The Feldstein-Horioka puzzle and the global recession period: Evidence from South Africa using asymmetric cointegration analysis

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THE FELDSTEIN-HORIOKA PUZZLE AND THE GLOBAL FINANCIAL CRISIS: EVIDENCE FROM SOUTH AFRICA USING ASYMMETRIC COINTEGRATION ANALYSIS

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ABSTRACT: In this study we examine the effects of the 2007-2008 global financial crisis on the Feldstein-Horioka coefficient for South Africa using momentum threshold cointegration and error correction techniques applied to quarterly national savings-investment time series collected between 2000:Q1 and 2017:Q1. Our empirical strategy consists of segregating the data into three samples; one corresponding to the full sample (1960:Q1 – 2016:Q4), another corresponding to the pre-crisis period (1960:Q1-2008:Q3) and the last corresponding to the post-crisis period (2008:Q4-2016:Q4). Our empirical results validate asymmetric cointegration effects for both the full and the pre-crisis periods while only accepting a linear cointegration relation for the post-crisis period. The saving-retention coefficient estimates produced are 0.59 (significant), 0.64 (significant), and 0.22 (insignificant) for the full, pre-crisis and post-crisis periods, respectively. These results imply that international capital mobility has increased in the post-crisis period and this may be primarily due to the effects of a redirection of private capital flows by investors to safe haven assets. Therefore policy plans of further relaxing of capital controls is inadvisable considering that capital is already highly mobile.

Keywords: Investment; Savings; Feldstein-Horioka puzzle; Threshold cointegration; Threshold error correction model; Sub-Saharan Africa; South Africa.

JEL classification codes: C22, C52, F21, F41.

1 INTRODUCTION

In their seminal paper, Feldstein and Horioka (1980) came up with a simple yet controversial proposition that advocated for a high degree of correlation between domestic savings and investment for 16 OECD economies during the post-war period. According to the authors, in absence of international capital mobility, both domestic savings and investment should be highly correlated since such circumstances would require investment to be solely financed by domestic savings (Bangake and Eggoh, 2011). Conversely, in a world of unfettered capital mobility, domestic savings would flow to countries with the highest rate of return and thus domestic savings would be financed by world pool of savings and this would be empirically reflected by a low correlation between savings and investment (Raheem, 2017). From a policy perspective, testing for the existence of an empirical relationship between domestic savings and investment is considered an important academic exercise since it has implications for single currency debates, tax policies on capital and savings, whether growth is constrained by the domestic savings rate and if fiscal deficits will have large crowding out effects private investment (Wang, 2013).

Following the initial conjecture, a large volume of empirical works have embarked on testing the validity of Feldstein and Horioka (1980)'s proposition. On one hand, there is exists a handful of studies which have confirmed the original Feldstein-Horioka hypothesis mainly for industrialized or OECD economies (Murphy (1984); Penati and Dooley (1984); Obstfeld (1986); Dooley et. al. (1987); Golub (1990); Tesar (1991); and Sinn (1992)). Notably most of these empirical conclusions have been drawn from econometric analysis of time series data collected prior to global financial liberalization era of the 1990's. On the other hand, there exists an even larger body of empirical evidence which has rejected the Feldstein-Horioka proposition and renders the hypothesis as a puzzle since increased openness and financial liberalization experienced globally throughout the 1990's would imply higher international capital mobility or a lower savings-retention coefficient (see Armstrong et. al. (1996); Jansen (1996); Krol (1996); Dekle (1996); Shiabata and Shintani (1998); and Sarno and Taylor (1998)). With respect to the latter group of studies, there exist a number of explanations for the

existence of the so-called "Feldstein-Horioka puzzle" which range from the effects of government policy which targets the current account balance (Golub, 1990) to output fluctuations in non-traded goods (Tesar, 1993) to population growth (Obstfeld, 1986), to financial constraints (Chang and Smith, 2014) and yet with all of these different arguments there still exists no definitive consensus pertaining to the debate.

A significant yet overlooked factor which may have altered the saving-investment relations over time relates to the role played by financial crisis periods. Take for instance, Kim and Jeon (2011) who establish higher international capital movements (i.e. lower savings-retention coefficients) in Asian countries for periods subsequent to the Asian financial crisis of 1997-1998. According to these authors, more international capital entered into the Asian region subsequent to the crisis as the International Monetary Fund (IMF) advised these countries to reform their financial and economic systems as well as drastically open their financial markets. These recommendations by the IMF were based on a phenomenon more popularly referred to as a "monetary policy tri-lemma", in which monetary authorities are unable to simultaneously attain capital mobility, a fixed exchange rate and independent monetary policy. At any given time policymakers can only achieve two of the three aforementioned policies (Phiri, 2016). Since economies worldwide were generally perceived as being reliant on international capital flows, most Asian economies opted for a combination of fixed exchange rates and open financial markets whilst other emerging economies, such as South Africa and Indonesia, opted for a combination of open market policies and independent monetary policy (i.e. inflation targeting policy regime).

The most recent financial crisis which emerged as a crash in the U.S. property market, triggered the collapse of the US banking system and eventually propagated adverse spillover effects to financial markets worldwide more concentrated so for European economies. In wake of the global financial crisis, emerged the global recessionary period of 2008-2009 as well as the European sovereign debt crisis of 2010. In terms of international capital flows, the sub-prime crisis led to a repatriation of international capital back to domestic countries and this

repatriation of finance may have possibly affected the way in which saving and investment move across countries (But and Moley, 2016). Empirical evidence recently presented by Katsimi and Zoega (2016) as well as But and Morley (2016) suggests that the global financial crisis increased the savings-retention coefficient or similarly lowered movements in international capital flows in the Euro Area. However, bearing in mind that the effects of the most recent financial crisis are likely to be different between developing and advanced economies, it is quite astonishing to discover that there exists no empirical works, to the best of our knowledge, which have exploited the possibility of a change in the savings-retention coefficients in the aftermath of the crisis for emerging and less developed economies. This becomes even more thought provoking considering the integration of many developing and emerging economies into worldwide financial markets over the last couple of decades via capital account liberalization, stock market liberalization and financial sector liberalization (Le Roux and Moyo, 2015).

In our manuscript, we consider as a case study the South African economy, which arguably boasts the most financially open economy on the African continent, and we empirically examine whether the global financial crisis altered the savings-retention coefficient for the economy. The singular previous South African case study recently presented by Gil-Alana et. al. (2016) used fractional cointegration to establish that the financial deregulation period of the 1980's loosened the steady-state relationship between domestic savings and investment. Nevertheless, the authors are unable to detect a break point corresponding to the global financial crisis and thus it is yet to be established whether the recent financial crisis did have a bearing on the savings-investment relationship for the South African economy. In our study, we examine the effect of the recent financial crisis on the Feldstein-Horioka coefficient for South Africa as a representative of an emerging and Sub-Saharan African (SSA) country. Methodologically, we deviate from the conventional norm of linear cointegration analysis, and use the momentum threshold autoregressive (MTAR) model of Enders and Siklos (2001) as our empirical framework. We consider this framework favourable since linear cointegration models may prove to be too restrictive in accounting for the dynamic relationship between

savings and investment. In particularly considering the long span time series employed in our empirical study, which covers a wide range of political, economic and monetary developments, it is highly unlikely that the savings-investment relationship is symmetric over the steady-state equilibrium.

Against this background, the remainder of the paper is structured as follows. The next section is dedicated to a review of the associated literature. The third section of the paper presents the empirical framework of the study whereas the time series data and empirical results of the study are presented in the fourth section of the paper. The paper is concluded in the fifth section by drawing policy conclusions from the empirical analysis.

2   LITERATURE REVIEW

Over the past 35 years or so, the Feldstein-Horioka puzzle has indeed been one of the most studied phenomenon in international macroeconomics and its place within the academic paradigm has been appropriately branded as one of the six major economic puzzles (Obstfeld and Rogoff, 2000). As a consequence there has been an overwhelming number of studies which have empirically investigated the Feldstein-Horioka hypothesis for a wide range of countries using data collected from various time periods and, for convenience sake, these studies can be segregated into four strands of empirical works, namely; i) studies on industrialized economies; ii) studies on emerging and/or developing economies, iii) studies on both industrialized and developing economies, and iv) nonlinear studies.

2.1 Review of studies on industrialized economies

An overwhelming majority of the existing empirical literature examining the Feldstein-Horioka hypothesis has typically focused on OECD and European Union (EU) countries. Indeed the pioneering work of Feldstein and Horioka (1980) was conducted for OECD

countries as is the case with a majority of other earlier empirical works conducted between the eighties and the early nineties (Feldstein (1983); Murphy (1984), Penati and Dooley (1984), Obstfeld (1986), Dooley et. al. (1987), Golub (1990), Tesar (1991) and Sinn (1992)). However, a vast majority of empirical works conducted for OECD countries came after the financial liberalization period of the 1990's (Argimon and Roldan (1994); Armstrong et. al. (1996); Jansen (1996); Krol (1996); Shiabata and Shintani (1998); Hussein (1998); Coiteux and Olivier (2000); Jansen (2000); Corbin (2001); Cadoret (2001); Kim (2001); Ho (2002); Georgopoulos and Hejazi (2005); Coakley et. al. (2005); Caporale et. al. (2005); Amirkhalkhali and Dar (2007); Pelgrin and Schich (2008); Kollias et. al. (2008); Rao et. al. (2010); Narayan and Narayan (2010); Kumar and Rao (2011); Chu (2012); Ketenci (2013); Johnson and Lamdin (2014); Holmes and Otero (2014); Darkos et. al. (2016); Katsimi and Zoega (2016); and But and Morley (2017)). Apart from these studies conducted for OECD countries there also exists a separate group empirical studies which conducted their empirical analysis on individual industrialized economies such as the US (Miller (1988); Gulley (1992); Moreno (1997); Levy (2000); de Vita and Abbott (2002) and Hoffman (2004)), the UK (Sarno and Taylor (1998); Abbott and de Vita (2003); Ozmen and Parmaksiz (2003)) and Japan (Yamori (1995); Dekle (1996); Narayan (2005) and Guzel and Ozdemir (2011)).

In collectively summarizing the above reviewed literature on these industrialized economies, we observe that the studies of Feldstein and Horioka (1980); Feldstein (1983); Murphy (1984); Penati and Dooley (1984); Obstfeld (1986); Dooley et. al. (1987); Golub (1990); Tesar (1991); Sinn (1992); Moreno (1997); Hussein (1998); Levy (2000); Hoffman (2004) and Darkos et. al. (2016) obtain a savings-retention coefficient of close-to-unity thus implying low capital mobility in OECD economies more prominently so for periods preceding the 1990's. On the other end of the spectrum, Miller (1988); Gulley (1992); Yamori (1995); Armstrong et. al. (1996); Jansen (1996); Krol (1996); Dekle (1996); Shiabata and Shintani (1998); Sarno and Taylor (1998); Coiteux and Olivier (2000); Corbin (2001); Kim (2001); Cadoret (2001); Ho (2002); de Vita and Abbott (2002); Abbott and de Vita (2003); Ozmen and Parmaksiz (2003); Caporale et. al. (2005); Coakley et. al. (2005); Narayan (2005);

Georgopoulos and Hejazi (2005); Amirkhalkhali and Dar (2007); Kollias et. al. (2008); Pelgrin and Schich (2008); Rao et. al. (2010); Kumar and Rao (2011); Guzel and Ozdemir (2011); Chu (2012); Ketenci (2013); Johnson and Lamdin (2014); Holmes and Otero (2014); Katsimi and Zoega (2016); and But and Morley (2017) all find a Feldstein-Horioka coefficient of close to zero or negative, implying high capital mobility in these countries.

2.2 Review of studies on emerging and developing economies

The empirical literature which exists for developing and emerging economies is not as extensive as that observed for industrialized economies and is primarily concentrated on Asian countries (Sinha (2002); Kim et. al. (2005); Kim et. al. (2007); Singh (2008); Brahmasrene and Jiranyakul (2009); Guillaumin (2009); Li (2010); Eslamloueyan and Jafari (2010); Khundrakpam and Ranjan (2010); Kim and Jeon (2011); Chan et. al. (2011); Wang (2013) and Jiang (2014)). A majority of the reviewed Asian literature has verified high mobility in international capital movements (Kim et. al. (2005); Singh (2008); Brahmasrene and Jiranyakul (2009); Guillaumin (2009); Li (2010); Eslamloueyan and Jafari (2010); Kim and Jeon (2011); Wang (2013) and Jiang (2014)) even though there are a few exceptional studies which find low capital mobility within the time series data (Kim et. al. (2007); Khundrakpam and Ranjan (2010); and Chan et. al. (2011)). What makes this group of studies on Asian countries particularly interesting in comparison to the reviewed studies on advanced economies, is the wide use of cointegration techniques used by the authors in obtaining their various empirical results. This renders this group of studies less susceptible to having obtained spurious regressions results in their respective analysis.

Another noteworthy strand of empirical literature for emerging and developing economies is those conducted for BRICS (Konya (2015) and Behera (2015)), Sub-Saharan African (SSA) (de Wet and Van Eyden (2005); Payne and Kumazawa (2005); Cyrille (2010); Adams et. al. (2016) and Raheem (2017)) and African countries (Adedeji and Thornton (2006)

for 6 African countries; Cooray and Sinha (2007) for 12 African countries; Bangake and Eggoh (2011) for 37 African countries; Nindi and Odhiambo (2014) for Malawi; Barros and Gil-Alana (2015) for Angola and Gil-Alana et. al. (2016) for South Africa). This particular reviewed strand of empirical literature is relevant to our study since they tend to include South African time series data in their respective analysis. Notably, most of these studies are panel studies and indicate high capital mobility (i.e. low savings-retention coefficient) among African countries (de Wet and Van Eyden (2005); Payne and Kumazawa (2005); Adedeji and Thornton (2006); Cooray and Sinha (2007); Cyrille (2010); Bangake and Eggoh (2011); Adams et. al. (2016) and Raheem (2017)).

It should be cautioned that the reviewed African panel studies which include South Africa in their dataset are susceptible to being criticized on the premise of including outliers which would influence a change in the relationship between saving and investment. A conspicuous example of this problem pertains to the inclusion of Luxemburg in OECD statistics, of which a number of studies find that exclusion of this country from OECD time series data significantly changes the savings-retention coefficient (Tesar (1991), Jansen (2000) and Coiteux and Olivier (2000)). Therefore individual-specific studies of Behera (2015), Konya (2015) and Gil-Alana et. al. (2016) are of even more interest to us since these works exclusively investigate the Feldstein-Horioka puzzle for South African time series and mutually reveal high levels of international capital mobility more specifically so for periods subsequent to the financial deregulation era of the 1980's.

2.3 Review of studies conducted for both industrialized and developing economies

There also exists a separate cluster of panel data studies which have investigated the saving-investment relationship simultaneously for developed and industrialized economies (Coakley and Kulasi (1997); Sinha and Sinha (2004); Chakrabarti (2006); Adedeji and Thornton (2008); Georgopoulos and Hejazi (2009); Herwartz and Xu (2010); Chang and Smith

(2014); Dzhumashev and Cooray (2016)). Typically, these empirical studies segregate their empirical data into different panels corresponding to industrialized and developing groups of economies and make comparisons of their empirical estimations of the time series afterwards. Interestingly enough, the general consensus derived from this reviewed cluster of studies is that developing or emerging economies have higher levels of international capital mobility (i.e. lower savings-retention coefficients) in comparison to industrialized counties. It is also worth noting that the works of Sinha and Sinha (2004); Herwartz and Xu (2010); Chang and Smith (2014); and Dzhumashev and Cooray (2016) all include South Africa as part of the panel of developing or emerging economies in their analysis.

2.4 Review of nonlinear studies

Even though it is not commonly acknowledged in the literature, the idea of a nonlinear savings-investment relationship was initially explored in the seminal work of Feldstein and Horioka (1980) who estimated a quadratic savings-investment regression and discovered that the quadratic 'savings-retention coefficient' term was insignificant hence indicating no existing nonlinearities. However, recent developments within econometric estimation techniques have resulted in more refined methods of capturing possible nonlinearities within time series data and this has resulted in a handful of studies which have applied highly specialized econometric techniques to capture existing nonlinearities within the saving-investment time series i.e. Ho (2003) for 23 OECD countries; Bautista and Maveyraud-Tricoire (2007) for 7 East Asian economies; Aka (2007) for Ivory Coast and Ghana; Fouquau et. al. (2008) 24 OECD economies; Di Iorio and Fachin (2014) for 18 OECD economies; Dursun and Abasiz (2014) for Turkey; Chen and Shen (2015) for 9 European countries; and Barros and Gil-Alana (2015) for Angola. Typically, these studies argue that linear econometric frameworks may not contain high enough testing power for estimation of time series variables which most probably have underlying nonlinear data generating processes.

Methodologically, the literature presents three main families of nonlinear econometric frameworks which have been used to substantiate nonlinear investment-savings correlations, namely, i) Markov-Switching (M–S) frameworks of Hamilton (1989) ii) the panel threshold autoregressive (PTAR) model and panel smooth transition regression (PSTR) estimation models of Hansen (1999) and Gonzalez et. al. (2005), respectively, iii) and the regime switching cointegration models of Gregory and Hansen (1996); Hansen and Seo (2002) and Hatemi-J (2008). The works of Telatar et. al. (2007), Bautista and Maveyraud-Tricoire (2007) and Chen and Shen (2015) use M-S models to derive two regime states from the data, namely, low and high capital mobility states. The two studies of Telatar et. al. (2007) and Chen and Shen (2015) confirm that most EU members have transitioned from low capital mobility to high capital mobility during periods corresponding to the creation of the EMU in 1994 whereas Bautista and Maveyraud-Tricoire (2007) draw similar sentiments for Asian economies which are found to have transitioned from low to high states of capital mobility during periods corresponding to the Asian financial crisis of 1998-1999. Aka (2007) uses the M-S VAR model to examine causality effects in to West African neighbouring countries (Ivory Coast and Ghana). The authors finds that investment granger causes savings for Ivory Coast in low volatility regime and vice-versa in high volatility regime whereas the author discovers no significant causal relations between the variables for Ghana in either of the two regimes.

Using PTAR and PSTR models respectively, Ho (2003) and Fouquau et. al. (2008) segregate their empirical data into two regimes corresponding to low and high share of GNP and the authors discover that the savings-retention coefficient for European countries becomes larger (i.e. less capital mobility) as the relative GNP share becomes larger, hence supporting the country-size argument of Murphy (1984). Moreover, Fouquau et. al. (2008) find that two other threshold candidates (i.e. degree of openness, country-size and current account ratios) can account for regime-switching behaviour in the savings-retention coefficient for Euro economies. Closer in nature to our study, Di Iorio and Fachin (2014) employ a bootstrap panel cointegration model with regime shifts of Gregory and Hansen (1996) whereas Dursun and Abasiz (2014) employ both the traditional threshold cointegration model of Hansen and Seo

(2002) as well as the single-break regime shift cointegration models of Gregory and Hansen (1996) and the double-break regime shift cointegration model of Hatemi-J (2008). On one hand, Di Iorio and Fachin (2014) find that there are no cointegration relations between savings and investment for a majority of OECD countries regardless of whether or not a structural break is accounted for in the cointegration relation. On the other hand, Dursun and Abasiz (2014) establish that when one break point is placed in the cointegration model for Turkish data, the savings-retention coefficient is close to unity hence validating the Feldstein-Horioka puzzle whereas when two structural break points are used the puzzle disappears such that there exists high international capital mobility after the Turkey financial crisis of 1994 to 1995. All-in-all, the studies of Di Iorio and Fachin (2014) and Dursun and Abasiz (2014) emphasize the importance of accounting for regime-switching cointegration behaviour in correcting the Feldstein-Horioka puzzle for both advanced and emerging economies.

3 EMPIRICAL SPECIFICATION AND MODELNIG TECHNIQUES

Denoting I, S and Y as investment, savings and national income, respectively, Feldstein and Horioka (1980), note that the savings-investment relationship can be represented by the following empirical long-run regression:

$$\frac{I}{Y} = \beta_0 + \beta_1 \frac{S}{Y} + \varepsilon_t \qquad (1)$$

Where (I/Y) and (S/Y) represents the investment and savings share in national income, respectively, and ε_t is a well-behaved error term. According to Feldstein and Horioka (1980) when β_1, the savings-retention coefficient, equals or is close to unity then this implies that an economy has low poor capital mobility such that it more-or-less resembles a financial autarky economy (with $\beta_1 = 1$ implying imperfect capital mobility or complete financial autarky).

Conversely when β_1 approaches zero then an economy exerts greater international capital mobility with a coefficient $\beta_1 = 0$ implying perfect international capital mobility.

According to the classic Engle and Granger (1987) theorem, in order to avoid the classic problem of spurious regression commonly associated with OLS estimates, the individual time series under investigation should be first difference stationary processes whereas the error term should be stationary with a zero mean. From equation (1), testing for cointegration can be achieved by the running the following regression of the error term:

$$\varepsilon_t = \rho\varepsilon_t + \xi_t, \quad \xi_t \sim \text{iid} \tag{2}$$

And thereafter, test the null hypothesis of the null hypothesis of no cointegration (i.e. -2 > ρ > 0) against the alternative of symmetric cointegration effects (i.e. -2 < ρ < 0). The Granger representation theorem guarantees that, if ρ is significantly different from zero, then equations (1) and (2) jointly imply the existence of the following error correction model (ECM) specifications:

$$\Delta\frac{I}{Y_t} = \alpha_0 + \gamma_1\xi_{t-1} + \sum_{i=1}^{n}\Phi_{i1}\Delta\frac{I}{Y_{t-i}} + \sum_{i=1}^{n}\psi_{i1}\Delta\frac{S}{Y_{t-i}} + \mu_{1t} \tag{3}$$

$$\Delta\frac{S}{Y_t} = \alpha_0 + \gamma_2\xi_{t-1} + \sum_{i=1}^{n}\Phi_{i2}\Delta\frac{I}{Y_{t-i}} + \sum_{i=1}^{n}\psi_{i2}\Delta\frac{S}{Y_{t-i}} + \mu_{1t} \tag{4}$$

Where ξ_{t-1} are the long-run error correction term whose coefficient, γ_i, is expected to be negative yet bounded within negative one (i.e. -1 < γ_i < 0) and provides a measure of the periodic rate of equilibrium correction in the face of a shock to the time series; Φ_i and ψ_i are the short-run dynamic coefficients and μ_{it} is a well-behaved disturbance term. Enders and Granger (1998) as well as Enders and Siklos (2001) all argue that the symmetric cointegration

tests may exert low power and thus the error correction representations may be misspecified if actual steady-state adjustment is indeed asymmetric. The authors suggest the following threshold autoregressive (TAR) cointegration regression as an alternative specification for the cointegration model represented in equation (2):

$$\varepsilon_t = I_t \rho_1 \varepsilon_{t-1} + (1 - I_t)\rho_2 \varepsilon_{t-1} + \eta_t \tag{5}$$

Where ρ_1 and ρ_2 are threshold error term coefficients such the sufficient condition for the stationarity of ε_t are ρ_1, $\rho_2 < 0$ and $(1 + \rho_1)(1 + \rho_2)$; η_t is disturbance term with properties $N(0, \sigma^2)$ and I_t is the Heaviside indicator function which assumes the following form:

$$I_t = \begin{cases} 1 \ if \ \varepsilon_{t-1} \geq \tau \\ 0 \ if \ \varepsilon_{t-1} < \tau \end{cases} \tag{6}$$

The TAR Heaviside indication function specified in equation (6) depends on the level of the lagged equilibrium error term, ε_{t-1}, and the unknown threshold value τ. Enders and Siklos (2001) and Caner and Hansen (2001) propose an alternative specification in which the Heaviside indicator function (M_t) depends on the lagged changes of ε_{t-1} such that momentum is given more to one side than the other (i.e. MTAR model):

$$M_t = \begin{cases} 1 \ if \ \Delta\varepsilon_{t-1} \geq \tau \\ 0 \ if \ \Delta\varepsilon_{t-1} < \tau \end{cases} \tag{7}$$

From equations (5) through to (7), the parameters of empirical interest are β_1, ρ_1, ρ_2 and τ. The estimation procedure begins with obtaining the unknown threshold value τ and this is

based on Chan's (1993) grid-search method which entails arranging the potential thresholds in ascending order and discarding the lowest and highest 15 percent of the observations. Thereafter the threshold regression is estimated using each potential threshold value of ε_{t-1} and the consistent or true threshold estimate is determined as the one which yields the lowest residual sum of squares. Once the true value of τ is obtained, backward substitution is performed in order to obtain the threshold error coefficient values of ρ_1 and ρ_2 as well as the savings-retention coefficient β_1.

As a means of validating symmetric ad asymmetric cointegration effects, Enders and Granger (1998) and Enders and Silkos (2001) propose the testing of two cointegration hypotheses. The first test involves testing the null hypothesis of no cointegration effects (i.e. $\rho_1 = \rho_2 = 0$) against the alternative of convergence effects (i.e. $\rho_1 \neq \rho_2 \neq 0$). The F-statistics used to test this hypothesis are denoted as F-Max* for the TAR model and F-Max*(M) for the MTAR model. Once the null hypothesis of no cointegration effects is rejected, one then proceeds to test the second null hypothesis of linear cointegration effects (i.e. $\rho_1 = \rho_2$) against the asymmetric convergence alternative(i.e. $\rho_1 \neq \rho_2$). These statistics are denoted as Φ^* and Φ^*(M) for the TAR and MTAR model, respectively. The critical values of the aforementioned hypotheses tests are reported in Enders and Siklos (2001).

According to the Granger representation theorem, the existence of cointegration implies the existence of an error correction mechanism between the time series variables. Once the null hypothesis of linear cointegration effects is rejected, Enders and Siklos (2001) suggest the estimation of the following threshold error correction mode (TECM):

$$\Delta \frac{I}{Y}_t = \alpha_0 + \sum_{i=1}^{n} \Phi_i \Delta \frac{I}{Y}_{t-i} + \sum_{i=1}^{n} \psi_i \Delta \frac{S}{Y}_{t-i} + I_t \gamma_1 \xi_{t-1} + (1 - I_t) \gamma_2 \xi_{t-1} + \mu_{1t} \quad (8)$$

$$\Delta \frac{S}{Y}_t = \alpha_0 + \sum_{i=1}^{n} \Phi_i \Delta \frac{I}{Y}_{t-i} + \sum_{i=1}^{n} \psi_i \Delta \frac{S}{Y}_{t-i} + I_t \gamma_1 \xi_{t-1} + (1 - I_t) \gamma_2 \xi_{t-1} + \mu_{2t} \quad (9)$$
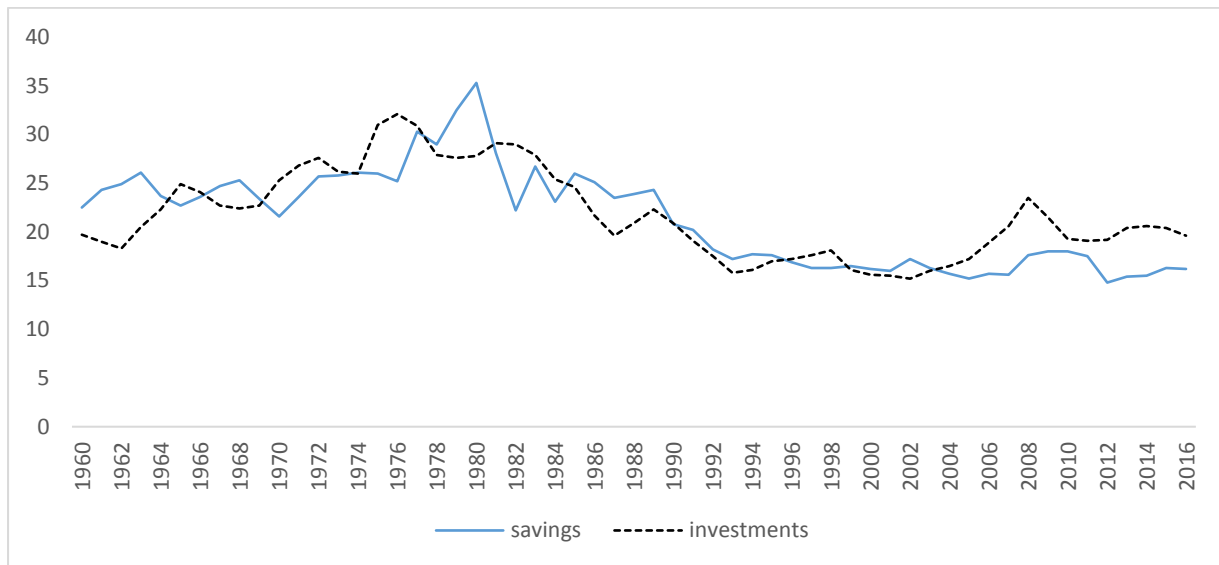
Where Δ is a first difference operator, ξ_{t-1} is the error correction term, I_t is the indicator function which assumes the TAR and MTAR identities defined in equations (6) and (7), respectively; and μ_{ti} is a well behaved error process. From the TEC regressions (8) and (9), three hypotheses are tested for. Firstly, we test the null of no threshold error correction as $\gamma_{1i} = \gamma_{2i}$, against the alternative of asymmetric error correction effects (i.e. $\gamma_{1i} \neq \gamma_{2i}$). This hypothesis is tested using a F-statistic denoted as $F[H_{30}]$. Secondly, we granger test the null of the investment rate not causing savings rate (i.e. $\Phi_i = 0$). Lastly, we granger test the null hypothesis of savings rate not leading the investment rate (i.e. $\psi_i = 0$). All aforementioned hypotheses are tested using F-statistics.

4   DATA AND EMPIRICAL RESULTS

4.1 Data description and unit root tests

The time series data employed in our study are the gross fixed capital formation as ratio of GDP (I/Y) and the ratio of gross savings to GDP (S/Y). All data has been retrieved from the South African Reserve Bank online statistical database over a quarterly interval ranging from 1960:Q1 to 2016:Q4. The time series plot of the variables used over the study period is presented in Figure 1 below.

Figure 1: Savings-investment patterns in South Africa: 1960 to 2016

Note that over the entire study period, the savings and investment time series variables appear to more or less move together over time. From the early 1960's to the mid-1970's both savings and investment were on an upward trend until they reached their peaks in the mid-1980's. At the time, South Africa enjoyed heavy foreign direct investment in mining and manufacturing and consequentially high investment during this period coincided with increased savings in part because the high gold price and high corporate profitability lead to high rates of savings. The descent of the time series variables from the mid-1980's until the mid-1990's corresponds to the periods of disinvestment associated with sanctions placed on the South African economy as a component of the-then anti-Apartheid campaigns which resulted in massive capital reversals. During this period, low savings rate were primarily due to a deteriorating household savings which were not compensated for by an increase in government savings or corporate savings ratios (Bonga-Bonga and Guma, 2017). Further exacerbating the worsening economic conditions were the deteriorating manufacturing and mining industries, deteriorating global economic conditions of the 1980's, the major "brain drain" of the 1980's, the 1985 debt crisis as well as the severe drought period of 1992.

Following the democratic transition of 1994, the savings and investment variables began to stabilize albeit at historically low levels. Despite a number of fiscal policies programmes implemented between 1994 and 2004 (i.e. Reconstruction and Development Programme (RDP), Growth Employment and Redistribution (GEAR) and Accelerated and Growth Initiative for South Africa (ASGISA)) which aimed at correcting the social imbalances caused by the former Apartheid regime, these policies did little to improve national savings shares in GDP. The reduction in national savings experienced the years subsequent to 1994, was mainly a result of deteriorating private corporate and household savings. On the other hand, due to the privatization programmes embedded in the RDP and GEAR strategies, investment share in GDP slightly improved in the post-democratic period of 1994 and yet never returned to their previously high levels experienced in the late 1970's.

Notably, between 2003 and 2008, the investment share of GDP drastically improved whilst the share of savings in GDP remained at relatively low rates. One of the factors influencing much of this boost in investment during this period can be attributed to the 2004 announcement of the South Africa's historic hosting of the Soccer World Cup in 2010 (Phiri, 2015). On the other hand, little improvement was exerted on the savings rate during this period because household savings relatively to disposable income greatly deteriorated (Odhiambo, 2009). Following the advent of the global financial crisis of 2008 as well as the resulting global recession period of 2009, both savings and investment shares in GDP began to deteriorate and despite a number of policy initiatives (New Growth Path (NGP) and New Development Plan (NDP)) put in place to foster infrastructure investment, seemingly the time series have not fully returned to their pre-crisis figures.

Table 1: Descriptive statistics of the time series

	sample periods					
	full sample		pre-crisis		post-crisis	
	I/Y	S/Y	I/Y	S/Y	I/Y	S/Y
Mean	21.77	21.42	22.04	22.24	20.19	16.55
std. dev.	4.56	5.35	4.85	4.85	1.31	1.35
observations	228	228	195	195	33	33

Notes: All computations are the authors own. Std. dev. represents the standard deviation

Complementary to Figure 1, we present some descriptive statistics for the time series variables in Table 1. Collectively, we observe that the characteristics of the descriptive statistics for the time series changes from the pre-crisis to the post-crisis periods. For instance, the pre-crisis period savings averages 22.24% of GDP slightly exceed the investment averaged 22.04% of GDP over the same period whereas for the post-crisis period, investment averages of 20.00% of GDP exceed the savings averages of 16.50% of GDP. In also considering the standard deviation statistics reported in Table 1, note that the pre-crisis values of 4.85 and 4.86 obtained for savings and investment respectively, far exceeds the post crisis values of 1.31 and 1.35 for the savings and investment variables, respectively. We thus conclude that both the averages and volatility measures for the respective time series variables was much higher in the pre-crisis period when compared to that experienced in the post-crisis period.

As is standardized in the empirical literature, it is imperative that we also examine the integration properties of the time series variables prior to utilizing them in our cointegration analysis. The ADF unit root test is the most frequently utilized procedure employed in previous empirical studies in determining stochastic trends in the time series (Gulley (1992), Coakley and Kulasi (1997), Moreno (1997), Shibata and Shintani (1998), Sarno and Taylor (1998), Hussein (1998), Levy (2000), Coiteux and Olivier (2000), Sinha (2002), De Vita and Abbott (2002), Ho (2002), Pelagidis and Mastroyiannis (2003), Abbott and De Vita (2003), Ozmen and Parmaksiz (200), Kim et. al. (2005), Adedeji and Thonton (2008), Singh (2008), Khundrakpam and Ranjan (2010), Narayan and Narayan (2010), Ketenci (2013), Konya (2015), Behera (2015) and Barros and Gil-Alana (2015)). In pursuit of these authors, we also apply the ADF test to our observed time series inclusive of (i) a drift; and (ii) a trend, and report our findings in Table 2.

As can be observed from Table 2, both time series fail to reject the unit root null hypothesis for all conducted tests regardless of whether the tests are performed with a drift or trend. Only when first differences are applied to the time series do we find that the variables

become stationary process thus rendering the time series as first difference stationary (i.e. I(1) processes). Notably, the finding of savings and investment being mutually I(1) time series is consistent with that obtained in previous empirical study of Behera (2015) which includes South Africa in the panel dataset. Overall, these findings permit us to proceed with our cointegration analysis without fear of obtaining spurious results in our regression analysis.

Table 2: ADF unit root test results

time series	(I/Y)		(S/Y)	
test type	drift	trend	drift	trend
sample period				
full	-1.75	-2.44	-1.69	-3.09
	(-6.64)***	(-6.65)***	(-12.22)***	(-12.20)***
pre-crisis	-1.58	-1.91	-1.64	-2.87
	(-5.47)***	(-5.43)***	(-11.36)***	(-11.36)***
post-crisis	-1.89	-1.96	-1.77	-1.34
	(-2.60)*	(-2.69)*	(-3.21)**	(-3.33)*

Notes: Significance codes: '***', '**' and '*' denote the 1, 5 and 10 percent critical levels, respectively. The unit root test statistics for first differences are reported in parentheses ().

4.2 Empirical results

Following the confirmation of stochastic trends in all our employed time series data, we carry out our empirical analysis in the following three modelling stages:

I.   Firstly, we perform three hypotheses tests for symmetric cointegration, asymmetric cointegration and threshold error correction effects for both TAR and MTAR specifications corresponding to data samples representative of the pre-crisis, the post-crisis and the full sample periods.

II. Secondly, we estimate threshold cointegration and error correction effects for the data samples which manage to reject all three tested hypotheses in stage I. For the case of the remaining data samples which manage to pass the hypotheses tests for symmetric cointegration we estimate linear cointegration and corresponding symmetric error correction models.

III. Lastly, we conducted causality tests for all estimated models carried out in stage II of our modelling procedure.

Table 3 below, presents the empirical results of the first stage of our modelling procedure. Beginning with the results reported for our TAR specifications, we obtain t-Max* statistics of 13.83, 13.28 and 21.93 for the full sample, pre-crisis and post-crisis periods, respectively. These statistics mutually reject the no cointegration null at all critical levels. However the $\phi*$ statistics produced for the full and pre-crisis periods are 0.53 and 0.27, respectively, and these statistics fail to reject the null of TAR cointegration for these two sub-samples. On the other hand, the $\phi*$ statistic associated with the post crisis period produces a highly significant figure of 11.39, hence rejecting the null of linear cointegration effects in favour of TAR cointegration for the post-crisis. And yet, in further testing for significant TEC effects for the post periods, we observe that the associated $F[H_{30}]$ statistic of fails to reject the null of TEC effects. Collectively, these results imply that TAR models fail to significantly capture the long-run and short cointegration dynamics between savings and investment in South African time series data.

The results produced from the MTAR models prove to be more encouraging/optimistic in nature. For instance, we find t-Max*(M) statistics of 17.71, 16.54 and 5.84 for the full, pre-crisis and post-crisis periods, respectively, all which are significant at critical levels of at least 5 percent. Similarly, the $\phi*$ statistics produce values of 8.12, 6.67 and 4.01 which are all significant at critical levels of at least 5 percent. Concerning the $F[H_{30}]$ statistic, we find values of 5.05 and 3.11 for the full and pre-crisis periods, respectively, and both these statistics are significant at 5 and 10 percent, respectively. Conversely, the $F[H_{30}]$ statistic associated with the post-crisis period is 0.02 which is insignificant and hence we cannot reject the no TEC effects hypothesis. All-in-all, we conclude that the MTAR cointegration and TEC models can be used to model the steady-state relationship between savings and investment in South Africa for the full and pre-crisis sample periods, whereas a linear cointegration framework is deemed to be more suitable for modelling the dynamic relationship for the post-crisis periods.

Table 3: Threshold cointegration and error correction tests

model type	Statistic	full sample	pre-crisis	post-crisis
	t-Max*	13.83	13.28	21.93
		(0.00)*	(0.00)***	(0.00)***
TAR	ϕ*	0.53	0.27	11.39
		(0.47)	(0.60)	(0.00)***
	F[H_{30}]	2.88	0.24	0.78
		(0.09)*	(0.62)	(0.39)
	t-Max*(M)	17.71	16.54	5.87
		(0.00)**	(0.00)***	(0.01)**
MTAR	ϕ*(M)	8.12	6.67	4.01
		(0.00)***	(0.01)**	(0.03)**
	F[H_{30}]	5.05	3.11	0.02
		(0.02)**	(0.08)*	(0.90)

Notes: Significance codes: '***', '**' and '*' denote the 1, 5 and 10 percent critical levels, respectively.

Having tested for threshold cointegration and error correction effects, we proceed to model and estimate the significant models identified in the first stage of our modelling process. To recall, we only found significant MTAR models for the full sample and pre-crisis whereas linear cointegration models are more suitable for the post-crisis period. Panel A of Table 4 reports the estimates of the savings-retention coefficients, β_1, as well as the threshold error coefficient parameters, ρ_1 and ρ_2, which measure the speed of adjustment back to steady-state equilibrium after positive and negative shocks to the current account, respectively. Notice from Panel A that all β_1 coefficients produce estimates of 0.59, 0.64 and 0.22 for the full, pre-crisis and post-crisis sample periods, respectively, and these estimates are consistent across both linear and threshold models. The first two savings-retention estimates are significant at all critical levels whereas the last estimate produces an insignificant value.

Generally, the aforementioned results can be considered highly credible since they are quite comparable to other savings-retention coefficients obtained in previous studies for African economies (i.e. de Wet and Van Eyden (2005); Payne and Kumazawa (2005); Adedeji and Thornton (2006); Cooray and Sinha (2007); Cyrille (2010); Bangake and Eggoh (2011); Adams et. al. (2016) and Raheem (2017)). Overall, our savings-retention estimates imply that the entire investigated period has been characterized by moderately high levels of international

capital mobility even though it appears that international capital flows became more/increasingly mobile subsequent to the 2007-2008 financial crisis. Further note that our results are contrary to those realized in the works of Katsimi and Zoega (2016) and Morley (2016) who find higher savings-retention coefficients in the post-crisis periods for Euro economies. We consider our results as being plausible since, as noted by Ostry et. Al. (2010), there has been a surge of capital inflows back to emerging markets as the global economy began to recover from the financial crisis although these capital flows are more short-term or speculative in nature.

In turning to the threshold error term estimates, we note that for both the full-sample and pre-crisis periods the threshold error coefficient estimates satisfy the convergence condition of $\rho_1, \rho_2 < 0$ and $(1 + \rho_1)(1 + \rho_2)$. We obtain ρ_1 estimates of -0.41 and -0.46 for the full and pre-crisis periods, respectively, whilst the ρ_2 estimates are -0.14 and -0.17, for the full and pre-crisis periods, respectively. Note that in both sample periods, $|\rho_1| > |\rho_2|$, a result which implies that positive shocks to the equilibrium are eradicated quicker than negative shocks. In other words, during periods where current account is improving, disequilibrium caused by savings and investment is easier to correct in comparison to periods where the current account is worsening. Thus greater persistence is observed in deteriorating current accounts when compared to improving ones.

Table 4: MTAR-TEC estimates

panel A	long-run cointegration model estimates				
sample period	full sample		pre-crisis		post-crisis
	linear	MTAR	linear	MTAR	linear
β_0	9.08	9.08	7.79	7.79	16.57
	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.00)***
β_1	0.59	0.59	0.64	0.64	0.22
	(0.00)***	(0.00)***	(0.00)***	(0.00)***	(0.21)
τ		1.748		1.979	
ρ_1		-0.41		-0.46	
		(0.00)***		(0.00)***	
ρ_2		-0.14		-0.17	
		(0.00)***		(0.00)***	

panel B	error correction estimates									
sample period	full sample				pre-crisis				post-crisis	
	linear		MTAR		linear		MTAR		linear	
dependent variable	$\Delta\frac{i}{y}$	$\Delta\frac{s}{y}$	$\Delta\frac{i}{y}$	$\Delta\frac{s}{y}$	$\Delta\frac{i}{y}$	$\Delta\frac{s}{y}$	$\Delta\frac{i}{y}$	$\Delta\frac{s}{y}$	$\Delta\frac{i}{y}$	$\Delta\frac{s}{y}$
ξ_{t-1}	-0.03	0.23			-0.02	0.26			-0.09	0.25
	(0.09)*	(0.00)			(0.07)*	(0.00)***			(0.09)*	(0.17)
$\xi_{t-1} > \tau$			-0.02	0.19			-0.01	0.23		
			(0.36)	(0.00)***			(0.83)	(0.00)***		
$\xi_{t-1} > \tau$			-0.08	0.53			-0.11	0.53		
			(0.06)*	(0.00)***			(0.02)**	(0.00)***		
R^2	0.41	0.09	0.54	0.25	0.04	0.17	0.16	0.28	0.03	0.14
DW	2.21	1.96	2.05	1.92	1.94	2.06	1.95	2.11	1.97	2.12
LB	0.01	0.02	0.00	0.00	0.00	0.02	0.02	0.00	0.01	0.00

Notes: Significance codes: '***', '**' and '*' denote the 1, 5 and 10 percent critical levels, respectively. DW and LB denote the Durbin Watson and Ljung Box test statistics, respectively, and both test statistics indicate that absence of serial correlated residuals.

Panel B of Table 4, present the error correction estimates associated with both linear and nonlinear cointegration models. As is evident from our findings, regardless of whether a linear or threshold model is considered, we find significant error correction estimates with the correct negative sign only when the investment variable is the driving variable in the system. In particular, we find error correction estimates of -0.03, -0.02 and -0.09 for the full, pre-crisis and post-crisis periods, respectively, thus implying that 3%, 2% and 9% of disequilibrium are corrected each quarter during these respective periods.

On the other hand, we find significant error correction estimates with the correct negative sign for the threshold models only when investment is the driving variable and the error correction estimate is above it's estimated threshold (i.e. $\xi_{t-1} > \tau$). In particular, we obtain ξ_{t-1} estimates of -0.08 and -0.11 for the full and pre-crisis periods thus implying that 8% and 11% of disequilibrium are corrected each quarter during these respective periods. Overall, these results obtained from the threshold error correction model are comparable with those obtained from the linear error correction models in that there are low levels percentage correction of steady state deviations each quarter (i.e. between 2% and 11%).

Table 5: Causality tests

model type	sample period	causality direction	
		savings to investment	investment to savings
linear	full	3.50 (0.01)**	1.76 (0.16)
	pre-crisis	3.21 (0.01)**	2.02 (0.11)
	post-crisis	2.09 (0.12)	1.43 (0.26)
MTAR	full	5.23 (0.00)***	1.46 (0.23)
	pre-crisis	4.03 (0.01)**	1.36 (0.26)

`Notes: Significance codes: '***', '**' and '*' denote the 1, 5 and 10 percent critical levels, respectively.

In the final stage of our empirical modelling process, we examine causality effects between the time series across both linear and threshold cointegration models and the results of this empirical exercise are reported in Table 5. As is shown for the linear models, the null hypothesis testing no causal effects from savings to investment produce test statistics of 3.50, 3.21 and 2.09 for the full, pre-crisis and post-crisis periods, respectively. Note that the first two reported statistics reject the null hypothesis at critical levels of at least 5 percent whereas the last statistic cannot reject the null at all significance levels. Contrary, when testing the null of no causal effects from investment to savings, we obtain F-statistics of 1.76, 2.02 and 1.43.

In turning to the causality results obtained from our threshold models, we obtain F-statistics of 5.23 and 4.03 for the full and pre-crisis periods and these statistics reject the null that savings does not granger cause investment at all significance levels. Conversely, when testing the null of no causality effects from investment to savings, our obtained F-statistics estimates of 1.46 and 1.36 for the full and pre-crisis periods, respectively, and these statistics fail to reject the null hypothesis at all critical levels.

In general, our causality test results the threshold models imply that savings granger caused investment in both the full and pre-crisis periods. Note that the causality results obtained from our threshold models concur with those obtained from the linear causality analysis. Moreover, the common finding of causality running from savings to investment in the pre-crisis period and full samples has been iterated in the studies of Argimon and Roldan (1994), Sinha (2002), Brahmasrene and Jiranyakul (2009) and Josic and Josic (2012) albeit for European Union countries, Asian countries, Thailand and Croatia, respectively. On the other hand, the finding of no causal effects between savings and investment in the post-crisis which is in line with findings of Grullon (2016) for 4 developing countries.

5   CONCLUSIONS

In deviating from the traditional belief of a linear steady-state analysis for the Feldstein-Horioka puzzle, this current paper examines the asymmetric cointegration relationship between savings and investment for South Africa within a MTAR framework for the quarterly periods 1960:Q1 – 2015:Q4. We consider our paper worthwhile due to the scarcity of empirical literature on the subject matter for South Africa as an individual economy, with the study of Gil-Alana et. al. (2016) being the only priori exception. In varying from these authors who find a breakpoint during the financial deregulation period of the 1980's, we examine the changing dynamics of the savings-investment relationship with respect to a more recent event, the 2007-2008 global financial crisis. Interestingly enough we obtain savings-retention coefficients of

0.64 (significant) and 0.22 (insignificant) for the pre and post crisis periods, respectively, whilst for the full sample we obtain a coefficient of 0.59 (significant). Collectively, these results imply increased international capital mobility in transcending from the pre-crisis to post-crisis periods. This result may be explained the repatriation of capital investment to safe haven assets in face of the recent financial crisis.

In terms of steady state dynamics, we find asymmetric convergence effects between savings and investment for the pre-crisis and full samples in which disequilibrium from the steady state is corrected quicker for improving current accounts whereas for deteriorating current accounts such adjustment is more persistent in nature. However, during the post-crisis period both worsening and improving current accounts are corrected symmetrically. Moreover, our causality analysis indicates that savings led to investment during the pre-crisis periods whereas during the post-crisis no causality exists between the variables. The particular finding of no causality between the variables further emphasizes the notion of increased mobility in the post-crisis period. Our study thus serves as a caution to policymakers to adopt effective capital management techniques and abandon the notion of further relaxing exchange controls as a means of attracting investment to finance the current budget. In an environment already characterized by high capital mobility and a downgraded international credit rating, such increased liquidity will most likely be channelled to increased consumption, imports and capital flight, which in turn, could exert adverse effects on already fragile exchange rates, economic growth and employment levels.

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