 (2), 143–153.

Cavaliere, G., Nielsen, M.O., Taylor, A.M.R., 2015. Boostrap score tests for fractional integration in heteroskedastic ARFIMA models, with an application to price dynamics in commodity spot and future markets. J. Econom. 187, 557–579.

Chaudhuri, K., Wu, Y., 2003. Random walk versus breaking trend in stock prices: evidence from emerging markets. J. Bank. Finance 27 (4), 575–592.

Chen, G., Firth, M., Rui, O., 2002. Stock market linkages: evidence from Latin America. J. Bank. Finance 26 (6), 1113–1141.

Christensen, B.J., Nielsen, M.O., Zhu, J., 2010. Long memory in stock market volatility and the volatility in mean effect: the FIEGARCH-M model. J. Empirical Finance 17, 460–470.

Claessens, S., Forbes, K., 2001. International Financial Contagion. Springer Editorial.

Corhay, A., Rad, T., Urbain, J., 1993. Common stochastic trends in European stock markets. Econ. Lett. 42 (4), 385–390.

Crato, N., Rothman, P., 1994. Fractional integration analysis of long run behavoiur for US macroeconomic time series. Econ. Lett. 45 (3), 287–291.

Dahlhaus, R., 1989. Efficient parameter estimation for self-similar process. Ann. Stat. 17, 1749–1766.

Dekker, A., Sen, K., Young, M., 2001. Equity market in the Asia Pacific region: a comparison of the orthogonalized and generalized VAR approaches. Glob. Finance J. 12, 1–33.

De Santis, G., Gerard, B., 1997. International asset pricing and portfolio diversification with time-varying risk. J. Finance 52 (5), 1881–1912.

Dickey, D.A., Fuller, W.A., 1979. Distributions of the estimators for autoregressive time series with a unit root. J. Am. Stat. Assoc. 74 (366), 427–481.

Diebold, F.X., Rudebusch, G.D., 1991. On the power of Dickey–Fuller test against fractional alternatives. Econ. Lett. 35 (1), 155–160.

Donadelli, M., 2013. Global integration and emerging stock market excess returns. Macroecon. Finance Emerg. Mark. Econ. 6 (2), 244–279.

Driessen, J., Laeven, L., 2007. International portfolio diversification benefits: cross-country evidence from a local perspective. J. Bank. Finance 31 (6), 1693–1712.

Dumas, B., Solnik, B., 1995. The world price of foreign exchange risk. J. Finance 50 (2), 445–479.

Elliot, G., Rothenberg, T.J., Stock, J.H., 1996. Efficient tests for an autoregressive unit root. Econometrica 64, 813–836.

Embrechts, P., McNeil, A., Straumann, D., 2002. Correlation and dependence in risk management: properties and pitfalls. In: Dempster, M.A.H. (Ed.), Risk Management: Value at Risk and Beyond. Cambridge University Press, Cambridge, pp. 176–223.

Engle, R.F., Granger, C.W.J., 1987. Cointegration and error correction: representation, estimation and testing. Econometrica 55, 251–276.

Fratzscher, 2001. Financial market integration in Europe: on the effects of EMU on stock markets. Int. J. Finance Econ. 7 (3), 165–193.

Garcia, R., Tsafack, G., 2011. Dependence structure and extreme comovements in international equity and bond markets. J. Bank. Finance 35 (8), 1954–1970.

Gerrits, R., Yuce, A., 1999. Short- and long-term links among European and US stock markets. Appl. Financ. Econ. 9 (1), 1–9.

Gjika, D., Horvath, R., 2013. Stock market comovements in Central Europe: evidence from the asymmetric DCC model. Econ. Modell. 33, 55–64.

Gil-Alana, L.A., 2003. Testing of fractional cointegration in macroeconomic time series. Oxf. Bull. Econ. Stat. 65, 517–529.

Gil-Alana, L.A., Hualde, J., 2009. Fractional integration and cointegration: an overview with an empirical application. Palgrave Handb. Appl. Econom. 2, 434–472.
Gil-Alana, L.A., Moreno, A., 2012. Uncovering the US term premium: an alternative route. J. Bank. Finance 36 (4), 1181–1193.
Gil-Alana, L.A., Robinson, P.M., 1997. Testing of unit roots and other nonstationary hypotheses in macroeconomic time series. J. Econom. 80, 241–268.
Gil-Alana, L.A., Robinson, P.M., 2001. Testing seasonal fractional integration in the UK and Japanese consumption and income. J. Appl. Econom. 16, 95–114.
Granger, C.W.J., 1980. Long memory relationships and the aggregation of dynamic models. J. Econom. 14, 227–238.
Granger, C.W.J., 1981. Some properties of time series data and their use in econometric model specification. J. Econom. 16, 121–130.
Granger, C.W.J., Joyeux, R., 1980. An introduction to long memory time series models and fractional differencing. J. Time Ser. Anal. 1, 15–39.
Hassler, U., Rodrigues, P.M.M., Rubia, A., 2014. Persistence in the banking industry: fractional integration and breaks in memory. J. Empirical Finance 29, 95–112.
Hasslers, U., Wolters, J., 1994. On the power of unit root tests against fractional alternatives. Econ. Lett. 45 (1), 1–5.
Hosking, J.R.M., 1981. Fractional differencing. Biometrika 68, 165–176.
Huang, B., Yang, C., Hu, J., 2000. Causality and cointegration of stock markets among the united states, Japan and the South China growth triangle. Int. Rev. Financ. Anal. 9 (3), 281–297.
Jeon, B.N., Von Furstenberg, G.M., 1990. Growing international co-movement in stock price indexes. Q. Rev. Econ. Bus. 30 (3), 15–30.
Johansen, S., 1988. Statistical analysis of cointegration vectors. J. Econ. Dyn. Control 12 (2–3), 231–254.
Johansen, S., 1991. Estimation and hypothesis testing of cointegration vectors in Gaussian Vector Autoregressive models. Econometrica 59 (6), 1551–1580.
Kenen, P., 1976. International Trade and Finance. Cambridge University Press, London.
Kim, C., Phillips, P.C.B., 1999. Fully-Modified Estimation of Fractional Cointegration Models Mimeo. Yale University.
Kim, C., Phillips, P.C.B., 2006, Log periodogram regression: the nonstationary case, Cowles Foundation Discussion Papers, no. 1597.
Koedijk, K., Campbell, R., Kofman, P., 2002. Increased correlation in bear markets: a downside risk perspective. Financ. Anal. J. 58 (1), 87–94.
Kundu, S., Sarkar, N., 2016. Is the effect of risk on stock returns different in up and down markets? A multi-country study. Int. Econom. Rev. 8 (2), 53–71.
Kwan, A.C.C., Sim, A.B., Cotsomitis, J.A., 1995. The causal relationships between equity indices on world exchanges. Appl. Econ. 27 (1), 33–37.
Kwiatkowski, D., Phillips, P.C.B., Schmidt, P., Shin, Y., 1992. Testing the null hypothesis of stationarity against the alternative of a unit root: how sure are we that economic time series have a unit root? J. Econom. 54, 159–178.
Lean, H.H., Smyth, R., 2007. Do Asian stock markets follow a random walk? Evidence from LM unit root tests with one and two structural breaks. Rev. Pacific Basin Financ. Mark. Policies 10 (15).
Lee, D., Schmidt, P., 1996. On the power of the KPSS test of stationarity against fractionally integrated alternatives. J. Econom. 73 (1), 285–302.
Levine, R., 2001. International financial liberalization and economic growth. Rev. Int. Econ. 9 (4), 688–702.
Marashdeh, H.A., Shrestha, M.B., 2010. Stock market integration in the GCC countries. Int. Res. J. Finance Econ. 37, 104–114.
Marinucci, D., Robinson, P.M., 2001. Semiparametric fractional cointegration analysis. J. Econom. 105, 225–247.
Martins, L.F., Rodrigues, P.M.M., 2012. Testing for persistence change in fractionally integrated models. An application to world inflation rates. Comput. Stat. Data Anal. 76, 502–522.
Masih, A., Masih, R., 1997. On the temporal causal relationship between energy consumption, real income, and prices: some new evidence from Asian-energy dependent NICs based on a multivariate cointegration/vector error-correction approach. J. Policy Modell. 19 (4), 417–440.
Mayoral, L., 2006. Further evidence on the statistical properties of real GNP. Oxf. Bull. Econ. Stast. 68, 901–920.
McLeod, A.I., Hipel, K.W., 1978. Preservation of the rescaled adjusted range. A reassessment of the Hurst phenomenon. Water Resour. Res. 14, 491–507.
Mensah, J.O., Alagidede, P., 2017. How are Africa's emerging stock markets related to advanced markets? Evidence from copulas. Econ. Modell. 60, 1–10.
Michelacci, C., Zaffaroni, P., 2000. (Fractional) beta convergence. J. Monetary Econ. 45 (1), 129–153.
Murthy, V., Washer, K., Wingender, J., 2011. Are stock prices in the US nonstationary? Evidence from contemporaneous unit root tests. Appl. Financ. Econ. 21 (22), 1703–1709.
Narayan, P.K., 2005. The saving and investment nexus for China: evidence from cointegration tests. Appl. Econ. 37 (17), 1979–1990.
Narayan, P.K., 2006. Panel data, cointegration, causality and Wagner's law: empirical evidence from Chinese provinces. China Econ. Rev. 19 (2), 297–307.
Narayan, P.K., Smyth, R., 2004. Cointegration of stock markets between New Zealand, Australia and the G7 economies: searching for co-movement under structural change. Aust. Econ. Papers 44 (3), 231–247.
Ng, A., 2000. Volatility spillover effects from Japan and the US to the Pacific-Basin. J. Int. Money Finance 19 (2), 207–233.
Ng, S., Perron, P., 2001. Lag length selection and the construction of unit root tests with good size and power. Econometrica 69, 1519–1554.
Patton, A.J., 2006. Modelling asymmetric exchange rate dependence. Int. Econ. Review 47 (2), 527–556.
Phylaktis, K., Ravazzolo, F., 2002. Stock market linkages in emerging markets: implications for international portfolio diversification. J. Int. Financ. Mark. Inst. Money 15, 91–106.
Phillips, P.C.B., Perron, P., 1988. Testing for a unit root in time series regression. Biometrika 75 (2), 335–346.
Pukthuanthong, K., Roll, R., 2009. Global market integration: an alternative measure and its application. J. Financ. Econ. 94 (2), 214–232.
Ratanapakorn, O., Sharma, S.C., 2007. Dynamics analysis between the US stock return and the macroeconomics variables. Appl. Financ. Econ. 17 (4), 369–377.
Robinson, P.M., 1994. Efficient tests of nonstationary hypotheses. J. Am. Stat. Assoc. 89, 1420–1437.
Robinson, P.M., 1995a. Gaussian semi-parametric estimation of long range dependence. Ann. Stat. 23, 1630–1661.
Robinson, P.M., 1995b. Log-periodogram regression of time series with long range dependence. Ann. Stat. 23, 1048–1072.
Robinson, P.M., Yajima, Y., 2002. Determination of cointegrating rank in fractional systems. J. Econom. 106, 217–241.
Sharma, S.C., Wangbangpo, P., 2002. Stock market and macroeconomic fundamental dynamic interaction: ASEAN-5 countries. J. Asian Econ. 13 (1), 27–51.
Sowell, F., 1992. Maximum likelihood estimation of stationary univariate fractionally integrated time series models. J. Econom. 53, 165–188.
Syllignakis, M.N., Kouretas, G., 2011. Dynamic correlation analysis of financial contagion: evidence from the Central and Eastern European markets. Int. Rev. Econ. Finance 20 (4), 717–732.
Tai, C.S., 2007. Market integration and contagion: evidence from Asian emerging stock and foreign exchange markets. Emerg. Mark. Rev. 8 (4), 264–283.
Taylor, M.P., Tonks, I., 1989. The internationalisation of stock markets and the abolition of U.K. exchange control. Rev. Econ. Stat. 71 (2), 332–336.
Velasco, C., 1999a. Gaussian semiparametric estimation of nonstationary time series. J. Time Ser. Anal. 20, 87–127.
Velasco, C., 1999b. Nonstationary log-periodogram regression. J. Econom. 91, 299–323.
Wong, W.K., Agarwal, A., Du, J., 2004. Financial integration for India stock market: a fractional co-integration approach. Finance India 18 (4), 1581.
Yang, J., Kolari, J.W., Min, I., 2003a. Stock market integration and financial crises: the case of Asia. Appl. Financ. Econ. 13 (7), 477–486.
Yang, J., Min, I., Li, Q., 2003b. European stock market integration: does EMU matter? J. Bus. Finance Account. 30 (9–10), 1253–1276.
Yi, Z., Heng, C., Wong, W.K., 2009. China's stock market integration with a leading power and a close neighbor. J. Risk Financ. Manage. 2 (1), 38–74.