The unemployment-stock market relationship in South Africa: Evidence from symmetric and asymmetric cointegration models

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THE UNEMPLOYMENT-STOCK MARKET RELATIONSHIP IN SOUTH AFRICA: EVIDENCE FROM SYMMETRIC AND ASYMMETRIC COINTEGRATION MODELS

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ABSTRACT: In this study, we examine linear and nonlinear cointegration and causal relations between unemployment and stock market returns in South Africa using quarterly data collected between 1994:Q1 and 2016:Q1. Our empirical results reveal significant cointegration effects between the time series in both linear and nonlinear models, even though both frameworks ultimately reject the notion of any causal relations between the variables. Collectively, our study rejects the notion of unemployment being a good predictor for stock market returns and neither do developments in the stock market have any effect on the unemployment rate. Such evidence advocates for weak-form efficiency in the JSE equity prices whereby unemployment data cannot help investors to predict the movement of future share prices and further suggests that policymakers cannot rely on stock market development as an avenue towards lowering the prevailingly high levels of unemployment as set in current macroeconomic policy objectives.

Keywords: Stock market returns; Unemployment; Cointegration; Causality effects; MTAR model; South Africa.

JEL Classification Code: C13; C22; C52; E23; E44.

1 INTRODUCTION

The Johannesburg Stock Exchange (JSE) is the 2nd oldest stock market in Africa, the 17th largest stock exchange in the world, the sixth largest among emerging economies and the largest within the African continent, with over 400 listed companies, over 900 securities and a market capitalization of over 900 billion US dollars in 2013 (Hussan, 2013). The JSE also has the largest number of cross-listed firms compared to other African stock exchanges and conducts trade on international platforms such as the London Stock Exchange (LSE), the New York Stock Exchange (NYSE), the Frankfurt Stock Exchange (FSE) and the SIX Swiss Exchange. Moreover, the JSE has recently introduced collation centres countrywide which allows for trades to be conducted 400 times faster and currently the number of trades is up by 57 percent, volumes are up by 4 percent and the value of trades up by almost 21 percent (Yartey, 2008). Notwithstanding the relative size and increasing sophistication of the JSE, unemployment and poverty in South Africa remains one of the highest in the world due to the lingering effects of the previous Apartheid legacy. The South African government is currently embarking on macroeconomic policies such as the New Growth Path (NGP) which aims to reduce unemployment from it's current figure of 25 percent to 15 percent by the year 2020. It would therefore be of great interest to policymakers and investors alike to deduce an empirical relationship between stock market development and unemployment in South Africa, with the hope that stock market returns could help foster a macroeconomic environment conducive towards lowering unemployment or that stock returns can be used as an indicator for future movements in stock returns.

Even though the current literature contends that stock market development is an important condition for economic growth in South Africa (Nyasha and Odhiambo (2015) and Phiri (2015b)), very little is known concerning the relationship between stock market activity and unemployment in the country. This is highly noteworthy since unemployment is traditionally known as a measure of the health of an economy and of recent has been viewed as a highly efficient predictor of stock market behaviour especially in developed stock market exchanges (Boyd et. al., 2005). There are two viewpoints to this debate. On one hand, unemployment can be found to granger cause stock market returns. Such evidence would violate the conventional view of the JSE being weak-form efficient (see (Appiah-Kusi and Menyah, (2003) and Phiri (2015a)) and implies that investors can base their future investment decisions on actual or expected unemployment data. On the other hand, if stock market returns are found to lead unemployment, stock market development can be thought of as a vehicle towards eradicating unemployment and poverty in the country. In also considering the historical combination of constant growth in stock returns and South Africa's high unemployment rate in post-Apartheid regime, it would not seem unreasonable to speculate that unemployment and stock market activity are positively correlated for the economy. Notably, South Africa's situation is similar to that of Nigeria where high stock performance has been accompanied with soaring unemployment rates (Bamidele, 2015). And yet, South Africa's situation is also contradictory to that of other leading African stock exchanges such as in Mauritius (Stock Exchange of Mauritius), Egypt (Egyptian Stock exchange) and Morocco (Casablanca Stock Exchange) which have highly developed stock exchanges in combination with low unemployment rates of 7, 9 and 13 percent, respectively.

Thus far, the bulk of the current empirical literature examining the relationship between unemployment and stock exchange activity is concentrated on industrialized economies (Farmer (2015) for the US; Farsio and Fazel (2013) for the US, China and Japan; and Fitoussi et. al. (2000) for 19 OECD countries). With exception of the work of Bamidele (2015) for Nigeria, there is virtually no other empirical research existing for other African countries on the subject matter. Moreover, a majority of previous empirical works have traditionally conducted their investigations by relying on symmetric cointegration frameworks of Engle and Granger (1987) and Johansen (1991) (see Jagannathan and Wang (1993), Jagannathan et. al. (1998), Farsio and Fazel (2013), Farmer (2015) and Bamidele (2015)). It has recently become well known that these linear cointegration frameworks have low testing power and inferior size properties in the presence of asymmetric adjustment between a pair of time series. Besides, previous empirical evidence of nonlinearity existing in the individual times series of unemployment and stock market returns data for South Africa has also been recently provided in the studies of Phiri (2014) and Phiri (2015a), respectively. Primarily motivated by this, our current study contributes to the literature by examining cointegration and causal relations between unemployment and stock market returns for South African data. In order to increase the robustness of our study we employ two empirical frameworks for our analysis; the first being the linear cointegration framework of Engle and Granger (1987) supplemented with cointegration tests proposed by Johansen (1991), and second is the momentum threshold autoregressive model of Engle and Granger (1998) and Enders and Silkos (2001). We conduct our empirical analysis on quarterly data collected post-Apartheid period of 1994:Q1 to 2016:Q1.

Against this background, we structure the remainder of the paper as follows. The next section provides a synopsis of the relationship between stock returns and unemployment in the form of a literature review. The third section of the paper focuses on stock market and unemployment developments in South Africa from a historic perspective. The research methodology is outlined in the fourth section of the paper whilst the data and the empirical results are given in the fifth section of the paper. The study is concluded in the sixth section in the form of policy discussions as well as avenues for future research.

2 LITERATURE REVIEW

Mainstream economy theory depicts on a strong link between stock market activity and unemployment. The capital asset price model (CAPM) as pioneered by Sharpe (1964), Lintner (1965a,b), Mossin (1966) and Black (1972) as well as the discounted cash flow (DCF) model of Graham and Dodd (1934) were amongst the first frameworks used to depict a causal relation between stock prices and unemployment. Within the CAPM model, correlating movements between stock market prices and unemployment is facilitated through one or more of the following three primitive factors; i) the risk-free rate of interest, ii) the expected growth rate of corporate earnings and dividends, and iii) the equity risk premium (Boyd et. al., 2005). On the other hand, the standard DCF model equates the stock price to the discounted present value of a firm's future cashflows, which in turn is linked to the labour demand of firms through the wage curve. A number of empirical papers have provided support for an equilibrium relationship between unemployment and stock market variables based on the channels depicted in the CAPM and DCF models. For instance, the earlier studies such as Fama (1981), Chen et. al. (1986), Geske and Roll (1983) and Mandelker and Tandon (1985) were able to demonstrate that a large number of economic and non-equity financial variables affect discount rates, the ability of firm's to generate cash flows, and future dividend payments. A latter group of studies exclusively found an equilibrium relationship between stock market returns and unemployment. Amongst these studies are the works of Jagannathan and Wang (1993), Jagannathan et. al. (1998), Phelps (1999), Farsio and Fazel (2013), Farmer (2015) and Bamidele (2015).

Another theoretical proposition linking stock market prices and unemployment is based on the sectoral shift hypothesis of Lilien (1982). According to this theory, unemployment is, in part, the result of labour shifts from those sectors where relative wages are declining to those sectors were relative wages are expanding. Initially Lilien (1982) demonstrated that the dispersion of unemployment across industries was a useful proxy in explaining movements in the unemployment rate. However, Lilien's (1982) index was criticized by Abraham and Katz (1986) on the basis of being contaminated with by aggregate demand influences. Consequentially, Black (1987), Loungani et. al. (1990), Brainard and Cutler (1993), Fourtin and Thivierge (1997) and Loungani and Trehan (1997) improved on Lilien's (1982) index by demonstrating that stock market dispersion is a much better proxy for the volume of intersectoral shifts since it gives an early signal of shocks that affect sectors differently and puts more weight on shocks that investors expect to be permanent. These developments resulted in a handful of studies investigating the effects of stock market diversion on unemployment and the evidence provided so far can at best be described as inconclusive. For instance, Dopke and Pierdzioch (2000) find that the influence of stock market diversion on output and unemployment is significant but rather small. Conversely, Chehal et. al. (2010) find that stock market dispersion leads to unemployment over the short-run but not over the long-run. On the other hand, Chen et. al. (2011) find that stock market dispersion accounts for a significant portion of both long-term and short term US unemployment even after controlling for aggregate factors, such that an increase in stock market dispersion leads to an increase in the unemployment level. Furthermore, Jorgensen et. al. (2012) find a positive but weak effect from US earnings dispersion to unemployment for the data and this result is similar to that obtained in Dopke and Pierdzioch (2000). And even more recently, Kalay et. al. (2015) found that US earnings dispersion is associated with higher unemployment and lower industrial production during recession periods, whereas during expansions dispersion has an insignificant impact on unemployment and production.

The final theoretical propositions linking unemployment to stock returns can be attributed to two framework's, the first being Blanchard's (1981) IS-LM model and the second being the Diamond-Mortensen-Pissarides (DMP) model of Diamond (1982), Mortensen (1985) and Mortensen-Pissarides. On one hand, Blanchard (1981) develop an IS-LM model which in equilibrium, macroeconomic news can be good or bad depending on the state of the economy. Cutler et. al. (1988), Orphanides (1992), McQueen and Roley (1993), Veronesi (1999), Boyd et. al. (2005), Cakan (2012), Krueger and Fortson (2003), and Cakan et. al. (2015) all offer support on the notion that stock returns react to unemployment news. On the other hand, the DMP model specifically relates unemployment to job-creation incentives. When the incentive for job creation falls, the labour market slackens and unemployment increases. The DMP model has been recently used as a theoretical workhorse to demonstrate the effects of unemployment on the stock market. For instance, Mukoyama (2009) demonstrate that discount factors of either entrepreneurs or workers are procyclical and these procyclical discount factors can magnify labour market volatility and thus influence unemployment. In particular, the author discovers that entrepreneurs discount factors exerts a larger influence on labour market volatilities compared to the discount factors for workers. In a different study, Kuehn et. al. (2012) build a general-equilibrium model which combines a DMP labour market with full treatment of financial markets. In the model, volatility in allocations resulting from amplified productivity shocks in the labour market causes financial volatility which then widens the equity premium in financial markets. The authors demonstrate that equity premium is countercyclical and can be predicted by labour market tightness. Furthermore, Hall (2014) use a DMP labour market to show that discount rate is the driving force of unemployment such that stock market falls during recessionary periods because the discount rate rises. However, the author is unable to account for why the discount rate falls so much during recessionary periods like the 2009 global recessionary period. Meanwhile, Kilic and Wachter (2015) develop a model with a DMP labour market with an ad-hoc sticky-wage specification as a means of further investigating the underlying force behind the cyclical behaviour and unemployment and vacancies in relation to equity markets. The authors find that during rare disaster events such as the global recession period, high unlevered equity premium is the source of labour and stock market volatility which simultaneously lowers stock market valuation and rises unemployment. Finally, Miao et. al. (2016) introduce credit constraint within a DMP labour market which produces multiple equilibria positions. In one equilibrium there exists bubble in the stock market which relaxes credit constraints and allows firms to increase investment and hire more workers. However, when the bubble bursts credit constraints tighten causing firms to decreasing investment and cut workers hence creating unemployment.

3 STOCK MARKET AND UNEMPLOYMENT DEVELOPMENTS IN SOUTH AFRICA: A POST-APARTHEID SYNOPSIS

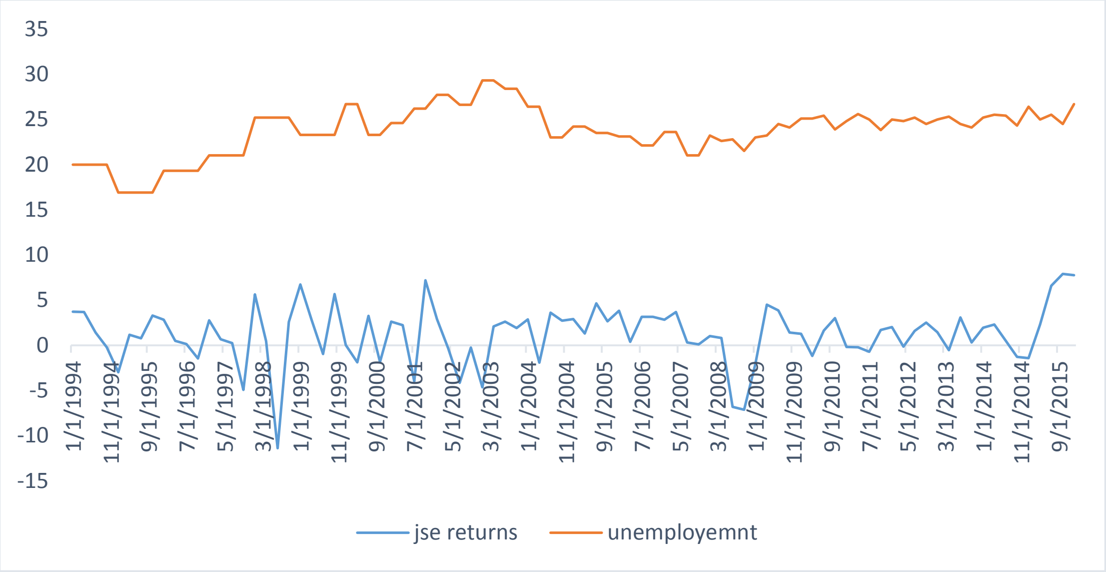
From a historical perspective, both stock market returns and the unemployment rate in the post-apartheid era appear to have been more-or-less positively correlated, with increases in stock market activity appearing to go hand-in-hand with increases in the unemployment rate. Following South Africa's democratic transition of 1994, unemployment in the country averaged just above 16 percent and at this time JSE stock returns averaged slightly over 2 percent. In 1994 the JSE proposed amendments to national government which were designed to improve the efficiency and the liquidity of the stock exchange. On the other hand, fiscal authorities implemented two policy programmes; firstly the Reconstruction and Development Programme (RDP) in 1994, and then secondly the Growth, Employment and Redistribution (GEAR) programme in 1996. The later programme was seen as an upgrade of the former and aimed to specifically create 400 000 new jobs every year through various public works programmes. However, between 1996 and 1998, unemployment had risen from 19 percent to just under 26 percent whilst at the same time JSE market returns increased from 2.5 percent to 3.5 percent. According to Von Fintel and Burger (2014) this sharp increase in unemployment in early post-Apartheid period is a result of long-run generational changes in which older 'more-employable' generations were exiting the labour market whilst the younger generation entered the labour market with a greater probability of remaining unemployed and this created a new high unemployment equilibrium. Improvements experienced in stock market returns during the 1996 to 1998 period can be attributed to i) the introduction of the electronic trading system, the JET system, which was an upgrade from the previous outcry system; ii) the launching of the real-time stock exchange news service (SENS) which enhanced transparency and investor confidence and iii) the opening of trading to foreign nationals. Collectively, this resulted in a drastic increase in stock trades volumes and market liquidity.

However, following the Asian financial crisis in 1998, stock prices returns in South Africa averaged -0.7 percent between 1998 and early 1999, and then picked up to 3.5 percent in late 1999. In 2001, the JSE entered into an agreement with the LSE and in 2002 began trading on the LSE using the LSE stock exchange electronic trading system (SETS). This was accompanied with exceptional stock market performance with returns averaging over 4 percent between 2001 and 2003 except for the period immediately following the September 11 attacks on the World Trade Center, when the JSE experienced a slump averaging -0.45 percent in stock returns. Also in the aftermath of the 9/11 event, South Africa experienced her worst unemployment rates, averaging a record-high 30 percent in 2002, thus ranking it as the 5th highest in the world. Further contributing to South Africa's woes was the shrinking mining sector which further exacerbated the already increasing unemployment rates. In 2004, the GEAR policy programme was phased out and ultimately replaced by Accelerated and Growth

Initiative for South Africa (ASGISA). This government programme was mandated to halve unemployment by the year 2014 mainly through the vehicle of job creation. Unemployment rates fell from averages of 28 percent to 21 percent between 2004 and 2007, which partially reflected the implementation of the ASGISA programme. On the other hand, stock prices were on an upward trend from 2003 up until early 2006 averaging 2 percent returns. In 2007, the JSE experienced a significant shift in her trading mechanism when the LSE leased yet another trading platform to JSE, the JSE TradeElect trading system. However, the collapse of Lehman Brothers in September 2008 eventually took a toll on the South African economy. During the period of 2007 to 2009, market return averages fell from 1.7 percent to -3 percent and unemployment increased from 21 to 25 percent with over 1 million jobs being lost during the recessionary period of 2009.

Nevertheless, the JSE began to recover from the recession in late 2009 averaging 1.9 percent in stock returns whilst unemployment slightly decreased to an average rate of 24 percent. In 2010, government announced the New Growth Path (NGP) programme which set explicit goals of creating 5 million jobs and reducing unemployment to 15 percent by 2030. However, between 2010 and 2013 unemployment and stock market returns both remained more-or-less at steady levels averaging 25 percent and 1.2 percent respectively during this period. In 2013, the JSE decided to shift the trading platform from London back to Johannesburg under a new trading platform, the Millennium exchange trading platform. It is under this platform that the JSE ushered in collation facilities which increased transactions by almost 400 times faster than the previous TradeElect system. This consequentially resulted in stock returns increasing to rates as high as 6 percent in 2015. On the other hand, government introduced the national development plan (NDP) in 2013 and this programme is working on a long-term plan to reduce poverty and eliminate inequality by 2030 through sectoral employment programmes. Nevertheless, unemployment has not improved since then and is currently averaging 25.1 percent thus ranking the country as having the 8th highest unemployment rate in the world. Figure 1 presents a graphical depiction of the seemingly positive co-movement between JSE stock market prices and unemployment in the post-Apartheid era.

Figure 1: JSE stock returns and unemployment: 1994-2015



4 EMPIRICAL FRAMEWORK

4.1 Linear cointegration model

Since it is well known that both unemployment and stock returns are endogenous variables, we base our empirical framework on the premise of specifying two long-run bivariate regression equations. Under the first regression, stock returns is set-up as the dependent variable i.e.

$$smr_t = \alpha_t + \beta unemp_t + e_t \quad (1)$$

Where smr_t are stock market returns, $unemp_t$ is the unemployment rate and e_t is the long-run regression error term. Under the second regression, unemployment is specified as the dependent variable i.e.

$$unemp_t = \alpha_t + \beta smr_t + e_t \quad (2)$$

According to Engle and Granger (1987), any long-run regression which is estimated for a pair of time series variables will produce spurious results if the times series are not found to be cointegrated over time. Therefore, Engle and Granger (1987) suggest that cointegration within the system of equations can be validated if the individual time series are first difference stationary (i.e. integrated of order I(1)) and the cointegration residuals are found to be levels stationary (i.e. integrated of order I(0)) such there exists a cointegration vector comprising of a linear combination of the time series. Furthermore, the residuals of the cointegration vector can be normalized for the time series through an error correction model (ECM) which measures the deviation of the series from its steady-state equilibrium. In its simplest form, a bi-variate error correction model (ECM) between the two time series, smr_t and $unemp_t$, assumes the following functional form:

$$\begin{pmatrix} \Delta smr_t \\ \Delta unemp_t \end{pmatrix} = \gamma ect_{t-1} + \sum_{i=1}^{p} \alpha \Delta smr_{t-i} + \sum_{i=1}^{p} \beta \Delta unemp_{t-i} + \mu_t \quad (3)$$

Where ect_{t-1} is the error correction term of the time series towards its long-run equilibrium and the coefficients α and β measure the short-run effects of the time series variables on the dynamic model. Granger causality can further facilitated under the ECM framework. In particular, smr_t granger cause $unemp_t$ if the α coefficients are found to be significantly different from zero whereas $unemp_t$ granger cause smr_t if the β coefficients are significantly different from zero.

4.2 Threshold cointegration model

Even though Engle and Granger's (1987) cointegration procedure is usually appraised for it's computational ease, it has come under severe criticism by the likes of Enders and Granger (1998) and Enders and Silkos (2001), who demonstrate that the conventional linear cointegration framework exhibits low power and poor size properties in the presence of asymmetric adjustment. As a means of circumventing this issue, Enders and Silkos (2001) allow the residual deviations (ξ_{ti}) of the long-run cointegration regression to behave as threshold processes. In particular, these authors propose two variants of the threshold process.

The first process is the TAR model, which captures deep movements in the equilibrium errors and is specified as follows:

$$e_t = \rho_1 e_t(e_t < \tau) + \rho_2 e_t(e_t \geq \tau) + \nu_t \tag{4}$$

The second process is the MTAR and is designed to capture sharp movements in the equilibrium errors:

$$e_t = \rho_1 e_t(e_t < \tau) + \rho_2 e_t(e_t \geq \tau) + \nu_t \tag{5}$$

Where τ is unknown threshold level which is estimated using the minimization criterion proposed by Hansen (2000), ρ_1 measures asymmetric adjustment when the equilibrium error is below its threshold and ρ_2 measures asymmetric adjustment when the equilibrium error is below its threshold. For both TAR and MTAR versions of the model, Enders and Silkos (2001) advice on the testing of two hypothesis, namely, for cointegration relations between the time series and then for asymmetric cointegration relations. The null hypothesis of no cointegration is given as H_{10}: $\rho_1 = \rho_2$ and is tested against the alternative of cointegration amongst the variables (i.e. H_{11}: $\rho_1 \neq \rho_2$) using a standard F-test denoted as ϕ. If the null hypothesis of no cointegration is rejected, then one can proceed to further test for the null hypothesis of linear cointegration (i.e. H_{20}: $\rho_1 = \rho_2 = 0$) against the alternative of asymmetric cointegration (H_{21}: $\rho_1 \neq \rho_2 \neq 0$) using a modified F-test denoted as ϕ^*. Once the null hypothesis linear cointegration is rejected in favour of asymmetric cointergation, then corresponding threshold error correction (TEC) models can be specified as follows:

$$\begin{pmatrix} \Delta smr_t \\ \Delta unemp_t \end{pmatrix} = \gamma_1 ect_{t-1}(ect_t < \tau) + \sum_{i=1}^{p} \alpha \Delta smr_{t-i} + \sum_{i=1}^{p} \beta \Delta unemp_{t-i} + \{\gamma_2 ect_{t-1}(ect_t \geq \tau) + \sum_{i=1}^{p} \alpha \Delta smr_{t-i} + \sum_{i=1}^{p} \beta \Delta unemp_{t-i}\} + \mu_t \tag{7}$$

$$\begin{pmatrix} \Delta smr_t \\ \Delta unemp_t \end{pmatrix} = \gamma_1 ect_{t-1}(ect_t < \Delta\tau) + \sum_{i=1}^{p} \alpha \Delta smr_{t-i} + \sum_{i=1}^{p} \beta \Delta unemp_{t-i} + \{\gamma_2 ect_{t-1}(ect_t \geq \Delta\tau) + \sum_{i=1}^{p} \alpha \Delta smr_{t-i} + \sum_{i=1}^{p} \beta \Delta unemp_{t-i}\} + \mu_t \tag{8}$$

Regressions (7) and (8) are formally known as the TAR-TEC and MTAR-TEC regressions, respectively. Based on these threshold error correction (TEC) regressions, two main sets of hypothesis can be tested for. Firstly, the null hypothesis of no asymmetric error correction model (i.e. H_{30}: $\gamma_1 \neq \gamma_2$) can be tested against the alternative of threshold error correction model (i.e. H_{31}: $\gamma_1 = \gamma_2$). Secondly, the direction of causality amongst the time series can be evaluated by testing whether the coefficient values of Δsmr_{t-i} and $\Delta unemp_{t-i}$ are significantly different from zero. In particular, the null hypothesis that smr_t does not granger cause $unemp_t$ is tested as H_{40}: $\alpha_i = 0$ whereas the null hypothesis that $unemp_t$ does not granger cause smr_t is tested as H_{50}: $\beta_i = 0$. The aforementioned granger tests are facilitated through F-tests.

5 EMPIRICAL ANALYSIS

5.1 Data and unit root tests

Our empirical analysis makes use of quarterly time series data of percentage change in the total share prices for all shares in South Africa and the unemployment rate for people aged 15 to 64 years old which is collected from 1994:Q1 to 2016:Q1 from the Federal Reserve Economic Data (FRED) online database. All analysis is performed on the raw data and we do not employ any transformations on the time series. As a preliminary step towards our coitnegration analysis, we firstly test the time series for unit roots. Since it is well known that conventional unit root tests such as the ADF and PP tests suffer from distortions when the data generating process is close to a unit root, we opt to rely on the so-called "second-generation" unit root tests of Schmidt and Phillips (1992) as well as that of Elliot et. al. (1996). The Schmidt and Phillips (1992) tests have been performed with r and p statistics whereas the Elliot et. al. (1996) DF-GLS tests are performed with a constant and a trend. The results of the performed unit root tests on the time series variables are recorded in Table 1 below and show that in their level none of the test statistics for either time series variables is able to reject the null hypothesis of a unit root at a 1 percent level of significance. However, in their first differences all test statistics manage to reject the unit root hypothesis for all the time series. Evidently our empirical results verify that both the stock market returns and unemployment variables are integrated of order I(1) which is a preliminary condition for cointegration.

Table: Unit root test results

time series	unit root test			
	sp		ers	
	$\hat{\tau}_\tau$	$\hat{\rho}_\tau$	constant	trend
smr_t	-2.58	-8.47	-2.47	-3.18
	(7.12)*	(205.09)*	(-6.37)*	(-6.45)*
$unemp_t$	-2.13	-8.73	-1.16	-2.05
	(6.89)*	(12.54)*	(-4.88)*	(-4.11)*

Notes: '*' represent the 1% significance level. Test statistics for first differences provided in parentheses.

5.2 *Linear cointegration analysis*

In light of verifying that both stock market returns and the unemployment rate are first difference stationary variables, we proceed to test for linear cointegration effects between the time series variables. The number of cointegration vectors (r) within the system of variables is examined through two likelihood ratio tests proposed by Johansen (1991). The first test is the lambda-maximum test which tests the null hypothesis that the cointegration rank is equal to r against the alternative that the cointegration rank is equal to r+1. The test statistic used is a maximum generalized eigenvalue which is computed as:

$$J_{max} = -T\, In(1 - \hat{\lambda}_i) \qquad (9)$$

Where T is the sample size and $\hat{\lambda}_i$ is the i^{th} largest canonical correlation. The second cointegration test is the trace test which tests the null hypothesis that the cointegration matrix is equal to r against an alternative of the cointegration rank being equal to k. The test statistic used is the trace of a diagonal matrix of generalized eigenvalue and is computed as:

$$J_{trace} = -T \sum_{r+1}^{n} In(1 - \hat{\lambda}_i) \qquad (10)$$

Two versions of aforementioned cointegration tests have been performed on our data, the first with a drift and the second with a trend, with the results being reported in Table 2 below.

Table 2: Maximum Eigen and trace cointegration test results

H_0	H_1	J_{max}	99% c.v.	J_{trace}	99% c.v.
with a constant					
$r \geq 1$	$r = 1$ $(r \geq 2)$	4.59	12.91	4.09	11.65
$r \geq 0$	$r = 0$ $(r \geq 1)$	35.30	20.20	39.39	23.52
with a trend					
$r \geq 1$	$r = 1$ $(r \geq 2)$	4.76	16.26	4.76	16.26
$r \geq 0$	$r = 0$ $(r \geq 1)$	36.86	23.65	41.62	30.45

Note: Lag length for the maximum Eigen and trace tests have is 3 s determined by the AIC and BIC.

As can be observed from the cointegration test results reported in Table 2 above, the computed Eigen and Trace statistics advocate for at least one cointegration vector between the time series variables. In particular, when testing the null of no cointegration effects for cointegration rank r=0 inclusive of a constant, we obtain J_{max} and J_{trace} statistis values of 35.30 and 39.39, respectively. Notably these values exceed their corresponding critical values at all levels of significance hence rejecting the null hypothesis of no cointegration vectors. Similarly, when the testing procedure is inclusive of a trend, the J_{max} and J_{trace} statistics exceed their critical values at all significance levels, with values of 36.86 and 41.62, respectively. However, in proceeding to test the null hypothesis of one cointegration relation against the alternative of two cointegration vectors, we obtain J_{max} and J_{trace} statistics of 4.59 and 4.09, respectively, when the test is performed with a constant, whereas both J_{max} and J_{trace} statistic produce a similar value of 4.76 when the test is performed with a trend. We note that these statistics, performed with a constant and a trend, fail to reject the null of one cointegration relation at all significance levels. Nevertheless, in view of verifying one cointegration relation we conclude on unanimous evidence of linear cointegration existing among the variables and that the estimation of long-run and short-run relations can be conducted without concern of obtaining spurious results. We therefore proceed to estimate the long-run regression equations (1) and (2) and their corresponding error correction model equations (3) and (4) for the time series variables and report the results of this empirical exercise in Table 3 below.

Table 3: long run regression results and linear error correction model analysis

	dependent variable			
	smr_t		$unemp_t$	
ψ_0	0.72		23.72	
	(0.81)		(0.00)***	
ψ_1	0.02		0.01	
	(0.88)		(0.88)	
error correction model				
	Δsmr_t	$\Delta uemp_t$	Δsmr_t	$\Delta unemp_t$
Δsmr_{t-1}	0.21	0.06	-0.57	0.03
	(0.31)	(0.49)	(0.00)***	(0.54)
Δsmr_{t-2}	0.03	0.06	-0.51	0.04
	(0.84)	(0.38)	(0.00)***	(0.37)
Δsmr_{t-3}	0.09	-0.01	-0.20	-0.02
	(0.44)	(0.81)	(0.08)*	(0.62)
$\Delta unemp_{t-1}$	-0.05	-0.16	0.08	-0.10
	(0.87)	(0.17)	(0.81)	(0.40)
$\Delta unemp_{t-2}$	-0.62	-0.17	-0.65	-0.13
	(0.03)*	(0.13)*	(0.04)	(0.27)
$\Delta unemp_{t-3}$	-0.38	-0.13	-0.30	-0.09
	(0.18)	(0.26)	(0.35)	(0.45)
ect_{t-1}	-1.03	-0.05	0.13	-0.10
	(0.00)***	(0.61)	(0.38)	(0.07)*
R^2	0.46	0.08	0.46	0.08
DW	1.96	1.88	1.96	1.88
	(0.77)	(0.58)	(0.77)	(0.58)
LB	0.06	0.79	0.06	0.79
causality results				
H_{01}:		0.37		
jse $\rightarrow$unemp		(0.55)		
H_{02}:		0.26		
unemp $\rightarrow$jse		(0.61)		

Notes: p-values are reported in parentheses. "***', '**' and '*' represent 1%, 5% and 10% significance levels, respectively. The Durbin Watson (DW) ststitic for serial correlation indicates that all regressions are free from serial correlation whereas the Ljung-Box (LB) statistic for autocorrelation shows that only the regressions with stock market returns (smr) are free from autocorrelation.

In referring to our long-run regression estimates, the slope regression coefficients are reported in the top portion of Table 3. We particularly find that when stock returns is the dependent variable we obtain an estimate of 0.02 whereas when unemployment is the dependent variable the coefficient estimate is 0.01. However, based on the corresponding p-values these estimates are rendered as being insignificant. The error correction terms, as report in the middle of Table 3, produce a significant negative estimate of -1.03 when stock market returns is the dependent variable and the driving variable in the system. This implies that 103 percent of deviations are corrected each quarter when shock is induced on stock returns. On the other hand, when unemployment is both the dependent variable and the driving variable in the system the error correction term produces a significant negative coefficient of -0.10 which implies that 10 percent of deviations from the steady state are corrected each quarter. Furthermore, we note the lack of significant short run effects in the error correction models hence implying there are no short-term equilibrium reverting effects between stock returns and unemployment. In finally turning to our causality results, as reported at the bottom of Table 3, we find F-values of 0.37 when testing the null that stock returns does not ganger cause unemployment and a F-value of 0.26 when testing the other null hypothesis that unemployment does not granger cause stock returns. Based on the p-values associated with both test statistics, we are unable to reject both null hypotheses thus implying that there are no causality effects between both time series variables. We note that this result in line with that obtained in Farsio and Fazel (2013) who also find that no causality effects between unemployment and stock returns for the more developed economies of the USA, Chain and Japan.

5.3 *Threshold regression analysis*

Having examined linear cointegration effects between the time series, we now turn our attention towards possible nonlinear cointegration relations between the variables. To do so, we firstly perform the two threshold cointegration tests of Enders and Siklos (2001). To recall, the first test is testing the null hypothesis of no cointegration effects against the alternative of linear cointegration effects using the ϕ statistics whereas the second test is testing the null hypothesis of linear cointegation against the alternative of threshold cointegration using the ϕ^* statistic. As reported in the top panel of Table 5, we find that the ϕ statistics produces values of 26.36 and 25.38 for the TAR and MTAR regressions, respectively when stock returns variable is employed as the endogenous variable in the system. Notably, these statistics indicate that we must reject the null of no cointegration at all significance levels. On the other hand, we obtain weaker ϕ statistic values of 3.40 and 2.79 for the TAR and MTAR regressions, respectively, when unemployment is the dependent variable and these statistics manage to reject the null of no cointegration at a 10 percent significance level. In proceeding to test for threshold cointegration effects we observe that the ϕ^* statistic produces values of 6.05 and 4.77 for the TAR and MTAR regressions, respectively when stock returns is the dependent variable whereas we find values of 1.94 and 0.78 for TAR and MTAR regressions, respectively, when unemployment is the dependent variable. Both ϕ^* statistics for the stock market regressions reject the linearity hypothesis at a 10 percent level of significance whereas both statistics under the unemployment regressions fail to reject the linearity hypothesis.

Considering that we find significant TAR and MTAR cointegration effects for the stock market returns regressions, we then exclusively estimate the associated long-run regression coefficients and the threshold error correction terms for these particular regressions. In similarity to the long-run coefficient obtained under our linear cointegration analysis, we find a positive yet insignificant estimate of 0.72. In turning to the estimates of our threshold error terms, we report significant estimates of -1.13 and -0.59 for ρ_1 and ρ_2, respectively under the TAR model whereas we also obtain significant estimates of -1.05 and -0.74 for ρ_1 and ρ_2 for the MTAR specification. Notably, these estimates satisfy the asymmetric equilibrium convergence condition ρ_1, $\rho_2 < 0$, which, according to Enders and Silkos (2001) ensures that the equilibrium error terms are stationary and yet exhibit asymmetric behaviour. In paying closer attention to the threshold equilibrium error term estimates we further note that for both TAR and MTAR specifications, $\rho_2 < \rho_1$, hence implying that positive deviations from the equilibrium are corrected quicker than negative deviations. Interpretively, our results imply that the variables revert to their steady state position by 113 percent in the case of a negative deviation and by 59 percent for a positive deviation, for the case of the TAR model. On the other hand, equilibrium reversion occurs by 105 percent for negative deviations and by 74 percent for positive deviations when dictated by MTAR dynamics. Overall, this implies that negative shocks, commonly in the form of adverse external shocks, to either stock returns or the unemployment rate are quickly absorbed in the system in comparison to positive shocks, which may be in the form of government and other regulatory policies.

Table 5: TAR and MTAR cointegration tests and estimates

	dependent variable			
	smr_t		$unemp_t$	
	TAR	MTAR	TAR	MTAR
ϕ	26.36	25.38	3.40	2.79
H_{00}: $\rho_1 = \rho_2 = 0$	(0.00)***	(0.00)***	(0.04)*	(0.07)*
ϕ^*	6.05	4.77	1.94	0.78
H_{01}: $\rho_1 = \rho_2$	(0.01)*	(0.03)*	(0.17)	(0.38)
τ	1.58	1.82	2.72	0.04
Ψ_0	0.72	0.72		
	(0.81)	(0.81)		
Ψ_1	0.02	0.02		
	(0.88)	(0.88)		
ρ_1	-1.13	-1.05		
	(0.00)***	(0.00)***		
ρ_2	-0.59	-0.74		
	(0.00)***	(0.00)***		

Notes: p-values are reported in parentheses. "***', '**' and '*' represent 1%, 5% and 10% significance levels, respectively.

In light of evidence supporting asymmetric adjustment existing between JSE stock returns and the unemployment rate, we present the threshold error correction (TEC) test and model estimates of the TAR and MTAR variants using the consistent threshold estimates obtained in our previous threshold cointegration regressions. In firstly testing the null hypothesis of no threshold error correction effects against the alternative of threshold error correction effects, we find insignificant statistics of 1.45 and 1.41, respectively, for TAR and MTAR variants of the TEC specifications when unemployment is the driving variable in the system. This result implies that there are no significant TEC effects when unemployment is the driving variable in the system. Conversely, when stock returns is the driving variable in the system, the statistics of 4.25 and 2.65, respectively obtained for the TAR and MTAR models are significant and hence imply that we cannot reject the null hypothesis of TEC effects. Moreover, when stock returns is the driving variable in the system we find a negative and significant error correction terms of -1.18 and -1.11 respectively for the TAR and MTAR models in the lower regime of the TEC specifications. On the other hand, the error correction terms for the TAR and MTAR specifications in the upper regime of the TEC system produce negative and significant estimates of -0.60 and -0.57, respectively, when stock returns is the driving variable in the system. Collectively, we treat this as evidence of long-run equilibrium reverting behaviour in the face of shock to the system. We also observe insignificant short-run coefficients regardless of whether stock returns or unemployment is the driving variable in the system for both TAR and MTAR variants of the TEC regressions. Note that this result is on par with that obtained in our linear error correction model estimates. In lastly testing for causality effects between unemployment and stock returns, none of the test statistics for the TAR and MTAR models is able to reject neither null that unemployment does not granger cause stock returns or that stock returns does not cause unemployment. Once again, these results are in coherence with those presented in our linear cointegration analysis and also with those presented by Farsio and Fazel (2013) for the USA, China and Japan.

Table 6: Threshold error correction tests and estimates

model type	TAR		MTAR	
dependent variable	Δsmr_t	$\Delta unemp_t$	Δsmr_t	$\Delta unemp_t$
Δsmr^{-}_{t-1}	0.11	0.04	0.07	0.04
	(0.63)	(0.67)	(0.77)	(0.70)
$\Delta unemp^{-}_{t-1}$	0.43	-0.06	0.42	-0.05
	(0.36)	(0.75)	(0.38)	(0.78)
ect^{-}_{t-1}	-1.18	-0.04	-1.11	-0.04
	(0.00)***	(0.58)	(0.00)***	(0.62)
Δsmr^{+}_{t-1}	0.15	-0.01	0.17	-0.01
	(0.33)	(0.88)	(0.28)	(0.87)
$\Delta unemp^{+}_{t-1}$	-0.11	-0.13	-0.02	-0.11
	(0.79)	(0.46)	(0.97)	(0.55)
ect^{+}_{t-1}	-0.60	0.09	-0.57	0.12
	(0.01)**	(0.30)	(0.03)*	(0.26)
R^2	0.46	0.04	0.45	0.04
DW	1.89	2.03	1.93	2.05
	(0.63)	(0.88)	(0.77)	(0.86)
LB	0.01	0.97	0.01	0.97
hypotheses tests				
H_{30}: $\gamma_1 \neq \gamma_2$	4.25	1.45	2.65	1.41
	(0.04)**	(0.23)	(0.10)*	(0.24)
H_{40}:	0.72	0.09	0.41	0.27
smr →unemp	(0.49)	(0.91)	(0.67)	(0.77)
H_{50}:	0.43	0.42	0.74	0.08
unemp →smr	(0.65)	(0.66)	(0.48)	(0.92)

Notes: p-values are reported in parentheses. "***', '**' and '*' represent 1%, 5% and 10% significance levels, respectively. The Durbin Watson (DW) ststitic for serial correlation indicates that all regressions are free from serial correlation whereas the Ljung-Box (LB) statistic for autocorrelation shows that only the regressions with stock market returns (smr) are free from autocorrelation.

6 CONCLUSION

This study becomes the first to investigate the empirical relationship between stock market returns and unemployment for the South African economy using post-Apartheid quarterly data collected from 1994:01 to 2016:01. To ensure a considerable level of robustness the empirical analysis was performed using both linear and nonlinear cointegration frameworks. Linear cointegration was conducted using Engle and Granger (1987) two-step cointegration procedure and this was supplemented with Johansen (1991) cointegration tests. On the other hand, nonlinear cointegration analysis was done through the testing and estimation of TAR and MTAR models as outlined in Enders and Granger (1987) as well as in Enders and Silkos (2001). Overall, our empirical results reveal the following. First of all, we find that both linear and nonlinear cointegration frameworks validate the presence of long-run steady-state equilibrium between the time series. However, we are unable to find any short-term cointegration effects between the variables under both frameworks. Secondly, the long-run relationship found between stock market returns and unemployment is positive and yet insignificant. This finding implies that any seemingly positive stock returns-unemployment relationship that may be visually observed by chartists, is purely coincidental. Thirdly, we do not find any causality effects between the time series; that is to say that the information from past values of stock market returns do not feed into the unemployment rate, and vice versa. So whereas we have established cointegration relations between the time series, changes in these variables will not affect the counter variable.

There are a number of interesting phenomenon that policymakers, investors and academics can derive from the empirical results obtained in our study. For instance, policymakers must be aware that the insignificant relationship between stock market returns and unemployment implies that the JSE cannot be used in any direct way to alleviate the ever-troublesome problem of high unemployment in the country. In other words, policymakers are encouraged to stick to their more direct conventional methods of dealing with unemployment such as infrastructural spending and other labour market related strategies. Another implication which can be drawn from our study is that investors or speculators cannot use the domestic unemployment rate to predict or to ‘beat’ the stock market and thus gain superior returns on their investments. This also implies that the JSE displays elements of weak-form efficiency in the sense that publically available information concerning unemployment cannot be used to predict the direction of stock market returns and this is not surprising given that the JSE has been recently ranked as the most efficiently regulated exchange in the world by the World Economic Forum. However, this matter concerning the weak-form efficiency of the JSE is not all conclusive seeing that the unemployment rate was the only macroeconomic variable that was tested for a relationship with stock returns. Therefore future research endeavours may directed towards testing the predictability of the stock market returns against a host of other financial and macroeconomic variables such as the exchange rate, the inflation rate, interest rates, economic growth.

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