Macroeconomic variables and the cross-section of Johannesburg Stock Exchange returns

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This study adopts the Chen, Roll & Ross prespecified variable approach to priced arbitrage pricing theory factor (APT) identification on the Johannesburg Stock Exchange (JSE). It is observed that the dichotomy in the return generating processes underlying South African mining and industrial shares leads to cross-sectional correlations in the residual errors of linear factor models that do not employ factor analytically extracted explanatory variables. As a result, a 'two residual market factor' approach is introduced in this study. Employing the iterated non-linear seemingly unrelated regression technique of McElroy & Burmeister (1988), it is found that the rand gold price, the rate on long bonds, the Dow-Jones Industrial Index and the level of gold and foreign exchange reserves together with the Industrial and All-Gold residual market factors represent priced sources of risk within the framework of the APT over the period 1985 to 1995. The pricing relationships estimated are found to be inconsistent with those implied by the capital asset pricing model. These results are robust across the 'unconstrained intercept' and 'zero beta' cross-sectional model specifications. The findings of the study, however, imply that the influence of macroeconomic variables on the JSE is most parsimoniously expressed in the two factor APT model of Van Rensburg & Slaney (1997).

Introduction and prior research

The majority of studies investigating the relationship between macroeconomic forces and the cross-section of equity returns have been conducted within the framework of the Ross (1976; 1977) arbitrage pricing theory (APT). As is the case in American literature, South African research concerning the number and nature of the priced APT factors can be broadly subdivided into (i) those studies adopting a factor analytic approach (see Roll & Ross, 1980; Reinganum, 1981; Brown & Weinstein, 1983; Chen, 1983; Cho, 1984; Cho, Elton & Gruber, 1984; Pari & Chen, 1984; Dhrymes, Friend & Gultekin, 1984; Conway & Reinganum, 1988; Lehmann & Modest, 1988; Connor & Korajczyk, 1988; Brown, 1989; Kryzanowski, Lalancette & To, 1994 *inter alia*) and (ii) those studies that preselect certain macroeconomic variables and proceed to test whether sensitivities to these variables are associated with statistically significant risk premia (see Chen, Roll & Ross, 1986; Chen, Chan & Hseih, 1985; Burmeister & McElroy, 1988; and McElroy & Burmeister, 1988). This study adopts the latter approach to APT factor identification on the JSE.

Although links between the above two branches of research are seldom made, factor analytic evidence is germane to the appropriate specification of a linear factor model that employs only prespecified macroeconomic variables as explanatory forces. In particular, factor analytic findings are relevant with regard to the possible violation of the 'diagonality' assumption of such a model. Prior South African factor analytic studies support the extraction of at least two but no more than three principal components as a parsimonious representation of the return generating process on the JSE (Page, 1986, 1989; Biger & Page, 1993; and Van Rensburg & Slaney, 1997). With regard to the economic interpretation of these statistical constructs, Page (1986) found that mining shares tended to load on the first principal component extracted while industrial firms tended to load on the second:

> 'What the findings do suggest is that the underlying macroeconomic variables determining the return generating process can be divided into those that influence the mining sector to a greater extent and those that influence the industrial sector to a greater extent' (1996: 42).

Van Rensburg & Slaney (1997) included observable marker variables in their promax rotation of the factor pattern and found that the JSE Actuaries All-Gold and Industrial Indices could be employed as observable proxies for the first two factor analytic factors on the JSE. It is argued that using these proxies (i) alleviates the primary difficulty of factor analytic procedures, that is the economic interpretability of estimated factor loadings and risk premia (Fama, 1991: 594–595); and (ii) allows access to the powerful system equation methodologies adopted by McElroy & Burmeister (1988) when conducting cross-sectional testing. These observable proxies are of relevance to the 'two residual market factor' approach introduced in later on in this article.

Prior research adopting the prespecified variable approach to APT factor identification on the JSE is limited to Van Rensburg (1996) who found that unexpected movements in the Dow-Jones Industrial Index, short-term interest rates, the term structure of interest rates and the 'residual market factor'[2] of Burmeister & Wall (1986) were associated with statistically significant risk premia over the decade of the 1980s. These findings do not preclude the possibility of other macrovariables also being priced over this period. For example, default premia (the difference in yields between corporate and government bonds) were consistently found to be an important explanatory variable in the US environment (Ross, S. – personal correspondence; see also Chen, Roll & Ross, 1986; and McElroy & Burmeister, 1988 *inter alia*). Due to the perceived poor quality of the data in the thinly traded South African corporate bond market, this variable was not selected for testing. In addition, the caveat is applied that, due to the fact that the term structure and short-term interest rate variables displayed a marked degree of multicollinearity, in the multiple regression context the former variable may actually be more representative of a 'long bond' rather than a term structure effect.

Most importantly, the prespecified variable approach adopted in Van Rensburg (1996) does not take account of the dichotomy in the return generating processes underlying mining and industrial shares on the JSE. Consequently, a factor analytic augmentation to the prespecified approach is suggested in Van Rensburg (1997) where the residual errors of linear factor models, employing the preidentified priced factors, are factor analysed. The factor scores derived in this manner are then employed alongside the prespecified macro-variables in order to provide a more comprehensive specification of the return generating process than would otherwise be the case. It was found that (i) the factors more than doubled the explanatory power of models employing only macroeconomic variables to explain both 'market' and individual asset returns; (ii) all but three shares loaded significantly on one of the residual factors; and (iii) one of the factor analytically extracted factors was also able to significantly explain the cross-section of share returns thereby justifying its inclusion in the LFM underlying the APT. As expected, following varimax and promax factor rotation, it was found that mining shares tended to load on the first 'residual factor' while industrials loaded on the second.

In a recent study, Van Rensburg (1999) reinvestigated the issue of the macroeconomic identity of candidate APT factors on the JSE. A comprehensive list of macroeconomic series were investigated in the context of the socioeconomic and institutional factors that shaped the South African economy over the turbulent period of democratic transition, 1965 to 1995. Based on a time-series analysis of the data, the following variables were selected as candidate APT factors for future cross-sectional testing: (i) the rand gold price, (ii) the three month Banker's Acceptance rate, (iii) the 10-year long bond rate, (iv) the Dow-Jones Industrial Index, (v) the balance on the current account, and (iv) the money market shortage.

This study extends prior South African research in the following ways. First, an independent and more recent share sample than that employed by Van Rensburg (1996, 1997) is analysed in this study. The search for candidate factors is also based on the more extensive time-series investigation conducted in Van Rensburg (1999) (the second section). Second, employing the vector autoregressive (VAR) methodology, unexpected movements in the factors are extracted accounting for the predictive power of other series (in addition to past factor realisations) in the formation of investor expectations (first part, third section). Third, to avoid the problem of contemporaneous correlation of residual errors in the models, *two* residual market factors are employed throughout the subsequent analysis. These ('All-Gold' and 'Industrial') residual market factors are employed as observable and economically interpretable proxies for the factor analytic augmentation proposed in Van Rensburg (1997). The econometric nature of the bias resulting from the omission of this procedure is elaborated on and illustrated through an example in the second part of the third section. Fourth, three cross-sectional model specifications are estimated using the systems equations technique of McElroy & Burmeister (1988): the 'exact APT restrictions' model, the 'unconstrained intercept' model and the 'zero beta' version of the APT. The results are compared across model specifications to draw inferences on the validity of the APT pricing restrictions and the robustness of the findings with respect to the proxy employed for the default free asset.

Sample selection

In comparison to investigators of the major US equity markets, thin-trading is a far more serious concern for South African researchers.[3] The implications for South African researchers are considered by Bradfield (1989) and Bowie & Bradfield (1993) in the context of the estimation of systematic risk on the JSE:

> 'The main cause of the bias associated with estimation problems in environments characterised by thin-trading is the fact that *recorded prices* are used to represent true *underlying prices*. For example, when a security has not been traded in a period in question then the recorded price of the security remains unchanged, and represents the outcome of some transaction in a previous period. The underlying (theoretical) price of the security, by contrast would reflect the arrival of any new information in the period in question' (Bradfield, 1989: 23).

Finance researchers must either correct for the presence of thin-trading, through the trade-to-trade approach (as suggested by Bowie & Bradfield, 1993: 19) or omit thinly traded securities from their sample, incurring a self-selection bias. In this study thinly-traded securities are omitted from the sample. A prominent environmental feature germane to interpreting the sample employed in this study, is the dominant and focused nature of South African institutional investment policy:

> 'The ... dominant position of financial institutions in trading activities on the JSE has been reflected in the buying policies of these institutions. The latter concentrate their buying activities on a narrow range of equities amounting to about 50 shares which are of a blue chip status, and reasonably marketable. The great majority of shares quoted on the JSE do not attract the attention of these institutions' (*Economic Focus*, 1990, 84[6]; see also *The JSE Centenary Publication*, 1987: 134).

The inferences drawn from this research are appropriately interpreted to apply to this coterie of JSE shares that constitute the bulk of investor trade and interest.

All of the shares comprising the JSE Actuaries All-Share Index on 3 January 1995 were initially considered for inclusion in the sample. Any shares listed after 31 January 1985 were excluded. Also, any share that had zero volume traded for more than 20 weeks out of the 520 weeks available in the sample period was excluded. In this manner the econometric problems associated with thin trading were avoided. This reduced the sample, from the 141 shares comprising the index as at 3 January 1995, to 55 that met both of these criteria. Gold shares were found to be, on average, more actively traded than industrials. This led to a higher representation of mining shares than industrials in terms of both market capitalization and the proportion of total shares listed in each sector. As a result, all of the shares listed in the non-mining sectors at 19 April 1995 were examined and all shares that met the above criteria were included in the sample. This procedure resulted in a final sample of 84 shares. Table 1 presents a overview of the sectoral composition of the sample.

Table 1 Sectoral overview of the sample

Sector	Percentage of sector market capitalization represented by sample	No. of shares in the sample as a percentage of total shares in the sector
Mining producers		
Coal	39.8%	11.1%
Diamonds	75.3%	11.1%
Rand and others	69.6%	25.0%
Evander	93.2%	50.0%
Klerksdorp	71.1%	44.4%
West Witwatersrand	98.3%	60.0%
OFS	12.8%	37.5%
Curtailed operations	0.0%	0.0%
Copper	83.0%	20.0%
Manganese	89.6%	33.3%
Platinum	75.3%	20.0%
Other metals & minerals	8.2%	15.4%
Mining financial		
Mining houses	72.9%	25.0%
Mining holding	94.3%	26.7%
Mining exploration	4.5%	6.7%
Financial		
Banks & fin. Services	64.3%	7.3%
Insurance	57.3%	12.5%
Investment trusts	43.6%	4.8%
Property	55.8%	14.3%
Property trusts	23.9%	25.0%
Property loan stock	0.0%	0.0%
Industrial		
Industrial holdings	37.9%	13.6%
Bev., hotels & leisure	48.9%	6.3%
Build., constr. & allied	36.6%	4.0%
Chem., oils & plastics	92.2%	30.8%
Clothing	0.0%	0.0%
Electronics	8.6%	2.3%
Engineering	34.7%	6.5%
Food	52.4%	15.6%
Furn., hous., & allied	0.0%	0.0%
Motor	0.0%	0.0%
Paper & packaging	73.1%	12.5%
Pharm. & medical	0.0%	0.0%
Print. & publishing	0.0%	0.0%
Steel and allied	26.8%	16.7%
Stores	20.3%	10.3%
Transport	0.0%	0.0%

Arithmetic returns were calculated as follows:

$$R_{it} = \frac{D_{it} + (P_{it} - P_{i(t-1)})}{P_{i(t-1)}} \qquad (1)$$

where: R_{it} = return on share i in period t

D_{it} = the dividend on share i for which the last date to register fell within period t

P_{it} = the price of share i at the end of period t.

Note that dividends are recognised in their 'ex dividend' rather than payment months (see Van Rensburg, Slaney & Hardy, 1997).

Following the prior research of Van Rensburg (1999), the first differences in levels or natural logarithms (prefixed '*D*' and '*DL*' respectively) of the following macroeconomic variables were initially selected as candidates for cross-sectional testing:

A: Share indices:

1. The JSE All-Share Index (DLALSI)
2. The Industrial Index (DLINDI)
3. The All-Gold Index (DLGOLDI)

B: Sector earnings

1. All-Share Indexed Earnings (DLAE)
2. Industrial Indexed Earnings (DLIE)
3. All-Gold Indexed Earnings (DLGE)

C: Economic forces

1. The Dow-Jones Industrial Index (DLDJ)
2. The rand gold price (DLGOLR)
3. The R150 10 year government bond (DR150)
4. The three month Bankers Acceptance rate (DRBAS)
5. The level of gold and foreign exchange reserves (DLGFX)
6. The money market shortage (DLMMS)

Note that sectoral earnings were included due to the strong theoretical prior that unexpected movements in these series should influence equity returns. The balance on the current account was excluded from the analysis as only quarterly data was available for this variable.

Month-end data was downloaded from the INET database at the University of Natal, Durban, over the period January 1980 to December 1994 while the cross-sectional analysis was conducted over the sub-period 1985.01 to 1994.12. In this way it was ensured that the inclusion of lagged variables in the modelling of unexpected movements in the factors did not consume any degrees of freedom in the ensuing cross-sectional tests. Indexed sector earnings were derived for each of the three JSE price indices from their earnings yields for example AE = $[E/P]_{ALSI}$ *ALSI where AE = the 'indexed accounting earnings' of the JSE All-Share Index, $[E/P]_{ALSI}$ = the earnings yield of the All-Share Index and ALSI = the level of the All-Share Index. All of the underlying series except for the level of gold and foreign exchange reserves[4] are publicly observable on a daily basis and, thus, their inclusion in the forecasting model, which uses monthly data, will not result in any meaningful Banz & Breen (1986) 'look-ahead' bias.

Results

Estimation of candidate factors

The linear factor model underlying the Ross (1976) Arbitrage Pricing Theory assumes that all risk factors have a mathematical expectation of zero i.e $E(f_{kt}) = 0$ for all factors $k = 1,\dots, K$) and are void of autocorrelation ($E(f_{kt}f_{kt+s}) = 0$ where $s \neq 0$) due to their theoretical depictions of innovations in certain pervasive risk factors.[5] In this study a vector autoregressive (VAR) model is employed to forecast future values of the candidate macrovariables and the residuals of this model are taken as a proxy for unexpected movements in the factors. This methodology is a generalisation of the Box-Jenkins transfer function technique and appropriate when forecasting interrelated economic series. Following the intuition of Sims (1980) each variable in the VAR model is treated symmetrically – all of the variables are specified as being endogenous. The time path of each series in the system is specified as being dependent on its own past realisations and both contemporaneous and past values of all of the other variables in the system. Using matrix notation, the general form of such a model in n variables and k lags can be depicted as:

$$By_t = \Gamma_0 + \sum_{j=1}^{k} \Gamma_j y_{t-k} + \varepsilon_t$$

Where:

$$B=\begin{bmatrix} 1 & b_{12} & \cdots & b_{1n} \\ b_{21} & 1 & & b_{2n} \\ \vdots & & \ddots & \vdots \\ b_{n1} & b_{n2} & \cdots & 1 \end{bmatrix}, y_t = \begin{bmatrix} y_{1t} \\ \vdots \\ y_{nt} \end{bmatrix}, \Gamma_0 = \begin{bmatrix} b_{10} \\ \vdots \\ b_{n0} \end{bmatrix}, \Gamma_j = \begin{bmatrix} \gamma_{11j} & \cdots & \gamma_{1nj} \\ \vdots & \ddots & \vdots \\ \gamma_{n1j} & \cdots & \gamma_{nnj} \end{bmatrix} \forall j, \varepsilon_t = \begin{bmatrix} \varepsilon_{1t} \\ \vdots \\ \varepsilon_{nt} \end{bmatrix}$$

B is a symmetric ($n \times n$) matrix of coefficients with unities on the main diagonal (to 'pick out' the dependent variable from y_t in the case of each equation in the system); y_t is the ($n \times 1$) vector of variables included in the VAR system; Γ_0 is a ($n \times 1$) vector of intercept coefficients; Γ_j is a ($n \times n$) matrix of coefficients for each lag $j = 1, \dots, k$ and ε_t is a ($n \times 1$) vector of residual disturbances. Such a system is characterised by feedback as contemporaneous values of y_{1t} and y_{2t}, for example, are allowed to influence each other. If $b_{12} \neq 0$ then a given unit shock to ε_{2t} will affect y_{2t} and indirectly y_{1t}. As a result, the error term ε_{2t} will be correlated with y_t which is a violation of the classical assumptions and, thus, the system cannot be directly estimated using conventional least squares analysis. However, this system may be manipulated into 'reduced form' (where each variable depends on its own lags, *lags* of other endogenous variables and error terms) by premultiplication by B^{-1}:

$$y_t = A_0 + \sum_{j=1}^{k} A_j y_{t-j} + e_t$$

Where: $A_0 = B^{-1}\Gamma_0$

$A_j = B^{-1}\Gamma_j$ for all $j = 1,\dots,k$

$e_t = B^{-1}\varepsilon_t$

Although it is now the case that $E(y_t'e_t) = 0$, due to the premuliplication by B^{-1} and the multicollinearity inherent in this overparameterised model, a typical element of the estimated coefficient matrix A_j cannot be directly interpreted. As a result, these estimates are not reported in the VAR results in Table 2. It is not the purpose of this analysis to draw economic inferences on the dynamics inherent in this model. Rather the VAR methodology is employed as a pragmatic approach to capturing the dominant lagged interrelationships between the variables to forecast the future values of the system's components.

All of the domestic economic series: DLAE (the growth rate of 'All-Share' earnings), DLIE (the growth rate of 'Industrial' earnings), DLGE (the growth rate of 'All-Gold' earnings), DLGOLR (rand gold returns), DR150 (changes in the rate on 10 year gilts), DRBAS (changes in the rate on three month Banker's Acceptances), DLGFX (the growth rate of the level of gold and foreign reserves), DLMMS (the growth rate of the money market shortage), together with DLINDI (returns on the Industrial Index) and DLGOLDI (returns on the All-Gold Index), were included in the VAR model. The equity return variables are included due to (i) the argument that they will reflect anticipations of the other economic series; (ii) to allow for the possibility of bi-directional causality between equity returns and the underlying macroeconomy; and (iii) efficient market considerations prompt an interest in the degree to which the system can predict equity returns.

The cross-correlations between all of the above series were examined in order to inform the decision as to the appropriate lag length of the VAR system. It was found that the majority of significant lagged relations occurred at the earlier lags. In order to capture the bulk of the forecasting power available within the system, it was decided to employ three lags in the VAR model. Thus, investors are assumed to make their (month-ahead) projections based on the most recent values of the economic series analysed as, for example, reported in the Reserve Bank's *Quarterly Economic Review*. Despite the benefits of longer lag structures, the addition of each lag consumes n degrees of freedom due to the multivariate nature of the model. The variable RALSI was also not included in the system as it was felt that the predictive power of this variable would be subsumed in DLINDI and DLGOLDI. Concerns of tractability motivated the decision not to experiment with the inclusion of additional economic series to aid the forecasting of the candidate factors. The results of the VAR model estimated are presented in Table 2.

The reported F statistics indicate that the system contains statistically significant predictive power regarding the future direction of the business cycle, short-term interest rates and the money market shortage. Consistent with notions of informational efficiency, the system does not successfully predict those economic series representing (or containing in their construction) market prices (DLINDI, DLGOLDI, DR150.

Table 2 Vector autoregression results

Panel A: Dependent variable DLAE		
Mean of dependent var: 0.008	R squared: 0.47	F statistic: 2.68
Stnd dev. of dependent var: 0.022	Adjusted R squared: 0.30	Prob (F): **0.00**
Stnd error of regression: 0.019	Log likelihood: 322.39	D.W. statistic: 1.88
Panel B: Dependent variable DLIE		
Mean of dependent var: 0.009	R squared: 0.48	F statistic: 2.76
Stnd dev. of dependent var: 0.025	Adjusted R squared: 0.31	Prob (F): **0.00**
Stnd error of regression: 0.021	Log likelihood: 310.06	D.W. statistic: 1.86
Panel C: Dependent variable DLGE		
Mean of dependent var: 0.002	R squared: 0.51	F statistic: 3.06
Stnd dev. of dependent var: 0.037	Adjusted R squared: 0.34	Prob (F): **0.00**
Stnd error of regression: 0.030	Log likelihood: 267.86	D.W. statistic: 1.84
Panel D: Dependent variable DLINDI		
Mean of dependent var: 0.018	R squared: 0.26	F statistic: 1.02
Stnd dev. of dependent var: 0.055	Adjusted R squared: 0.01	Prob (F): 0.45
Stnd error of regression: 0.055	Log likelihood: 195.35	D.W. statistic:1.88
Panel E: Dependent variable DLGOLDI		
Mean of dependent var: 0.006	R squared: 0.25	F statistic:0.99
Stnd dev. of dependent var: 0.095	Adjusted R squared: 0.00	Prob (F): 0.50
Stnd error of regression: 0.095	Log likelihood: 129:56	D.W. statistic: 2.00
Panel F: Dependent variable DR150		
Mean of dependent var: 0.000	R squared: 0.30	F statistic: 1.28
Stnd dev. of dependent var: 0.005	Adjusted R squared: 0.07	Prob (F): 0.18
Stnd error of regression: 0.005	Log likelihood: 495.30	D.W. statistic 1.99
Panel G: Dependent variable DLRBAS		
Mean of dependent var: –0.001	R squared: 0.51	F statistic: 3.06
Stnd dev. of dependent var: 0.005	Adjusted R squared: 0.34	Prob (F): **0.00**
Stnd error of regression: 0.004	Log likelihood: 498.18	D.W. statistic: 1.99
Panel H: Dependent variable DLGFX		
Mean of dependent var: 0.008	R squared: 0.30	F statistic: 1.27
Stnd dev. of dependent var: 0.08	Adjusted R squared: 0.06	Prob (F): 0.19
Stnd error of regression: 0.08	Log likelihood: 154..59	D.W. statistic: 1.97
Panel I: Dependent variable DLMMS		
Mean of dependent var: 0.003	R squared: 0.36	F statistic: 1.64
Stnd dev. of dependent var: 0.426	Adjusted R squared: 0.14	Prob (F): **0.04**
Stnd error of regression: 0.395	Log likelihood:- 40.82	D.W. statistic: 2.00
Panel J: Dependent variable DLGOLR		
Mean of dependent var: 0.007	R squared: 0.24	F statistic: 0.94
Stnd dev. of dependent var: 0.046	Adjusted R squared: –0.02	Prob (F): 0.56
Stnd error of regression: 0.046	Log likelihood: 215.54	D.W. statistic: 1.92

(p values of F statistics significant at the 95% level are in bold)

construction) market prices (DLINDI, DLGOLDI, DR150, DLGFX and DLGOLR). The fact that the standard error of each of these series is identical to the standard error of the residual of its forecasting equation (rounded to three places) clearly confirms this inference. In these cases the mean of the dependent variable is its optimal forecast value.

The multi-equation residuals were taken as estimates of unexpected movements in the candidate series and prefixed by a '*U*'. All were found to be characterised by a mean value of zero. Ljung-Box statistics conducted over 12 lags also found all of these series to be void of autocorrelation, with the exception of UDLGOLR. The correllogram of the latter series revealed significant fifth and sixth order autocorrelation and partial autocorrelation. As a result, UDLGOLR was regressed on values of itself lagged by five and six periods and the residuals of this model were taken as revised values of this factor. The international series DLDJ (which was not included in the VAR model) was found to be void of autocor-

Table 3 Correlation matrix: candidate factors (1985–1995)

	UDLAE	UDLIE	UDLGE	UDLDJ	UDRBAS	UDR150	UDLGFX	UDLMMS	UDLGOLR
DLALSI	**0.20**	0.13	0.13	**0.41**	–0.01	**–0.23**	**0.18**	–0.02	**0.23**
DLINDI	0.06	–0.02	0.08	**0.54**	–0.06	**–0.18**	**0.17**	–0.01	–0.01
DLGOLDI	**0.20**	**0.23**	**0.14**	0.07	–0.06	**–0.15**	**0.17**	–0.01	**0.37**
UDLIE	**0.58**	**1**							
UDLGE	**0.35**	0.10	**1**						
ULDJ	0.03	–0.05	0.01	**1**					
UDRBAS	0.08	**0.20**	0.11	–0.06	**1**				
UDR150	**–0.16**	–0.13	–0.13	–0.06	**0.36**	**1**			
UDLGFX	0.03	0.06	0.03	0.01	0.05	–0.10	**1**		
UDLMMS	0.12	**0.15**	0.09	0.10	0.11	**–0.18**	–0.08	**1**	
DLGOLR	**0.24**	**0.30**	**0.18**	**–0.17**	**0.15**	–0.03	**0.25**	–0.04	**1**

(relationships significant at the 95% level are in bold)

relation over 12 lags consistent with Fama (1970) 'weak form' market efficiency.

The mean value of DLDJ over the sample period was, thus, subtracted from its realised values to derive a proxy for unexpected movements in this variable.The correlations of the candidate factors estimated in the manner described above both with each other and with DLALSI, DLINDI and DLGOLDI are reported in Table 3.

The upper portion of Table 3 presents the correlations of each of the candidate factors derived above with returns on the major sectoral equity indices. As the variables UDRBAS and UDLMMS were not found to have significant relations with the time series of equity returns over the sample period, they were omitted from further analysis. All of the correlations observed have signs consistent with those implied by the dividend discount model of equity valuation.

The lower portion of Table 3 tabulates the correlations of the candidate explanatory variables with each other. Due to the close relation between UDLAE and UDLGE, only the former was retained for cross-sectional testing. The most pronounced cases of multicollinearity between the candidate factors exist between UDLGOLR and UDLGFX and between UDLGOLR and UDLAE. The variables UDLGFX and UDLAE also share significant correlations with UDR150 explaining the dilution of their *ceteris paribus* influence in the multiple regression model presented in Panel A of Table 4 in the following section.

Factor analytic augmentation revisited

Van Rensburg (1997) points out that the assumption that $E(\varepsilon_{it}\varepsilon_{jt}) = 0$ for all $i \neq j$, which underlies all linear factor models, is likely to be violated in those specifications of the JSE return generating process that employ prespecified macroeconomic series as explanatory forces in the manner of Chen, Roll & Ross (1986).[6] A factor analytic augmentation to the prespecified variable approach is suggested. The rationale underlying this procedure can be presented as follows: assume that we have (mis)specified the following return generating process:

$$R_{it} = \alpha_i + \sum_{g=1}^{G} b_{ig} f_{gt} + \varepsilon_{it} \tag{3}$$

where: $E(\varepsilon_{it}\varepsilon_{jt}) \neq 0$ for all $i \neq j$.

The resulting contemporaneous correlation of residuals across shares can be thought of as being generated by the following process:

$$\varepsilon_{it} = \left(\sum_{h=1}^{H} b_{ih} f^{*}_{ht} + \xi_{it} \right) \tag{4}$$

where: f^{*}_{ht} = the principal factor score of factor h at time t
b_{ih} = the factor loading of asset i to factor h
ξ_{it} = the residual error at time t, where $E(\xi_{it}\xi_{jt}) = 0$ for all $i \neq j$.

Seen in this light, the specification error associated with (3) can be seen as a problem resulting from the omission of relevant explanatory variables. A natural solution is to substitute (4) into (3) and to estimate the resulting model:

$$R_{it} = \alpha_i + \sum_{g=1}^{G} b_{gm} f_{mt} + \sum_{h=1}^{H} b_{ih} f^{*}_{ht} + \xi_{it} \tag{5}$$

Fortunately, the nature of the problem ($E(\varepsilon_{it}\varepsilon_{jt}) \neq 0$ for all $i \neq j$) is such that its factor analytic identification and incorporation within the model allows not only its solution but also reveals potentially important information regarding the return generating process underlying the security. Information regarding the comovement of securities returns, which was previously relegated to the error term of the LFM, is explicitly utilised (as manifested in a set of factor scores) as a contributory explanatory force in the LFM. The properties of the estimators in the underspecified model 3 are discussed in Gujarati (1988: 403–404) and are derived in Kmenta (1990: 443–446):

(i) The estimator of α_i will be unbiased as the mean value of $(f_{gt}^{*}) = 0$, for each $g = 1, \ldots, G$.

(ii) The estimators of b_{ik} will be unbiased assuming that $E(f_{gt}\varepsilon_{it}) = 0$, for each $g = 1, \ldots, G$.

(iii)The variance of the estimators of both α_i and b_{ig} will be upwardly biased as $\sigma_\varepsilon > \sigma_\xi$. As a result, *prior South African research employing prespecified macroeconomic variables to explain JSE returns has overly tended to accept the null hypothesis of no relation being present between the series investigated and JSE returns!* This malady is avoided by using the two residual market factor approach outlined below which can be viewed as an econometric correction for omitted variable bias.

Following the findings of Van Rensburg & Slaney (1997), reviewed earlier in the article, returns realised on the JSE Industrial and All-Gold Indices are employed as convenient observable proxies for the first two factor analytically extracted factors. Accordingly, two 'residual market factors' are estimated using the residuals of ordinary least squares and included in the analysis reported in Table 4:

$$DLGOLDI_t = \alpha_{gol0} + b_{gol1}UDLAE_t + b_{gol2}UDLDJ_t + b_{gol3}UDR150_t + b_{gol4}UDLGFX_t + b_{gol5}\,DLGOLR_t + \varepsilon_{golt}$$

where: $UG_t = \varepsilon_{golt}$ = *the 'All-Gold' residual market factor.*

$$DLINDI_t - \alpha_{ind0} + b_{ind1}UDLAE_t + b_{ind2}UDLDJ_t + b_{ind3}UDR150_t + b_{ind4}UDLGFX_t + b_{ind5}\,UDLGOLR_t + \varepsilon_{indt}$$

where: $UI_t = \varepsilon_{indt}$ = *the 'Industrial' residual market factor.*

Table 4 compares the results when the 'All-Gold' and 'Industrial' residual market factors are included as explanatory forces when describing the return generating process on the JSE All-Share Index.

It is evident that, while the coefficent estimates of the underspecifed model in panel A are identical to those estimated in the augmented model reported in panel B, the standard are smaller in the latter case errors (and consequently the t statistics are larger). As a result, the *ceteris paribus* influences of DLAE and UDLGFX now pass tests of statistical significance at the 5% level. Had this augmentation not been conducted, the resulting estimation bias would mislead the researcher into inferring that these variables do not contribute to explaining the time-series of equity returns. Further, the large t statistics associated with UG and UI and the surge in adjusted R^2 from 0.26 to 0.90 when these factors are included in the model, confirm the relevance of this adjustment when employing macroeconomic forces to describe the return generating process operational on the JSE.

Table 5 reports the results when the series UDLINDI and UDLGOLDI are employed as explanatory variables instead of the residual market factors UI and UG. UDLINDI and UDLGOLDI respectively represent deviations of DLINDI and DLGOLDI from their mean values and Ljung Box statistics indicate that both of the series are void of serial correlation up to twelve lags. Unlike the residual market factors, UI and UG, these variables represent the entire variation in the Industrial and All-Gold Indices without controlling for the influence of the prespecified macrovariables.

Table 4 Multiple regression results (dependent variable: DLALSI)

Panel A: Linear factor model results				
Variable	Coefficient	Standard error	t-statistic	2-tail significance
C	**0.019**	0.005	4.074	**0.000**
UDLAE	0.336	0.285	1.178	0.241
UDLDJ	**0.510**	0.102	5.002	**0.000**
UDR150	**−2.871**	1.182	−2.428	**0.017**
UDLGFX	0.094	0.071	1.336	0.184
UDLGOLR	**0.371**	0.130	2.863	**0.005**
Mean of dependent var: 0.017		R squared: 0.29		F statistic: 9.34
Stnd dev. of dependent var: 0.056		Adjusted R squared: 0.26		Prop (F):**0.00**
Stnd error of regression: 0.050		Log likelihood: 191.14		D.W. statistic: 1.97
Panel B: 'Augmented' linear factor model results				
Variable	Coefficient	Standard error	t-statistic	2-tail significance
C	**0.019**	0.002	11.026	**0.000**
UDLAE	**0.336**	0.105	3.189	**0.002**
UDLDJ	**0.510**	0.038	13.535	**0.000**
UDR150	**−2.871**	0.437	−6.570	**0.000**
UDLGFX	**0.094**	0.026	3.615	**0.001**
UDLGOLR	**0.371**	0.048	7.746	**0.000**
UG	**0.317**	0.020	15.461	**0.000**
UI	**0.666**	0.039	16.988	**0.000**
Mean of dependent var: 0.017		R squared: 0.91		F statistic: 151.16
Stnd dev. of dependent var: 0.056		Adjusted R squared: 0.90		Prop (F):**0.00**
Stnd error of regression: 0.02		Log likelihood: 310.65		D.W. statistic: 1.97

(coefficients and p values of t and F statistics significant at the 95% level are in bold)

Table 5 Multiple regression results (dependent variable: DLALSI)

Economic forces and the two index model				
Variable	Coefficient	Standard error	t-statistic	2-tail significance
C	**0.018**	0.002	10.924	**0.0000**
UDLAE	0.178	0.103	1.732	0.0860
UDLDJ	0.055	0.043	1.281	0.4174
UDR150	–0.476	0.432	–1.102	0.273
UDLGFX	–0.021	0.026	–0.814	0.417
UDLGOLR	0.023	0.050	0.468	0.641
UDLGOLDI	**0.321**	0.020	16.170	**0.0000**
UDLINDI	**0.659**	0.038	17.335	**0.0000**
Mean of dependent var: 0.017	R squared: 0.91		F statistic: 165.77	
Stnd dev. of dependent var: 0.056	Adjusted R squared: 0.90		Prop (F):**0.00**	
Stnd error of regression: 0.02	Log likelihood: 314.48		D.W. statistic: 1.97	

(coefficients and p values of t and F statistics significant at the 95% level are in bold)

The t statistics of the coefficients reported in Table 5 indicate that none of the macrovariables offer a marginal contribution to the two index model (using only returns on the JSE Industrial and All-Gold Indices as explanatory variables) that is significant at the 95% level of confidence. A test for omitted variables with F statistic of 1.526 with a p value of 0.188 confirms that the cumulative contribution of the macrovariables is not significant at conventional levels. Thus, the important result can be inferred that *the two index model proposed by Van Rensburg & Slaney (1997) subsumes the influence of the other macroeconomic variables and, in this sense, offers a parsimonious representation of the influence of economic forces on JSE listed shares.*

Cross-sectional analysis

The defining characteristic of a priced factor is that it explains the cross-section of expected returns or, equivalently, that it is associated with a non-zero risk premium. The cross-sectional analysis is conducted using the iterated non-linear seemingly unrelated regression (ITNLSUR) methodology. In this model the sensitivity coefficients (b_{ik}s) for each share in the sample and the risk premia (λ_ks) associated with each factor k are measured simultaneously. Those factors associated with statistically significant risk premia are identified as being priced. The motivation for adopting this more powerful methodology over the two-step Fama-MacBeth (1973) technique used by Chen, Roll & Ross (1986) and Page (1986) is outlined by McElroy & Burmeister (1988: 32-33):

(i) Incorporating seemingly unrelated regression techniques into the non-linear regression allows across-equation restrictions to be specified which are the exact pricing restrictions postulated by the APT model.
(ii) There is no need to partition the assets into portfolios in order to avoid the 'errors in the variables' problem.[7] This procedure consumes a large number of degrees of freedom, considerably diminishing the power of the cross-sectional analysis. For South African researchers this is a serious practical problem due the relatively small size of JSE share samples. Further, both Campbell (1979) and Chen, Roll & Ross (1986) find that the (essentially arbitrary) criteria under which the portfolios are formed meaningfully impact on the results of the cross-sectional analysis.
(iii) The NLSUR procedure is robust with respect to the non-normality of the distribution of asset returns and factor values. Even in the absence of normally distributed error terms, the estimates obtained through this technique will be strongly consistent, asymptotically normally distributed and able to be utilised in standard hypothesis testing.
(iv) If it is the case that the error terms are normally distributed then iterating on the contemporaneous covariance matrix (ITNLSUR) provides full information maximum-likelihood estimators.

The following system was initially estimated using the ITNLSUR technique:

$$R_{it} - R_{ft} = \sum_{k=1}^{K} b_{ik}\lambda_k + \sum_{k=1}^{K} b_{ik} f_{kt} + \varepsilon_{it} \qquad (6)$$

for i = 1, ... n and t = 1, ... T.

where: R_{it} = realised returns on asset i in time period t
R_{ft} = the risk free rate of return at time t
b_{ik} = the sensitivity of asset i to factor k
λ_k = the risk premium associated with factor k
f_{kt} = the unexpected movement in factor k at time t
ε_{it} = the error term for asset i at time t

Equation 6 may be written in matrix notation as:

$$\rho_i = \sum_{k=1}^{K} (\lambda_k \iota_T + f_k) b_{ik} + \varepsilon_i$$

where: $\rho_i = (R_{i1} - R_{f1}, \ldots, R_{iT} - R_{fT})'$ for i = 1, ... n
$f_k = (f_{k1}, \ldots, f_{kT})'$ for k = 1, ... K
$\varepsilon_i = (\varepsilon_{i1}, \ldots, \varepsilon_{iT})'$ for i = 1, ... n
and ι_T is a T dimensional column vector of ones.

Using the Kronecker or direct product operator, $\otimes$, this system may be condensed as :

$$\rho_i = X(\lambda)b_i + \varepsilon_i$$

where: $X(\lambda) = (\lambda' \otimes \iota_T) + F$

$\lambda = (\lambda_1, \ldots, \lambda_K)'$

$F = (f_1, \ldots, f_K)$

$b_i = (b_{i1}, \ldots, b_{iK})'$ for $i = 1, \ldots n$

Stacking the n equations yields

$$\begin{bmatrix} \rho_1 \\ \rho_2 \\ \vdots \\ \rho_n \end{bmatrix} = \begin{bmatrix} X(\lambda) & 0 & \cdots & 0 \\ 0 & X(\lambda) & & \vdots \\ \vdots & & \ddots & 0 \\ 0 & \cdots & 0 & X(\lambda) \end{bmatrix} \begin{bmatrix} b_1 \\ b_2 \\ \vdots \\ b_n \end{bmatrix} + \begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \vdots \\ \varepsilon_n \end{bmatrix}$$

Which can, in turn, be more concisely stated as:

$$\rho = [I_n \oplus X(\lambda)]b + \varepsilon$$

where: $\rho = (\rho_1, \ldots, \rho_n)'$

$b = (b_1, \ldots b_n)'$

$\varepsilon = (\varepsilon_1, \ldots, \varepsilon_n)'$

and $E(\varepsilon) = 0_{nT}$ and $E(\varepsilon\varepsilon') = [\Sigma \otimes I_T]$, where Σ is the $n \times n$ variance-covariance matrix of the contemporaneous residuals of assets $i = 1, \ldots n$ and $j = 1, \ldots n$. NLSUR estimators may be obtained in three steps:

(i) The linear K factor model is estimated via share-by-share OLS. This is the same procedure as followed in the first step of the Fama & MacBeth (1973) 'two step' procedure. As a result a vector of coefficients $b_i = (b_{i0}, b_{i1}, \ldots, b_{iK})$ and a T dimensional vector of residual errors ε_i is estimated for each share i.

(ii) Unlike in the two step procedure, the output utilised is not b_i but the residual vectors which are used to estimate S with typical element $[\delta_{ij}^2] = [T^{-1}\varepsilon_i\varepsilon_j]$.

(iii) The estimated variance-covariance matrix is plugged in to the following quadratic form, Q:

$$Q(\lambda, b, \Sigma) = [\rho - (I_n \oplus X)(\lambda)b]'\left[\Sigma^{-1} \oplus I_T\right][\rho - (I_n \oplus X(\lambda)b)] \quad (7)$$

Values for λ and b are chosen so as to minimise the value of this expression. It can be seen that a sum of squared errors, weighted by the estimated variance-covariance matrix Σ, is being minimised with respect to λ and b. The procedure above may be repeated, iteration occurring between the estimates of Σ and the parameters λ and b. Note that, unlike in the initial step (i) described above, in these iterations the across equation restrictions manifested in equation (7) are imposed on the system. Residuals from the most recent estimates of λ and b are used to update the estimate of Σ. This in turn updates the quadratic form Q, allowing revised estimates of λ and b. Iteration occurs until estimates of the covariance matrix, Σ, stabilise. This estimation technique is called an iterated non-linear seemingly unrelated regression (ITNLSUR).[9]

The following three model specifications were estimated using the ITNLSUR technique:

(i) The 'exact arbitrage pricing restrictions' model as presented above:

$$R_{it} - R_{ft} = \sum_{k=1}^{K} b_{ik}\lambda_k + \sum_{k=1}^{K} b_{ik}f_{kt} + \varepsilon_{it} \quad (8)$$

(ii) The 'unconstrained intercept' model:

$$R_{it} - R_{ft} = \lambda_0 + \sum_{k=1}^{K} b_{ik}\lambda_k + \sum_{k=1}^{K} b_{ik}f_{kt} + \varepsilon_{it} \quad (9)$$

The 'zero beta' version of the APT:

$$R_{it} = \lambda_0 + \sum_{k=1}^{K} b_{ik}\lambda_k + \sum_{k=1}^{K} b_{ik}f_{kt} + \varepsilon_{it} \quad (10)$$

In each case, $i = 1, \ldots n$ and $t = 1, \ldots T$.

The specification of the 'unconstrained intercept' model follows directly from a suggestion of McElroy, Burmeister & Wall (1985: 274) and is used as a convenient test of the APT pricing restrictions: '...the APT dictates that the remaining intercept common to each portfolio be zero. The model is readily estimated and the zero-intercept restriction tested'.

The 'exact APT' model uses the three-month treasury bill rate as a proxy for the default free rate. Firer (1993) conducted a review of South African studies that required a proxy for the 'risk free asset'[x]: De Villiers, Lowlings, Pettit & Affleck Graves (1986) utilised 360 day treasury bills; Bradfield, Barr & Affleck Graves (1988) the 12 month fixed deposit rate; while Affleck Graves, Burt & Cleasby (1988) and Page & Palmer (1991) used the 90 day treasury bill rate. Van Rensburg (1997) found that risk premia estimates were insensitive to changes in default free proxies, comparing the findings when the 12 month fixed deposit rate, the 90 day treasury bill rate and the interest rate on 10 year gilts were successively utilised for this purpose. In this study, the sensitivity of the analysis to the default-free proxy is investigated by comparing the premia estimated by the 'zero-beta' specification of the APT. In this model the intercept term, λ_0, represents the 'premium' associated with an asset with a zero sensitivity to all factors (where $b_{i1} = b_{i2} \ldots = b_{iK} = 0$) and is estimated from the data. Like the premia associated with the risk factors, λ_0 is constrained to be constant across time and assets.

Following the arguments in the previous section, the 'two residual market factor' approach is adopted throughout the analysis. The ITNLSUR results are summarised in Table 6.

In all cases, it was found that the signs and magnitudes of the estimated risk premia were robust across the 'exact APT', 'unconstrained intercept' and 'zero beta' model specifications. Further, where the unconstrained intercept model was estimated, the intercept term was found to be insignificantly different from zero, implying the validity of the APT pricing restrictions.

As expected, the two index model results reported in panel A of Table 6 are consistent with those of Van Rensburg & Slaney (1997) who found both factor analytic proxies to be priced. Panels B to F display three factor model results where UDLAE, UDLDJ, UDLGFX, UDR150 and UDLGOLR are successively included as candidate APT factors together with their associated dual residual market factors. It was found that, while UDLAE was not priced, the remaining series were found to be associated with risk premia significantly different

Table 6 Cross-sectional results (estimates of risk premia are reported)

	Exact APT restrictions			Unconstrained intercept model			'Zero beta' APT model		
	Coefficient	t statistic	p>|t|	Coefficient	t statistic	p>|t|	Coefficient	t statistic	p>|t|
Panel A: The two index model									
λ_0	–	–	–	–0.00099297	–0.41	0.6850	0.010330	4.15	0.0001
λ_{UIND}	0.015109	24.23	0.0001	0.00673415	2.67	0.0088	0.00673415	2.67	0.0088
λ_{UGOLD}	0.00580355	8.01	0.0001	0.016037	6.84	0.0001	0.016037	6.84	0.0001
Panel B: UDLAE and the two index residual market factor									
λ_0	–	–	–	–0.00089003	–0.36	0.7176	0.010428	4.17	0.0001
λ_{UDLAE}	–0.00116744	–1.09	0.2791	–0.00112817	–1.04	0.2993	–0.00112817	–1.04	0.2993
λ_{UIND}	0.015103	24.63	0.0001	0.0015930	6.79	0.0001	0.0015930	6.79	0.0001
λ_{UGOLD}	0.00719541	4.80	0.0001	0.00798023	2.88	0.0047	0.00798023	2.88	0.0047
Panel C: UDLDJ and the two index residual market factor									
λ_0	–	–	–	0.00135158	0.53	0.5986	0.012632	4.85	0.0001
λ_{UDLDJ}	–0.00700294	–2.78	0.0064	–0.00796424	–2.80	0.0060	–0.00796424	–2.80	0.0060
λ_{UIND}	0.019230	12.76	0.0001	0.018528	8.44	0.0001	0.018528	8.44	0.0001
λ_{UGOLD}	0.00687572	7.93	0.0001	0.00566158	2.24	0.0269	0.00566158	2.24	0.0269
Panel D: UDLGFX and the two index residual market factor									
λ_0	–	–	–	–0.00281888	–1.08	0.2807	0.00876325	3.19	0.0018
λ_{UDLGFX}	–0.00864763	–1.92	0.0576	–0.011740	–2.46	0.0154	–0.011740	–2.46	0.0154
λ_{UIND}	0.015894	21.65	0.0001	0.018817	7.12	0.0001	0.018817	7.12	0.0001
λ_{UGOLD}	0.00828852	5.42	0.0001	0.0011970	3.62	0.0004	0.0011970	3.62	0.0004
Panel E: UDR150 and the two index residual market factor									
λ_0	–	–	–	–0.00071474	–0.29	0.7756	0.010654	4.18	0.0001
λ_{UDR150}	0.00062786	3.14	0.0022	0.00063418	3.12	0.0023	0.00063418	3.12	0.0023
λ_{UIND}	0.016768	20.56	0.0001	0.017455	7.34	0.0001	0.017455	7.34	0.0001
λ_{UGOLD}	0.00795211	7.58	0.0001	0.00868213	3.33	0.0012	0.00868213	3.33	0.0012
Panel F: UDLGOLR and the two index residual market factor									
λ_0	–	–	–	–0.00102073	–0.41	0.6850	0.010341	4.12	0.0001
$\lambda_{UDLGOLR}$	0.00687407	3.64	0.0004	0.00704165	3.64	0.0004	0.00704165	3.64	0.0004
λ_{UIND}	0.015056	23.92	0.0001	0.016023	6.77	0.0001	0.016023	6.77	0.0001
λ_{UGOLD}	0.00259098	1.21	0.2285	0.0029773	1.15	0.2523	0.0029773	1.15	0.2523
Panel G: Priced macrovariables and and the two index residual market factor									
λ_0	–	–	–	–0.00096989	–0.35	0.7302	0.010307	3.61	0.0005
λ_{UDLDJ}	–0.00761949	–2.84	0.0054	–0.00703940	–2.33	0.0217	–0.00703940	–2.33	0.0217
λ_{UDLGFX}	–0.011099	–2.21	0.0292	–0.011829	–2.28	0.0244	–0.011829	–2.28	0.0244
$\lambda_{UDLGOLR}$	0.00529544	2.51	0.0135	0.00541099	2.51	0.0134	0.00541099	2.51	0.0134
λ_{UDR150}	0.00056332	2.58	0.0112	0.00054974	2.48	0.0148	0.00054974	2.48	0.0148
λ_{UIND}	0.201524	12.46	0.0001	0.02117	8.86	0.0001	0.02117	8.86	0.0001
λ_{UGOLD}	0.00992252	3.51	0.0007	0.010692	3.00	0.0033	0.010692	3.00	0.0033

(coefficients and p values significant at the 95% level are in bold)

from zero at the 95% level of confidence. The specification of UDLGOLR resulted in UGOLDI being unable to explain any of the remaining cross-sectional variation in returns. This implies that much of the priced risk in UGOLDI represents gold price risk.

In panel G all of the candidate factors found to be priced in the preceding analysis are included in the ITNLSUR model. It was found that each of the macroeconomic factors found to be priced individually are also priced in the presence of the other candidate series. The risk premia estimated are again found to be robust in terms of size and sign. Interpreting the signs associated with the premiums estimated, it is evident that while UDLDJ exhibits a positive correlation with JSE equity returns (see Table 3), the premium associated with exposure to this variable is significantly negative. The converse is the case of UDR150. Both of these results violate the pricing implications of the CAPM as derived in Sharpe (1984).[10] These results differ from those of Van Rensburg (1996, 1997), who found that gold price risk was not priced while short-term interest rate risk was priced over the period 1980–1990. In addition, the earlier studies found the Dow-Jones Industrial Index to be associated with a significantly positive risk premium rather than the negative premium estimated in this study. The comparison of these results reveals the dynamic nature of the process underlying the cross-section of JSE returns and motivates the appropriateness of time-varying parameter models in future JSE research.

Conclusion

The dichotomy in the return generating process underlying South African mining and industrial share returns will lead to a violation of the diagonality assumption of all linear factor models used to model the time-series of JSE returns that do not employ factor analytically derived explanatory variables. Using the Industrial and All-Gold Indices as observable proxies for the first two factors, an approach employing *two* Burmeister & Wall (1986) residual market factors was introduced in this study to avoid this misspecification error. It was argued that not only does this procedure significantly improve the explanatory power of models using prespecified macroeconomic variables, but also that its omission leads to upward bias in the variances of the coefficient estimators of these models. This result has far reaching and far from obvious implications, suggesting that prior research investigating the relationship between macroeconomic forces and the JSE has been biased towards accepting the null hypothesis of no significant relation being present. It was also found that a two index model, using returns on the All-Gold and Industrial Indices as explanatory variables subsumes the influence of the other prespecified macrovariables on equity returns. In this sense, the two index model of Van Rensburg and Slaney (1997) provides a parsimonious account the relationship between macroeconomic variables and the JSE.

The cross-sectional analysis revealed that the Dow-Jones Industrial Index, the rate on long bonds, the rand gold price and the level of gold and foreign reserves (together with the All-Gold and Industrial residual market factors) represented priced sources of risk on the JSE over the period 1985–1995. In the case of the former two variables the signs of the risk premia estimated were inconsistent with those implied by the CAPM. These findings were found to be robust across the 'unconstrained intercept' and 'zero beta' model specifications.

The direction for future research most clearly implied by this study is the empirical investigation of the ability of accounting ratios and other 'anomalies' to explain the cross-section of JSE returns beyond that of the two factor APT model of Van Rensburg & Slaney (1997).

Notes

1. The diagonality assumption postulates that the contemporaneous residual errors of a particular linear factor model are uncorrelated across securities: 'There is one assumption that makes a factor model less than vacuous. It is assumed that the ε values are uncorrelated with one another. In other words the returns of two securities will be correlated ... only through common reactions to one or more of the factors' (Sharpe, 1984: 182).
2. The residual market factor represents that variation in the 'market' that cannot be explained by a particular set of prespecified variables and is estimated by extracting the residuals of an ordinary least squares regression of a market proxy on these variables. When used as an explanatory force alongside the variables, the residual market factor play a useful 'catch all' role and allays fears of specification errors due to omitted variables. However, as will be argued below, a single residual market factor does not adequately capture the mining-industrial dichotomy on the JSE.
3. Bradfield (1989) found that over the 500 week period from 1 January 1978 to 31 August 1987, on average about one third of the survivors were not traded for at least one *consecutive* week in every four weeks considered. In 1989 it was estimated that the liquidity of the JSE (as measured by the turnover of shares as a percentage of market capitalisation) was only 5.5% compared to the 53% and 74% enjoyed on the New York and London Stock exchanges respectively (*Economic Focus*, 1990). McGregor estimates that only a small subsegment, representing about 10–15% of the shares listed on the JSE, are actually actively traded (*The JSE Centenary Publication*, 1987: 134). Of relevance to future researchers is the sharp rise in turnover that has occurred since the introduction of the JET system in 1996.
4. Since April 1978 information on the South African Reserve Bank's level of gold and foreign reserves has been calculated at the end of each month at 90% of the mean of the previous 10 days London gold price fixings and published in the 'Statements and Liabilities' of the SARB. The Reserve Bank publicly announces its month-end levels on the first working day of the next month. This will result in a *one day* 'look ahead' effect (personal correspondence, Andrew Cantor, Rand Merchant Bank Asset Management). Based on expectations formed on the previous working day, the price adjustments resulting from the unexpected portion of this announcement will occur on the first day after the month to which the announcement refers. The relative intensity of these price adjustments, compared to those experienced over the remaining 'synchronous' days of the month, empirically motivate whether it is more appropriate to lag this variable in the ensuing analysis. Unreported tests were conducted comparing the influence of both lagged and contemporaneous growth in the level of gold and foreign exchange reserves on ALSI returns. It was found that, while the lagged (announcement) influence was negligible, the contemporaneous (rest of month) relationship was statistically significant at the 99% level of confidence. The vector autoregressive model described below was also conducted using both contemporaneous and lagged values of DLGFX. It was found that using contemporaneous values did not improve the forecasts of any of the other macroeconomic series at conventional levels of significance, implying that the magnitude of any possible 'look ahead' bias is not such that it will meaningfully distort the extraction of unexpected movements in the factors. Thus, the data suggests that it is more appropriate to employ contemporaneous rather than lagged values of this variable in the analysis.
5. 'Apart from spanning the space of returns the most important property required of appropriate factor measures is that they have a zero expectation at the beginning of month t. In particular a macroeconomic factor measure cannot be predictable from its own past' (McElroy & Burmeister, 1988: 31).
6. Empirical investigation is the only appropriate avenue to ascertain whether this conjecture is, in fact, appropriate for the particular LFM specification under consideration. Regressions conducted on market model residuals by Page (1986) and Van Rensburg & Slaney (1997) find that its contemporaneous errors are significantly correlated across industrial and mining shares.
7. This problem arises in the Fama-MacBeth approach due to the fact the factor sensitivities estimated in time-series regressions are used as dependent variables in the cross-sectional 'second step'. The formation of portfolios prior to the second step is

employed to diversify away the estimation errors of the first step.

8. McElroy & Burmeister (1988: 31–32). See Greene (1990: 509–540) for a textbook account of seemingly unrelated regression techniques. The methodology was originally developed by Zellner (1962) and its application to asset pricing estimation suggested by Gibbons (1982).
9. Despite common terminology, these proxies for the 'risk free' asset are, more precisely, measures of the returns on **default free** securities, that is where there is no uncertainty as to the (nominal) magnitude and timing of the future cash-flows associated with ownership of the asset. Revisions in inflation expectations (which could be proxied by a term structure variable) can, for example, be reasonably conjectured as a macroeconomic risk factor to which both share and gilt returns are likely to be exposed (see Fama & French, 1993). The more correct nomenclature is adopted forthwith.
10. Sharpe (1984) demonstrates that if both the CAPM and the APT pricing restrictions hold, the following relationship must hold for all priced factors $k = 1,...,K$:

    $\lambda_k = \lambda_m \beta_k$

    where: λ_m= the 'market' risk premium; β_k = the CAPM beta of factor k.

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