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# **STOCK RETURN PREDICTABILITY IN SOUTH AFRICA: AN ALTERNATIVE APPROACH**

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## **ABSTRACT**

There is considerable debate internationally as to whether share returns are predictable. The limited evidence in South Africa (Gupta and Modise, 2012a, b and 2013) reveals that valuation ratios have no forecasting power but the Treasury bill rate, term spread and money supply have been found to be able to predict share returns at a relatively short horizon. In this study, the consumption aggregate wealth ratio of Lettau and Ludvigson (2001) is applied to South African share returns to assess its forecasting power using in-sample tests over both short and long horizons. The forecasting power of this composite variable is compared to a number of traditional variables. Similarly to the developed market evidence, the results indicate that the consumption aggregate wealth ratio is a significant predictor of returns and combined with the term spread, can explain a substantial component of the variation in future share returns. The implications of these findings for practitioners and policy makers are discussed.

## **1. Introduction**

The efficient market hypothesis maintains that share returns are not predictable using publicly available information such as valuation ratios or macroeconomic variables. If share returns can be forecasted, it suggests that markets are not fully efficient and investors can earn abnormal returns. An alternative view contends that predictability reflects the rational response of investors to time-varying investment opportunities which vary with cycles in risk aversion. In this context, macroeconomic variables are likely to reflect changing patterns and therefore play an important role in forecasting future share returns (Lettau and Ludvigson, 2001). Given that there is substantial evidence that share prices not only act as a leading indicator of output and inflation but

that there are also spillover effects from the share market to the real sector (Gupta and Modise, 2012b), obtaining accurate forecasts of changing cycles in risk can enable policy makers to devise and implement appropriate policies to minimise the impacts of market downturns. Irrespective of the theoretical stance on the driving force behind the forecastability of returns, the importance thereof for making money in financial markets and/ or more effective policy decisions has prompted a resurgence of research on this topic.

The evidence regarding predictability is mixed with only limited evidence that financial ratios such as dividend-to-price and earnings-to-price and other measures such as short-term interest rates and the term spread can forecast share returns (Rasmussen, 2006; Lettau and Ludvigson, 2010). Similar weak forecasting results have been obtained for South Africa (Gupta and Modise, 2012a, 2012b, 2013). However, Cochrane (2008) maintains that this relatively weak evidence may not reflect that share returns are not predictable but rather than these traditional forecasting measures are poor. Lettau and Ludvigson (2001) proposed an alternative measure to forecast share returns, known as the consumption aggregate wealth ratio (CAY). They found that CAY could explain future short- and long-run aggregate real share returns better than any of the traditional measures in the United States (U.S). Ioannidis *et al.* (2006), Gao and Huang (2008) and Sousa (2012) have confirmed the success of CAY in predicting share returns in other developed markets.

No study, however, has examined whether share returns in South Africa can be predicted using CAY and whether this variable is more successful than those identified by Gupta and Modise (2012b and 2013). In fact, the research on CAY has been limited to developed markets only. Thus, the purpose of this study is to ascertain whether CAY can be used to predict share returns in the emerging market of South Africa.

## **2. Literature Review**

### **2.1 The Link between the Macroeconomy and the Stock Market**

The relationship between macroeconomic factors and the stock market has received extensive attention from financial economists. If the price of a share represents the discounted present value of expected future cash flows then it must ultimately be affected by real economic activity measured by gross domestic product or industrial

production (Binswanger, 2000). Consequently, it has been argued that because share prices reflect investors' expectations of future economic conditions stock market movements can serve as a leading indicator of economic activity (Moolman and Du Toit, 2005). Chen *et al.* (1986), however, contended that this relationship is not unidirectional but that equally economic factors affect the stock market by impacting on discount rates and future dividends. Similarly, Fama (1981) argued that stock price changes must be linked to shocks to economic variables that affect the consumption and investment opportunity set. This suggests that by modelling the appropriate interactions between macroeconomic variables and stock market returns it could be possible to predict stock market movements (Moolman and Du Toit, 2005). Obviously being able to predict movements in the stock market represents both a theoretical challenge to the efficient market hypothesis (although several authors including Fama and French, 1988a and Balvers *et al.*, 1990 have demonstrated that predictability is not necessarily inconsistent with market efficiency in the context of intertemporal models) and an applied opportunity to earn abnormal returns. As a result, the explanatory power of macroeconomic variables to predict stock returns has received extensive empirical attention (Ferreira and Santa-Clara, 2011; Rapach *et al.* 2013).

## **2.2 Share Return Predictability with Traditional Financial Variables**

The evidence regarding predictability is mixed – not only in terms of whether share returns are predictable, but which variables can be used to predict returns and over what time horizons. In the U.S, early tests found that the dividend-to-price and earnings-to-price ratios were able to forecast future returns (Fama and French, 1989; Hodrick, 1992). However, Lamont (1998) demonstrated that the dividend-to-price ratio had greater predictive power than earnings-to-price. More recently, Ang and Bekaert (2007) and Lettau and Ludvigson (2010) confirmed the ability of the dividend-to-price ratio to predict excess returns over short and long horizons, but only when combined with a measure of the short-term interest rate. The results of Ang and Bekaert (2007), Lettau and Ludvigson (2001) and Lettau and Ludvigson (2010) mirrored the results of earlier studies by Hodrick (1992) and Lamont (1998) that short-term interest rates were able to predict returns, especially at short horizons. Keim and Stambaugh (1986) showed that

the term and default spreads could predict future returns, whereas Lettau and Ludvigson (2010) found that these variables had little forecasting power when combined with the dividend-to-price ratio and the short-term interest rate.

Very little research had been conducted on return predictability using fundamental information in South Africa until a series of studies by Gupta and Modise (2012a, b and 2013). Their study on valuation ratios revealed that the dividend-to-price and earnings-to-price metrics had no predictive power over the short- or long-run (Gupta and Modise, 2012a). Using financial variables, Gupta and Modise (2012b) found some evidence of predictability using the Treasury bill rate and term spread, but the forecasting power of these two variables was relatively weak. Finally, Gupta and Modise (2013) examined the use of macroeconomic variables and found that various measures of the interest rate and money supply had some success in forecasting future period share returns. Thus the South African findings appear to be similar to those documented for the U.S.

## **2.2 Share Return Predictability: An Alternative Measure**

Lettau and Ludvigson (2001) proposed an alternative variable to predict share returns, the consumption aggregate wealth ratio (CAY), which measures the transitory deviation from the long-run relationship between consumption, asset wealth and labour income. Lettau and Ludvigson (2001) found that CAY was able to explain approximately 9% of the variation in one-period ahead future returns. The inclusion of traditional forecasting variables resulted in only a marginal increase in the adjusted R-squared ( $\bar{R}^2$ ) of the forecasting regression to 10%, with the relative Treasury bill yield significant but the earnings-to-price ratio, dividend-to-price ratio and the term spread were insignificant. CAY had a significant positive relationship with expected future returns indicating that if returns were expected to decrease in the future, investors who desired to smooth out consumption patterns over time allowed consumption to temporarily decrease below its long-term relationship with asset wealth and labour income to protect future consumption from lower returns. The opposite was true if returns were expected to increase in the future (Lettau and Ludvigson, 2001).

Lettau and Ludvigson (2010) demonstrated that CAY also has predictive power over longer horizons. Moreover, Hodrick and Zhang (2001) also showed that the predictive

power of CAY far exceeded that of typical macroeconomic indicators - industrial production and gross national product. Out of country evidence in support of the forecasting power of CAY has also been obtained, such as that of Ioannidis *et al.* (2006) for Australia, Canada and the United Kingdom (U.K). Gao and Huang (2008) and Sousa (2012) confirmed this evidence for the U.K; however, Gao and Huang found CAY to be less successful in predicting returns in the Japanese market.

Brennan and Xia (2005) contend that the forecasting results of CAY are biased upwards as data which is not in the investor's information set at the time of the forecast is used to predict share returns (that is, CAY is estimated over the full time period of the studies and then used to predict returns during the same period). However, Lettau and Ludvigson's (2001) out-of sample tests and further tests in Lettau and Ludvigson (2004) of CAY estimated using an alternative procedure dispute this point. Moreover, Lettau and Ludvigson (2005) argued that the traditional method of computing CAY is correct from an econometric perspective, as cointegration requires that the full sample of data is used to estimate the true long-run relationship between the variables that would have been known to the representative investor; bias would only arise if information was ignored.

### **3. Methodology and Data**

In light of the success of CAY in predicting share returns in developed markets, the purpose of this study was to examine whether this composite macroeconomic variable has the same forecasting ability in emerging markets, with particular attention on the Johannesburg Stock Exchange (JSE).

#### **3.1 Data**

Quarterly data for the period 1990:03 to 2013:01 was used, with the frequency of the data necessitated by the use of macroeconomic data which was not available at any higher frequency. Consumption was measured as final expenditure by households on non-durable goods and services, with consumption on durable goods excluded as the theory applies to the flow of consumption whereas expenditure on durable goods

represents an addition to stock (Hassan and van Biljon, 2012). The seasonally adjusted current price series was utilised so as to remove the effects of predictable seasonal patterns, which are particularly relevant to consumer consumption, which tends to peak at year-end. This series was obtained from the South African Reserve Bank (SARB).

For labour income, the SARB's seasonally adjusted compensation for residents series was used, but net social benefits, net other current transfers, miscellaneous current transfers and taxes were also accounted for. To incorporate the transfer payments and taxes, where only annual data was available, a cubic spline<sup>i</sup> was used to interpolate quarterly observations that were then used to compute the total labour income measure.<sup>ii</sup>

Li *et al.* (2011) measured asset wealth as the difference between total household financial and non-financial assets and liabilities. This same measure was obtained for South African households from the SARB, but again only annual information was obtained. However, the SARB also provides quarterly estimates of the ratio of net household wealth to gross domestic product. By making use of the appropriate gross domestic product series, a quarterly series for net household wealth was computed by multiplying the ratio by gross domestic product. The current price series for these three variables were adjusted to real prices using the consumer price index, obtained from Statistics South Africa. Thereafter, each series was converted to natural logs.

To assess the ability of CAY to predict share returns on the JSE, a measure of returns was needed. For this purpose, the J203 FTSE/JSE All Share Index was used to represent the market. The excess market return was computed by subtracting the quarterly risk-free rate, measured as the return on the three-month Treasury bill. Thereafter, the nominal excess market return was converted to a real return using the consumer price index.

Several other variables that have been found to have predictive power for share returns, both internationally and in South Africa, were also examined so as to be able to compare against the performance of CAY in forecasting future returns. The relative Treasury bill yield, term spread, dividend-to-price, earnings-to-price and the one period previous excess real market return were selected. The relative Treasury bill yield was calculated as the three-month Treasury bill yield less the 12-month moving average

(Rapach *et al.*, 2005). The term spread was measured as the difference between the long-term (10-year) government bond yield and the three-month Treasury bill yield (Lettau and Ludvigson, 2010). The dividend-to-price and earnings-to-price ratios for the All Share Index were obtained, but these series did not take into account seasonality in dividends and earnings. As such, the dividend-to-price and earnings-to-price series were multiplied by the All Share Index price to obtain the equivalent quarterly dividend and earnings values. Thereafter, the dividend-to-price and earnings-to-price ratios were computed to account for seasonality as follows

$$D/P_t = \ln(D_t^4) - \ln(S_t) \text{ and} \quad (1)$$

$$E/P_t = \ln(E_t^4) - \ln(S_t), \quad (2)$$

where  $D/P_t$  and  $E/P_t$  are the dividend-to-price and earnings-to-price ratios respectively at time  $t$ ,  $S_t$  is the nominal stock price and  $D_t^4$  is the four-quarter dividend moving average computed as the sum of the dividends in the current quarter and three preceding quarters (i.e.  $D_t^4 = \sum_{t-3}^t D_t$ ) (Ang and Bekaert, 2007:654).  $E_t^4$  is defined analogously. The conditioning variables were taken as real values as they are computed as ratios or the difference between two series such that the effect of inflation is cancelled out. The exception to this is the lagged excess market return which was converted to a real return. All of the predictor variables were normalised (by subtracting the mean and dividing by the standard deviation) to aid interpretation.

## 3.2 The Computation of CAY

### 3.2.1 The Theoretical Development of CAY

The intertemporal budget constraint of investors is as follows

$$W_{t+1} = (1 + R_{wt})(W_t - C_t), \quad (3)$$

where  $W_t$  is total wealth and  $R_{wt}$  are the gross returns to total wealth (Campbell and Mankiw, 1989). This budget constraint demonstrates that an investor's total wealth is determined by the total wealth invested in the previous period (i.e. that which is not consumed) grown by the total returns from investing the funds. Campbell and Mankiw (1989) derived a formulation for the log consumption wealth ratio from this budget constraint. To do this, they introduced logs and obtained a first-order Taylor series



expansion of equation 3 to impose linearity, and obtained an estimate for the log differenced aggregate wealth as

$$\Delta w_{t+1} \approx (r_{wt+1}) + (1 - \frac{1}{p_w})(c_t - w_t), \quad (4)$$

where  $\Delta w_{t+1}$  is the change in the log of wealth and where  $p_w$  is the steady-state ratio of invested to total wealth  $(W_t - C_t)/W_t$  (Lettau and Ludvigson, 2001). All variables in lowercase are measured in natural log. By solving this difference equation forward, taking expectations and imposing a transversality condition ( $\lim_{i \rightarrow \infty} p_w^i (c_{t+i} - w_{t+i}) = 0$ ), Campbell and Mankiw (1989) expressed the log consumption wealth ratio as

$$(c_t - w_t) \approx E \sum_{i=1}^{\infty} p_w^i (r_{wt+1} - \Delta c_{t+1}), \quad (5)$$

where  $(c_t - w_t)$  represents the consumption wealth ratio.<sup>iii</sup> Assuming that the returns to total wealth and the consumption growth rate are stationary, equation 3 implies that consumption and wealth, the two non-stationary variables (in their price formats) must be cointegrated (Lettau and Ludvigson, 2010). Drawing from Granger's representation theorem, equation 5 reveals that any deviations in this long-run relationship between consumption and wealth in the current period will lead to changes in the return to total wealth or consumption growth in the following period. This intertemporal relationship implies that the consumption-wealth ratio should be able to predict future values of either the returns to wealth or consumption growth rate (Lettau and Ludvigson, 2001).

The limitation with this specification of the consumption wealth ratio is that aggregate wealth is not directly observable. To overcome this limitation, Lettau and Ludvigson (2001) decomposed total wealth into asset ( $A_t$ ) and human capital ( $H_t$ ) wealth such that  $W_t = A_t + H_t$ , with log aggregate wealth approximated as  $w_t \approx \omega a_t + (1 - \omega)h_t$ , where  $\omega$  represents the share of asset wealth in total wealth ( $A_t/W_t$ ). The return to aggregate wealth can be decomposed into the return on its two components

$$1 + R_{wt} = \omega(1 + R_{at}) + (1 - \omega)(1 + R_{ht}), \quad (6)$$

and this can be rewritten into an equation for log returns as follows

$$r_{wt} \approx \omega r_{at} + (1 - \omega)r_{ht}. \quad (7)$$

Substituting the log aggregate wealth decomposition into the left-hand side of equation 5 and equation 7 into the right-hand side yields the following specification

$$c_t - \omega a_t - (1 - \omega)h_t = E \sum_{i=1}^{\infty} p_w^i (\omega r_{at} + (1 - \omega)r_{ht} - \Delta c_{t+1}). \quad (8)$$

Drawing on Jagannathan and Wang's (1996) assertion that human capital is marketable, Lettau and Ludvigson (2001) assumed that human capital is a function of labour income such that  $h_t = \kappa + y_t + z_t$ , where  $y_t$  is the log of labour income and  $z_t$  is assumed to be a zero mean stochastic stationary variable.<sup>iv</sup> Substituting this into equation 8 (ignoring the constant) and rearranging

$$c_t - \omega a_t - (1 - \omega)(y_t + z_t) = E \sum_{i=1}^{\infty} p_w^i (\omega r_{at} + (1 - \omega)r_{ht} - \Delta c_{t+1})$$

$$c_t - \omega a_t - (1 - \omega)y_t = E \sum_{i=1}^{\infty} p_w^i (\omega r_{at} + (1 - \omega)r_{ht} - \Delta c_{t+1}) + (1 - \omega)z_t. \quad (9)$$

where  $c_t - \omega a_t - (1 - \omega)y_t$  is the consumption aggregate wealth ratio (CAY) (Lettau and Ludvigson, 2001). Similarly to Campbell and Mankiw's (1989) formulation for the consumption wealth ratio in equation 5, the fact that the variables on the right-hand side of equation 9 are stationary implies that the three non-stationary variables on the left-hand side must be cointegrated. This means that they share a common stochastic trend, with the coefficients  $\omega$  and  $1 - \omega$  the parameters of this shared trend. Thus, these three variables may deviate from one another in the short-run when expectations of future returns change, but they have a long-run relationship captured in the cointegrating vector. The deviation of the variables from this long-run relationship is captured by CAY. The parameters of the cointegrating vector,  $\omega$  and  $1 - \omega$ , should sum to one, but this is unlikely to hold in testing this relation because proxies are used for the variables. In particular, this is likely to arise due to the use of consumption on non-durable goods and services rather than total consumption, given the difficulty associated with measuring the flow from durable goods (Lettau and Ludvigson, 2010).

As with equation 5, equation 9 implies that CAY must forecast growth in labour income, consumption growth and/or asset wealth. Moreover, CAY will forecast only those components of these variables that have significant transitory components given

the cointegrating framework in which CAY is derived. Given that share returns comprise a major component of returns to total asset wealth (Lettau and Ludvigson, 2005), the returns to aggregate equity are used as an approximation of the returns to asset wealth in the model (Lettau and Ludvigson, 2010). Accordingly, equation 9 indicates that CAY may be able to predict share returns. This forecasting power should be more pronounced provided consumption growth and returns to human capital in the following period are not too volatile, which appears to be the case in practice (Lettau and Ludvigson, 2001; Brennan and Xia, 2005).

### 3.2.2 Testing for Cointegration

The consumption, asset wealth and labour income series were tested for the presence of a unit root using the Augmented Dickey-Fuller (ADF) test, with the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) employed for confirmatory purposes. For both tests, an intercept and trend were included, where appropriate, and the optimal number of lags for the ADF test was determined using the Akaike information criterion.

For the purposes of estimating the cointegrating relationship, the single equation method proposed by Stock and Watson (1993) was used. This method involves dynamic least squares, where leads and lags of the differenced independent variable are added as additional explanatory variables in the long-run relationship estimated using ordinary least squares. This is shown as follows

$$c_t = \alpha + \beta_a a_t + \beta_y y_t + \sum_{i=-k}^k b_{a,i} \Delta a_{t-i} + \sum_{i=-k}^k b_{y,i} \Delta y_{t-i} + u_t, \quad (10)$$

where  $k$  refers to the number of lead/ lag terms of the explanatory variables (Lettau and Ludvigson, 2001:822; Camacho-Guiterrez, 2010:8), with  $k$  chosen so as to minimise the Akaike information criterion. The addition of leads and lags of the asset wealth and labour income as explanatory variables eliminates the effects of regressor endogeneity yielding super-consistent estimates of the cointegrating relationships. Moreover, with dynamic least squares, asymptotically valid standard errors can be computed using the Newey-West approach which adjusts for heteroscedasticity and autocorrelation (Gao and Huang, 2008). Following Lettau and Ludvigson (2001), equation 10 was estimated with an intercept but without a trend term. Although there has been some debate as to the validity of imposing the restriction of no deterministic trend in the cointegrating

relationship for CAY (see Hahn and Lee, 2006), because the true data generating process can never be known, the validity of such an assumption is always open to debate. The Phillips and Ouliaris (1990) test of the stationarity of the residuals of equation 8 was then conducted to determine if the variables were cointegrated.

Lettau and Ludvigson (2010) acknowledge that if the two components of wealth in equation 4 – asset and human capital wealth – were themselves cointegrated and if labour income captured the trend in the latter, then it is plausible that a second cointegrating relationship may exist in the sample. If a second cointegrating relationship exists but only a single-cointegrating relationship has been estimated (as is the case with dynamic least squares), then the estimates of the coefficients of the cointegrating vector will be incorrect as they will reflect a linear combination of the two relationships. Although very little evidence of the existence of a second relationship has been documented (Ioannidis *et al.*, 2006), to ensure that the results of this test were not sensitive to this, the systems-based method of Johansen (1988) was also used to test for the presence (and number) of cointegrating relationships. Johansen’s (1988) trace and maximum-eigenvalue tests were conducted for this purpose.

### 3.3 Assessing the Ability of CAY and other Ratios to Predict Share Returns

Only in-sample tests of the predictive ability of financial ratios were conducted in this study, as although these have been criticised (Goyal and Welch, 2007), Inoue and Kilian (2005) demonstrate that these tests actually have greater power asymptotically than out-of-sample tests. The excess real market were examined for predictability using in-sample tests. The regression used for this purpose takes the following form

$$r_{m,t+1} = \gamma'z_t + \varepsilon_{1,t+1}, \quad (11)$$

where  $r_{m,t+1}$  are the excess real returns on the market,  $z_t$  is a vector of lagged predictor variables and  $\gamma'$  represents a vector of coefficients (Lettau and Ludvigson 2010:633).

This regression was initially estimated separately for each variable, and then a multivariate regression combining the predictor variables was undertaken to assess their joint ability to predict share returns. The null hypothesis that the predictive variable had no explanatory power ( $\gamma = 0$ ) was examined against a two-sided alternative that the

variable was able to significantly predict future returns ( $\gamma \neq 0$ ). The explanatory power of the instruments was also assessed using  $\bar{R}^2$ . The use of CAY as an explanatory variable in equation 11 does not require an adjustment to the standard error computation, despite the fact that it is a generated regressor, because cointegrating parameters converge to their true values at a rate of T (Johansen, 1988).

In addition to the one-quarter ahead regressions, the forecasting power of the variables was also examined over longer horizons. This is important because the varying nature of share returns over different horizons may provide biased results if only one-horizon is examined (Richardson, 1993); single-period estimates may be subject to noise (Valkanov, 2003); and the long-run regressions also provide a means of illustrating the economic implications of forecasting (Cochrane, 2005:395). For this purpose, the cumulative returns over two, four, six, eight and twelve quarters ahead were examined in the following model

$$r_{m,t+H,H} = \gamma_H' z_t + \varepsilon_{1,t+H,H}, \quad (12)$$

where  $r_{m,t+H,H}$  is the H-quarter continuously compounded excess real return equal to  $r_{m,t+1} - r_{f,t+1} + \dots + r_{m,t+H} - r_{f,t+H}$  (Lettau and Ludvigson, 2010:635). Newey-West standard errors were used to resolve the serial correlation that arises because of the use of overlapping returns.

If the dependent and independent variables in these regressions are non-stationary it can give rise to inaccurate assessments of the predictive power of the variables in the tests. To assess whether the variables in this study satisfied the stationarity criterion, the ADF and KPSS tests were used. In the literature, it has been found that financial ratios such as dividend-to-price and earnings-to-price frequently contain a unit root or at the very least are highly persistent. Moreover, cumulative returns may also exhibit this property because of the use of overlapping data. While the use of non-stationary variables obviously gives rise to spurious regressions which cannot be reliably interpreted (Cochrane, 2005:395), even the use of explanatory variables which are highly persistent can lead to incorrect inferences because the effects of persistence accumulate over time yielding coefficients and  $\bar{R}^2$  values which rise monotonically with the horizon (see Cochrane, 2005:394-395). To account for this, the  $\bar{R}^2$  measure of Hodrick (1992) was

computed as this provides an implied measure of the explanatory power from a long-run regression.

## 4. Results and Analysis

### 4.1 Analysis of CAY

The results for the ADF and KPSS tests confirmed that consumption, asset wealth and income were non-stationary in levels but stationary in first differences<sup>v</sup> and therefore the cointegration tests were performed. As shown in Table 1, for the Phillips-Ouliaris test the null hypothesis of no cointegration was rejected in favour of the alternative that the three variables were cointegrated.<sup>vi</sup> The results from the Johansen (1988) cointegration tests were largely consistent with this conclusion, as there was evidence (at 10% level) of one cointegrating relationship, but no evidence of a second relationship. The finding of only a single cointegrating vector between the three variables is consistent with the observation of Lettau and Ludvigson (2010) that it is rare to find a second relationship between consumption, asset wealth and labour income, with Ioannidis *et al.*'s (2006) finding for the U.K of two vectors the outlier in this regard.

This conclusion thus indicates that the coefficients from the dynamic least squares regression are correct and can be interpreted as they are super-consistent and the standard errors are asymptotically valid. The equation for the cointegrating relationship is as follows (with  $t$ -statistics computed using the Newey-West standard errors shown in brackets)

$$c_t = 0.24 + 0.21a_t + 0.74y_t. \quad (13)$$

(0.56) (4.12) (8.41)

The coefficients indicate that positive relationships exist between consumption and the two variables, which is in line with expectations as an increase in labour income and asset wealth should result in higher consumption expenditure. As predicted, the coefficients sum to less than one, but the relative magnitudes of the coefficients reflects a stronger relationship between labour income and consumption than asset wealth and consumption suggesting that labour income drives consumption more than asset wealth. This is the same pattern identified by Lettau and Ludvigson (2001) and Hahn and Lee (2006) for the U.S and Gao and Huang (2008) for the U.K and Japan respectively. In

contrast, Ioannidis *et al.* (2006) and Li *et al.* (2011) found the relationship between asset wealth and consumption to be stronger.

*Table 1: Cointegration Tests*

	Test Statistics	
Phillips-Ouliaris Test	$\tau$ -statistic: -5.46**	
Johansen Test:	Trace statistic:	Maximum Eigenvalue statistic:
$r = 0$	24.74*	19.65*
$r = 1$	7.10	6.93

\*, \*\* indicate significance at 10% and 5% respectively for the  $\tau$  statistic based on the MacKinnon (1996) critical values and for the trace and maximum-eigenvalue tests based on the MacKinnon *et al.* (1999) critical values.

Deviations from the shared trend between consumption, asset wealth and labour income will occur in the short-run, as captured by CAY. To ascertain whether these deviations represent transitory movements in consumption, asset wealth and/or income, a cointegrated vector autoregression was estimated, with the coefficients on CAY representing the adjustment parameters (or error correction mechanisms) showing how each of the three variables adjusts to restore equilibrium in the long-run relationship. These coefficients, shown in Table 2, indicate that the error correction mechanism was significant in the asset wealth equation (at 5%) and in the consumption equation (at 10%). This indicates that short-term deviations in the long-run relationship can be viewed as transitory movements principally in asset wealth and partially in consumption but not labour income. Moreover, the observation of a positive coefficient for the adjustment term in the equation for asset wealth is consistent with the theoretical relationship that an increase in CAY should lead to an increase in asset wealth. Assuming asset wealth and share returns are highly positively correlated, this results suggests that CAY may have power to explain future returns; the extent to which this is true is examined in the following section.

*Table 2: Estimates of the Error Correction Estimates*

	$\Delta \ln c_t$	$\Delta \ln a_t$	$\Delta \ln y_t$
$cay_{t-1}$	-0.1307* (-1.7264)	0.6795** (2.8306)	0.0293 (0.3043)

$\Delta \ln c_t$ ,  $\Delta \ln a_t$  and  $\Delta \ln y_t$  reflect the first difference in consumption, asset wealth and labour income respectively. The coefficients from the error correction vector autoregression are shown, with  $t$ -statistics provided in round parentheses thereunder. \* and \*\* indicate significance at 10% and 5% respectively.

## 4.2 Predictive Regressions

The summary statistics for the excess market returns over the various horizons and the forecasting variables are shown in Tables 3 and 4 respectively. As is evident, the autocorrelation in the market returns increased as the time horizon increased, which is partly due to the use of overlapping returns. However, despite the persistent nature of the returns, both the ADF and KPSS tests confirm that these cumulative returns were stationary. The five other predictor variables, including CAY, exhibited substantial persistence over time, but they satisfied the condition of stationarity. CAY had very low correlations with the contemporaneous values of the excess real market return, term spread and relative Treasury bill yield; however, it had a high negative correlation with both the dividend-to-price and earnings-to-price ratios, with these two financial metrics themselves highly correlated. These strong relationships suggest that CAY may track analogous predictable components of the share returns captured by the financial ratios.

*Table 3: Summary Statistics for the Excess Real Market Returns*

	Horizon (H) in quarters					
	1	2	4	6	8	12
Mean (%)	0.85	1.87	3.78	5.46	7.12	11.16
Std Dev (%)	9.74	13.92	19.51	23.78	27.04	32.49
Autocorrelation	0.02	0.49	0.76	0.83	0.87	0.90
ADF Test	-9.30**	-3.57**	-2.82**	-2.45**	-2.12**	-1.88*
KPSS Test	0.04	0.14	0.12	0.06	0.07	0.09

\* and \*\* indicate significance at 10% and 5% respectively for the ADF and KPSS tests.

*Table 4: Summary Statistics for the Predictor Variables*

	$r_{mt}$	Relative Treasury bill yield	Term spread	D/P	E/P	CAY
Panel A: Univariate Statistics, Unit Root and Stationarity Tests						
Mean (%)	0.85	-0.53	0.98	-3.62	-2.67	2.22
Std Dev (%)	9.74	2.09	1.69	0.21	0.20	0.03
Autocorrelation	0.02	0.83	0.85	0.84	0.84	0.78
ADF Test	-9.30**	-4.98**	-3.37**	-3.99**	-3.10**	
KPSS Test	0.04	0.05	0.11	0.09	0.07	



Panel B: Correlations						
$r_{mt}$	1					
Relative Treasury bill yield	-0.34	1				
Term spread	0.30	-0.50	1			
D/P	-0.35	-0.17	-0.31	1		
E/P	-0.39	-0.11	-0.41	0.94	1	
CAY	0.16	-0.30	0.10	0.54	0.57	1

$r_{mt}$ , E/P and D/P refer to the excess real market return, earnings-price and dividend-price ratio respectively. \* and \*\* indicate significance at 10% and 5% respectively for the ADF and KPSS test using the MacKinnon (1996) and Kwaitowski *et al.* (1992) critical values respectively. The ADF and KPSS tests are not shown for CAY because the cointegration tests have already proven the stationarity of this series.

The results from the predictive regressions are shown in Table 5. The lagged market return had no ability to forecast future returns irrespective of the time-horizon. This result is consistent with the low autocorrelation in the series and indicates that there was no evidence of mean reversion over time. The U.S evidence is mixed with regards to the predictive power of the lagged market return, as although the early work of Fama and French (1988b) and Lettau and Ludvigson (2001a) found significant univariate forecasting power, the more recent findings of Lettau and Ludvigson (2010) contradict this. The relative Treasury bill yield was identified to have predictive power on the JSE for one-quarter ahead returns and then for horizons longer than six quarters. In contrast, the term spread only had significant (at 10%) predictive power for one-quarter ahead returns; however, the finding that this variable was more closely related to short-term rather than long-term business cycles is similar to Fama and French's (1989) results. The signs for both variables were consistent with the view that spreads and short-term interest rates were positively and negatively correlated respectively with future business conditions.

As documented previously, Gupta and Modise (2012b, 2013) found that the term spread and relative Treasury bill had predictive power for returns in South Africa and thus the findings from this analysis are consistent with their results. Moreover, Gupta and Modise (2012b, 2013) also noted that the term spread's forecasting ability was limited to short-run horizons, while the relative Treasury bill was able to predict returns at both short- and long-horizons (although in this study it was less successful at two and four quarters ahead). The term spread and relative Treasury bill yield could explain approximately 4% of the variation in returns in one-quarter ahead, as measured by  $\bar{R}^2$ ,

and Hodrick's (1992)  $R^2$  confirmed that the explanatory power of these variables was not inflated by any persistence in these forecasting variables. Although this explanatory power is low, it is comparable to international studies such as Lettau and Ludvigson (2010), who found that the relative Treasury bill yield, for example, could explain 6% of the one-quarter ahead variation in returns, with this declining as the forecast horizon increased (based on Hodrick's (1992)  $R^2$ ).

The dividend-to-price and earnings-to-price ratios were found to exhibit no forecasting power over a one-quarter horizon; however, over longer horizons both financial ratios were seen to be significant predictors of returns, with positive coefficients consistent with the view that these ratios move with future business cycles. The  $\bar{R}^2$  values confirmed that for periods longer than four quarters, these two variables could explain a substantial component of the variation in the future risk premium. However, Hodrick's (1992)  $R^2$  values provide contradictory evidence, as they indicate that neither ratio could capture substantial variation in returns. These results thus reveal that the significance of the coefficients of the predictive regressions and high  $\bar{R}^2$  estimates using the dividend-to-price and earnings-to-price measures may be a statistical artefact arising from the persistence of these ratios. The finding of limited forecasting power after accounting for the near roots in this series, mirrors the results of Gupta and Modise

*Table 5: Forecasts of Multiple Quarter Excess Real Market Returns*

Regressors	Forecast horizon ( $H$ ) in quarters					
	1	2	4	6	8	12
$r_{mt}$	0.02 (0.22) [-0.01] {0.00}	-0.06 (-0.04) [-0.01] {0.00}	0.51 (0.30) [-0.01] {0.00}	-1.09 (-0.51) [-0.01] {0.00}	-0.63 (-0.25) [-0.01] {0.00}	-1.72 (-0.60) [-0.00] {0.00}
Relative Treasury bill yield	-2.06 (-2.12)** [0.04] {0.04}	-2.62 (-1.46) [0.03] {0.00}	-4.14 (-1.50) [0.04] {0.04}	-7.01 (-1.92)* [0.08] {0.02}	-10.49 (-2.47)** [0.15] {0.04}	-9.79 (-1.72)* [0.09] {0.01}
Term spread	1.98 (1.86)* [0.04] {0.04}	2.26 (1.33) [0.02] {0.04}	3.18 (1.10) [0.02] {0.00}	4.68 (1.31) [0.03] {0.00}	4.09 (1.07) [0.01] {0.00}	-3.37 (-0.71) [-0.00] {0.01}
D/P	1.41 (1.55) [0.01]	2.92 (1.94)* [0.03]	5.59 (2.47)** [0.07]	8.42 (2.56)** [0.12]	11.21 (2.77)** [0.20]	20.86 (4.87)** [0.44]

	{0.05}	{0.05}	{0.07}	{0.05}	{0.06}	{0.03}
E/P	0.99 (1.13) [-0.00] {0.01}	2.07 (1.41) [0.01] {0.06}	3.60 (1.74)* [0.02] {0.07}	5.60 (1.85)* [0.05] {0.07}	8.34 (2.09)** [0.09] {0.03}	18.73 (4.37)** [0.35] {0.02}
CAY	2.88 (3.38)*** [0.08] {0.08}	3.98 (2.98)*** [0.07] {0.09}	6.90 (3.36)*** [0.11] {0.07}	8.13 (2.21)** [0.10] {0.04}	9.89 (2.64)*** [0.12] {0.03}	15.27 (4.08)*** [0.22] {0.05}
Relative Treasury bill yield	0.19 (0.15) 2.58	0.54 (0.12) 3.59	1.28 (0.44) 5.96	-0.17 (-0.06) 7.97	-5.24 (-1.57) 5.09	-5.68 (-1.39) -0.53
Term spread	(2.05)** 0.73 (0.59)	(1.69)* 2.41 (0.85)	(1.79)* 4.86 (1.38)	(2.21)** 8.63 (2.04)**	(1.67)* 10.35 (2.19)**	(-0.18) 17.88 (3.29)***
D/P	2.80 (2.36)** [0.11] {0.06}	3.19 (1.78)* [0.10] {0.07}	5.30 (2.09)** [0.17] {0.09}	4.26 (1.15) [0.23] {0.08}	2.89 (0.74) [0.29] {0.06}	3.10 (0.61) [0.46] {0.07}
CAY						

This table shows the coefficients from the predictive regressions. Beneath each coefficient in round parentheses is the  $t$ -statistic computed using the Newey-West (1987) adjusted standard errors. The regression  $R^2$ , adjusted for degrees of freedom,  $\bar{R}^2$ , is shown in square parentheses, with Hodrick's (1992)  $R^2$  presented thereunder in curly parentheses. \*, \*\* and \*\*\* indicate significance at 10% and 5% respectively for the  $t$ -tests.

(2012a) based on their bootstrapping procedure. Moreover, this is also broadly consistent with the findings in the U.S after similar adjustments for the dividend-to-price and earnings-to-price ratios.

The forecasting results for CAY are shown in row 6 of Table 5. A significant coefficient was obtained for the one-quarter ahead horizon, consistent with the conclusions drawn from the error correction mechanism that CAY can forecast future returns. The coefficient was positive in accordance with the theoretical relationship that if market returns are forecast to increase in the future, then investors who desire smooth consumption levels will allow consumption to temporarily increase above its long-term relationship with asset wealth and labour income on the basis that future consumption will be supported from higher future returns. The opposite is true if returns are expected to decrease, with investors reducing consumption below the long-term level with asset wealth and labour income so as to protect future consumption levels against lower returns (Lettau and Ludvigson, 2001a). The explanatory power for the one-quarter

ahead returns was 8%, as measured by  $\bar{R}^2$ , which is comparable to the 9% and 8% documented by Lettau and Ludvigson (2001a; 2010) in their studies of the U.S. Gao and Huang (2008) obtained a lower  $\bar{R}^2$  for the U.K of 4%, and a 0%  $\bar{R}^2$  for Japan where CAY had no explanatory power. In the earlier regressions, the term spread was the most successful variable for predicting one-period ahead returns and was found to be able to explain 4% of the variation in one-quarter ahead returns. Thus, it is clearly evident that CAY by itself is a superior predictor of one-quarter ahead returns on the JSE than any of the traditional variables.

As the results in Table 5 confirm, the success of CAY in forecasting share returns was not only limited to the short-run, as it was able to explain 12% and 22% of the variation at eight- and twelve-quarters ahead, although this is not as substantial as the predictive power documented by Lettau and Ludvigson (2010) for the U.S of 28% and 34% for the same horizons. However, after accounting for the persistent nature of the measure, the explanatory power was notably reduced over longer horizons, as captured by Hodrick's (1992)  $\bar{R}^2$ . This finding does differ from that documented by both Rasmussen (2006) and Lettau and Ludvigson (2010) who found that CAY retained its forecasting power over long-horizons on the U.S market after accounting for the persistence in the series.

The joint predictive power of the forecasting variables in this sample was also assessed, with the results thereof shown in the final row of Table 5. The earnings-to-price and dividend-to-price ratios were not examined jointly because of their high correlation, but the high correlation between the dividend-to-price ratio and CAY was not found to be problematic. Only the regression with the dividend-to-price ratio is shown, in the interests of brevity, as it was found to perform better than earnings-to-price. The lagged market return was excluded as the combination without this variable yielded higher explanatory power. As can be seen CAY retained its significance but only for one-, two and four quarters ahead. Lettau and Ludvigson (2010) found CAY to still be a significant determinant of future period returns when combined with the dividend-to-price ratio; however, the results from this study suggest that while this was true for short horizons, at longer horizons of over a year, CAY became insignificant in the joint regressions as the effects of the dividend-to-price ratio crowded out CAY. The term spread was significant at the one-quarter ahead horizon when analysed individually, but

when combined with the other variables it was also significant at two, four and eight-quarters ahead. Interestingly, however, when combined with the other variables, the relative Treasury bill yield had no forecasting ability. This certainly also confirms some co-movement between the forecasting variables based on the interest rate and CAY. Accordingly these results confirm that CAY does contain important information about future period returns that is not contained in the traditional forecasting variables but over longer horizons much of this information appears to also be contained in the dividend-to-price ratio with the latter dominating potentially because of its near unit-root properties.

## **5. Conclusion**

The evidence from the studies of Gupta and Modise (2012a, b and 2013) provide little support in favour of the assertion that share returns on the JSE are predictable, although they did identify that the short-term interest rate had some forecasting ability. As a follow-up to these studies, this research sought to determine whether the consumption aggregate wealth ratio (CAY) could be used to predict share returns in the South African market. The results of this analysis generally confirmed the findings of Gupta and Modise (2012b, 2013) that the term spread and relative Treasury bill yield have some power to predict returns, over the short-run and long-run respectively. Any forecasting ability of the dividend-to-price and earnings-to-price ratios appeared to largely be a statistical artefact, especially at long horizons. In contrast, CAY was found to be a significant predictor of returns at short horizons of less than a year, although its power to forecast returns at longer horizons was limited.

These tests thus reveal that Lettau and Ludvigson's (2001) CAY, which captures the deviations from the long-run relationship between consumption, asset wealth and labour income, can be used to predict share returns on both developed and emerging markets. Thus, although participation levels in the market may be low, there are sufficient investors in the market adjusting their holdings and consumption levels in response to expectations of future market returns to give rise to CAY's significant predictive power for following period returns.

There is substantial evidence that share prices not only act as a leading indicator of output and inflation but that there are also spillover effects from the share market to the real sector. This study has shown that policy makers can use CAY to predict future business returns so as to be able to better implement policies to minimise the impacts of market downturns. Furthermore, the fact that share returns can be predicted using CAY, which is based on publicly available information, means that investors can structure asset allocation decisions so as to earn higher risk-adjusted returns.

## **Policy Brief**

The financial and real sectors of the economy are inextricably linked; with strong evidence to suggest that there are spillover effects from the stock market to the real economy, with the stock market usually leading the real sector. As such, substantial stock market downturns can negatively impact output. This is of particular concern in an emerging country such as South Africa where maintaining and supporting growth in gross domestic product is critical for the continued development of the country and the allied lowering of poverty levels. Accordingly, for policy makers the ability to forecast market downturns accurately enables them to implement appropriate policies to limit the effects of these stock market downturns on the real economy. Current measures used to forecast share returns such as the Treasury bill yield and dividend-price ratio have limited success in this regard.

The study of Charteris and Strydom (2016) presents a novel approach to forecasting stock market returns on the South African market that links the real and financial sectors directly through consumption, labour income and share returns. They find that the consumption aggregate wealth ratio can predict share returns one quarter ahead and provides a better measure than any of the traditional measures that policy makers currently utilise. Accordingly, these authors recommend that financial economists make use of the consumption aggregate wealth ratio in identifying changing patterns in risk as this is likely to provide a more accurate reflection of future returns. In so doing, this will provide policy makers with the opportunity to implement appropriate monetary or fiscal policy to prevent falloffs in production such as those experienced post the financial crisis of 2007/ 2008.

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<sup>i</sup> Rather than assuming the series grew equally during each quarter of the year, the more accurate technique for interpolation that is commonly employed in economics involving a spline was used (Kushnirsky, 2009). A spline is a polynomial between each pair of observed data points, where the coefficients are determined so as to ensure a smooth fitting function up to some order of derivative. A cubic spline fits a continuous curve with a piecewise series of cubic polynomial curves which are continuous up to the second derivative (Kurshnirsky, 2009).

<sup>ii</sup> For the period 1990:03 to 1994:04, labour income was only adjusted for taxes and not transfer payments, as this information was not explicitly recorded by the SARB prior to 1995.

<sup>iii</sup> The constant in this equation is excluded from the derivation as it simplifies the analysis.

<sup>iv</sup> This relation is drawn from the work of Campbell and Shiller (1989).

<sup>v</sup> These results are available from the author.

<sup>vi</sup> These results were not found to be sensitive to the choice of lead and lag parameters included in the dynamic least squares regression or the number of lags used in the computation of the test statistic. These results are available from the author.