European Research Studies Vol.
III, Iss. (3-4), 2000, pp. 131-144

The Pricing Of Risk Factors And The UK Insurance Stocks' Performance A Nonlinear Multivariate Approach

Elias Dinenis and Sotiris K. Staikouras[*] [♣]
City University Business School

Abstract

The objective of the present study is to examine the impact of exchange and interest rate changes on the common stock returns of the insurance companies in the UK. All general and life insurance firms listed in the London Stock Exchange are selected for this purpose. An augmented market model with the additional variables of the interest and exchange rate indices is employed to test both the pricing question and the factor sensitivity of the particular sample. A seemingly unrelated regression (SURE) multivariate estimation with both cross–equation restrictions and within equation nonlinear constraints on the parameters is employed. This method eliminates the errors in variable (EIV) problem and the estimates are strongly consistent and asymptotically normal even without the assumption of normally distributed errors. The two main implications of this investigation are as follows. First both kinds of insurance companies are negatively and equally affected by unanticipated changes in interest rates. Second the changes in exchange rates seem to inversely affect the general insurance companies, while the life insurance firms seem to be insensitive.

Key Words: Insurance stock returns, Interest and exchange rates, APT, Kalman filter, Nonlinear SURE modelling.

1. Introduction

The increased trend in interest rate volatility during the past two decades has been contended by academics and practitioners to be one of the possible reasons for the variations in equity prices. Common stocks of insurance companies could be cited as being susceptible to interest rate risk. This arises because of their propensity to undertake an asset transformation and intermediation function. An explanation for the above issue is offered by the nominal contracting hypothesis (French *et al.* 1983) where if interest rates change and the firm holds a large proportion of long–term assets, the market value of its assets will change more than if

[*] An earlier version of this paper was presented at the 1996 V Tor Vergata Financial Conference, Financial Markets: Imperfect Information and Risk Management, University of Rome, Italy.

[♣] We would like to thank Dr Giovanni Urga for his valuable comments and suggestions.
Centre for Mathematical Trading & Finance Frobisher Crescent, Barbican Centre, London EC2Y 8HB, Tel: 0171–4778694 – Fax: 0171–4778885.

the assets were short–term. Consequently, as the market value of a firm's common stock is tied to the value of its assets the more volatile their values, the more volatile the market value of its stock. A discussion related to the above issues is presented in section II. However, the relationship between stock prices and changes in market yields is less clear–cut. Nevertheless, based on empirical evidence, some generalizations can be made. On an ex–post basis, recent research indicates a relationship between changes in interest rates and the market value of common stocks [Bae (1990), Saunders & Yourougou (1990), Dinenis & Staikouras (1996, 1998)].

The aim of this paper is to investigate the sensitivity of UK insurance companies' stock returns to surprises in interest and exchange rates. It is motivated primarily by the lack of similar studies for the UK by examining both the sensitivity and the pricing relationship of the insurance industry. The remainder of the paper is constructed as follows. The next two sections describe the intuition behind the nominal contracting hypothesis, the data used and the methodology employed. A description of the expectation generating process and the results of the multivariate estimation for the three factors are shown in sections IV and V respectively. The last section presents an overview of our study, a general summary of the issues and draws some implications for future research.

2. The duration mismatch hypothesis

The concept of the maturity or better the duration of an income stream has been employed in a variety of contexts in applied economics and finance. In the current work, the term refers to the structure of the balance sheet to give a possible explanation of the interest rate sensitivity. It is actually true that securities, which are claims on monetary assets, such financial intermediary common stock should exhibit a covaration with the movements of market yields[1]. The impact, therefore, of changes in market yields on the common stock value of these firms will depend on the maturity composition (in the sense of "time to repricing") of their assets and liabilities (Flannery & James 1984). As an extreme example, consider an unlevered firm whose only asset is a consol. If all earnings are promptly paid out as dividends, the firm's stock price should move precisely with the consol's market value. More generally, the financial intermediary's common stock will be priced like a bond whose duration equals the average net duration of its assets and liabilities. That is, the nominal contracting hypothesis predicts that cross–sectional variation in the effect of unanticipated interest rate changes on stock prices should be related to differences in balance sheet compositions. A testable implication of the nominal contracting is the maturity mismatch hypothesis, which postulates that differences in the maturity composition of net nominal assets cause differences in the interest rate sensitivity of common stock returns. Actually, the above theory finds its origins in the duration theorem stated independently by Samuelson (1945) and later by

1 This might be less true for industrial firms which hold real assets. However, firms generally have a variety of nominal assets and liabilities. For example, cash, accounts receivable, depreciation tax shields, and contracts to sell products at fixed prices are nominal assets. On the other hand, debt, accounts payable, labour contracts, raw material contracts, and pension commitments may be nominal liabilities.

Hicks (1946): an increase (decrease) in interest rates will increase net worth if the weighted duration of the liability stream is greater than (less than) the weighted duration of the asset stream[2]. The Hicks–Samuelson duration theorem suggests that duration plays a critical role in planning the current structure of the balance sheet. If x_t is the Sterling amount of payment due to at date $t=1,2,...,T.$ and $d=\frac{1}{1+r}$ is the rate of discount defined for the interest rate $r>-1$, then $V(x)=\sum x_t d^t$ is the present value of the stream of payments $\{x_t\}$. The elasticity of *V* with respect to *d* is $D(x)=\frac{\sum t x_t d^t}{\sum x_t d^t}$ i.e. the weighted average date to maturity or duration of the stream $\{x_t\}$. To prove the duration theorem, let $\{\mathrm{A_t}\}$ and $\{\mathrm{L_t}\}$ be asset and liability streams, respectively, whose present values calculated at the common discount rate *d* are *V(A)* and *V(L).* Net worth is then *V=V(A)–V(L)*. If D_A and D_L are the duration of these streams, then differentiating the net worth with respect to the market yield we obtain the following:

$$\frac{\partial V}{\partial r}=d[V(L)D(L)-V(A)D(A)] \tag{1}$$

and the above statement follows at once.

If we assume *t* be a continuous variable on the closed interval [0,T] and R_t be the market rate of interest currently quoted on loans maturing at date *t* in the future, then the investor faces two types of risk in his investment horizon planning process. The first is that his expectation may not be realised. That is, his/her asset and liability streams may be different as debtors can default on their obligations and creditors can revise payments schedules on debts incurred. We shall assume away this kind of risk in the discussion that follows.

The second kind of risk is present when interest rates can change. When *t* is a discrete variable, the structure of yields must be one for which the long rate R_t is the geometric average of current and one period forward short rates (r_t).

$$R_t=\left\{\left(1+r_1\right)+\left(1+r_2\right)+.....+\left(1+r_t\right)\right\}^{1/t}-1 \tag{2}$$

When *t* is a continuous variable and interest is compounded continuously then the relationship is as follows:

$$R_t=\begin{cases} r_t=\text{constant} & t=0 \\ \left[\int_0^t r(x)dx\right]\Big/t & t\succ 0 \end{cases} \tag{3}$$

---

2 If the appropriately weighted durations of these streams are equal, then net worth will be unaffected by small changes in interest rates.

We shall refer to the function R_t as the market yield curve spanning the time interval [0,T]. Thus, the investor's wealth under this assumption is given by the expression:

$$W = \int_0^T \left(A_t - L_t\right) exp\left(-R_t \cdot t\right) dt \tag{4}$$

In the proof follows we shall assume that both L and W are positive so that $A>L>0$. Wealth is the difference between the present values of the investor's asset and liability streams, and changes in interest rates can affect these values differently. To understand why this is happening, assume that in equation (3) $R_t=R$ and differentiate W with respect to R. If the definitions:

$$\begin{aligned} D_A &= \frac{1}{A}\int_0^T t \cdot A_t \, exp\left(-R_t \cdot t\right) dt \\ D_L &= \frac{1}{L}\int_0^T t \cdot L_t \, exp\left(-R_t \cdot t\right) dt \end{aligned} \tag{5}$$

of the duration of these two streams are employed in the resulting expression, the Hicks–Samuelson theorem follows at once: the sign of the derivative dW/dr is the sign of the difference $LD_L–AD_A$ and the importance of the duration of asset and liability streams is apparent.

Several possibilities have been suggested under this framework. If D_A and D_L can be adjusted by the investor so that hedge or immunise wealth against changes in interest rates. On the other hand, an investor can also speculate on the yield movements. That is, if his/her expectations are that interest rates will rise, he/she can gamble on this guess by choosing $LD_L–AD_A>0$ and will be better off if his/her guess turns out to be correct. Similarly, he/she can gamble on a guess that interest rates will fall by choosing $LD_L–AD_A<0$. Nevertheless, the net effect of the aforementioned hypotheses is an empirical issue and beyond of the scope of the present study.

The other important issue is whether the exchange rate factor plays an important role in the insurance industry. As far as its impact is concerned, it is most probable to affect more companies dealing with foreign customers than firms dealing in the local market. This is the corner–stone of our second hypothesis that general insurance firms should be more harmed by any change in the exchange rates, while life insurance companies is expected either to be less affected or not at all.

3. Data and methodology

The sample examined in the present study consists of 21 insurance companies (7 life and 14 general) all listed in the London Stock Exchange. The sample spans the period from the first week of January 1989 through the last week of December 1998.

Weekly data are employed to avoid the skewness of daily returns[3] [Roll & Ross (1980)], and then are used to test the pricing relation and the impact of interest and exchange rates on the common stock returns of the insurance industry. Compounded returns are calculated by taking the prices at the last trading date of each week and the returns on the FTSE All–Share Index, the widest equity market index in the UK, are used as a proxy of the return on the market portfolio. The unexpected changes in one– and three– month Treasury bill rates are later used as the interest rate variable in equation (8). The exchange rate variable is represented by the change in the rate of US dollars per one Sterling pound in the same equation.

The arbitrage pricing theory (APT) postulates that the difference between actual and expected return on an asset *i* is a linear function of *k* economic variables. The actual returns of all assets in the market are governed by their sensitivities to the set of unexpected changes in these *k* factors. That is, the above can be formulated as follows:

$$R_{it} = E(R_i) + \beta_{i1} f_{1t} + \beta_{i2} f_{2t} + ... + \beta_{ik} f_{kt} + \varepsilon_{it} \tag{6}$$

The endogenous variable is the random return on asset *i*, $E(R_i)$ is the expected return on the same asset and the rest are the *k* economic forces. The betas measure the percentage change of assets' returns with respect to the change in the particular factor and ε_i is the white noise error term with the known properties. Ross (1976), in his seminal paper, showed that in equilibrium the expected returns are modelled as:

$$E(R_i) = \gamma_o + \gamma_{i1}\beta_{i1} + \gamma_{i2}\beta_{i2} + ... + \gamma_{ik}\beta_{ik} \tag{7}$$

The gamma coefficients are commonly referred to as the price of risk (premium or discount) and is a measure of association between the sensitivity of each asset *i* to a particular economic factor and the expected return on that asset. Assuming a three index model and substituting (7) into (6) we obtain the following model to examine the sensitivity and pricing questions:

$$R_{it} = \gamma_0 + \gamma_1\beta_{i1} + \gamma_2\beta_{i2} + \gamma_3\beta_{i3} + \beta_{i1}R_{mt} + \beta_{i2}SI_t + \beta_{i3}XR_t + \varepsilon_{it} \tag{8}$$

Where: R_{pit} = weekly return on an insurance stock i in week t.
R_{mt} = weekly return on the market index in week t.
SI_t = the orthogonalised surprises in interest rates.
XR_t = the orthogonolised innovation in US dollars per Sterling pound.
$\gamma_{1,2,3}$ = the prices of risk on the factors employed.

Equation (8) is estimated as a system of all available regressions. Although disturbances within each equation may be independent across observations, an nonzero correlation between corresponding disturbances from different equa-

[3] Fama (1976) and Trzcinka (1986) also argued that daily returns are not well described by the normal distibution and Roll & Ross (1980) showed that APT test were improved when every two other observations were used.

tions may be present. Assume for a particular observation *t*, that $cov(u_{it,} u_{jt}) = \sigma_{ij}$ whilst $cov(u_{it}, u_{jm}) = 0; t \neq m$. Then, this implies that there is no serial correlation and serial cross correlation. However, if $\sigma_{ij} \neq 0; i \neq j$ then there is contemporaneous correlation. The existence of contemporaneous correlation of disturbances explains the use of the term SURE, which achieves an improvement in efficiency by taking into explicit account that correlation (Zellner 1962, Burmeister & McElroy 1988). The process of stacking, assuming *N* equations, is:

$$\begin{bmatrix} Y_1 \\ \vdots \\ Y_N \end{bmatrix} = \begin{bmatrix} X_1 & & \\ & \ddots & \\ & & X_N \end{bmatrix} \times \begin{bmatrix} b_1 \\ \vdots \\ b_N \end{bmatrix} + \begin{bmatrix} u_1 \\ \vdots \\ u_N \end{bmatrix}$$

The same system can be written in a shorthand format as:

$$\mathrm{Y = X\,b + u}$$

where $Y = NT \times 1$ *matrix*

$X = NT \times \sum_{i=1}^{N} k_i$ *matrix*

$b = \sum_{i=1}^{N} k_i \times 1$ *matrix*

$u = NT \times 1$ *matrix*

According to the assumption of SURE model:

$$E(u_i u_j') = \Omega \otimes I_T \quad \text{with} \quad \Omega = \left[\sigma_{ij}\right] i, j = 1,2,\ldots\ldots,N$$

The most efficient estimation of the system is obtained by applying GLS to get:

$$\hat{b} = [X'(\Omega^{-1} \otimes I_T)X]^{-1}\,[X'(\Omega^{-1} \otimes I_T)Y]$$

with

$$E[(\hat{\beta} - \beta)(\hat{\beta} - \beta)'] = \left(\sigma^{ij} X_i' X_j\right)^{-1}\ [4]$$

The objective function is to choose *b* and Ω so as to maximise the log likelihood function

$$\ell(b,\Omega) = -(TN/2)log(2\pi) + (T/2)log\left(\Omega^{-1}\left(-0.5 \sum_{t=1}^{T} (Y_t - X_t b)' \Omega^{-1} (Y_t - X_t b)\right.\right.$$

which implies to choose *b* so as to minimise $Q(b) = u(b)'(\Omega^{-1} \otimes I_T)u(b)$ where *u(b)* is the vector of stacked residuals (a function of the parameters *b*), Ω is the esti-

4 This is equivalent to $Var(b) = [X'(\Omega^{-1} \otimes I_T)X]^{-1}$

mated covariance matrix of the residuals and I is the identity matrix. If Ω is recomputed from b_i at each iteration, this estimator converges to the ML estimator if the residuals are assumed to be multivariate normal. The SURE multivariate regression approach is employed in the present study, with a system of 21 equations, each one representing the ex–post stock return on individual insurance companies. An econometric presentation of the system of equations employed to simultaneously estimate the risk premia and factor sensitivities is presented in appendix 1.

4. Generating unanticipated components

Since the APT postulates that it is the innovation in each factor that affect the returns of each asset, an expectation generating process must be adopted to originate this component. A dynamic system in a particular form called the state space representation is employed. The Kalman filter (Kalman 1961) is an algorithm for sequential updating a linear projection for the system[5]. In particular, is a recursive procedure for calculating the optimal estimator of the state vector given all the information which is currently available.

Although the exposition follows presents the general univariate model, the general kalman filter approach applies to multivariate series as well. The Kalman filter consists of a system of two equations which take the form of:

$$\underset{(N\times1)}{Y_t} = \underset{(N\times m)}{X_t}\ \underset{(m\times1)}{\beta_t} + \underset{(N\times1)}{e_t} \tag{9}$$

$$\underset{(m\times1)}{\beta_t} = \underset{(m\times m)}{d}\ \underset{(m\times1)}{\beta_{t-1}} + \underset{(m\times1)}{v_t} \tag{10}$$

The errors from the above equations are serially uncorrelated with the following properties:

$$E(e_t) = 0\,Var(e_t) = H \quad 6$$
$$E(v_t) = 0\,Var(v_t) = Q$$

The above system is known as the observation and the state equation referring to equation (9) and (10) respectively. Equation (10) actually shows how the systems updates the information with respect to the coefficients:

$$\beta_t = d\beta_{t-1} + v_t = d(d\beta_{t-2} + v_{t-1}) + v_t = ... = d^t\beta + \sum_{i=1}^{t} d^{t-i} v_i \tag{11}$$

The following two assumptions complete the theoretical specification of the state space system:

---

5 The Kalman filter can be applied to a broad range of models. In our case the state variables are the regression coefficients from the expectations generating model.

6 If $Var(v_t) = 0$, then there is no time variation. The most common way to estimate the Kalman filter is to set Q = 0, H is assumed to be constant σ^2 while the β_0 and P_0 are obtained using estimates through part of the sample.

1. The initial state vector β_0 has the following properties:

$$E(\beta_0) = 0 \, Var(\beta_0) = P_0$$

2. The disturbance of both equation are uncorrelated with each other in all time periods, and uncorrelated with the initial state:

$$E(e_t v_\tau') = 0 \, \forall t, \quad \tau E(e_t \beta_0') = 0 \quad E(v_t \beta_0') = 0$$

Combining the intuition behind equation (9), (10) and (11) the Kalman filter model reduces to:

$$Y_t = (d^t \beta)' X_t + \sum_{i=1}^{t} (d^{t-i} v_i)' X_t + e_t$$

Apart from the aforementioned method, other econometric expectation generating processes have also been applied to test the robustness of the results[7].

Note, however, that the estimation of the three–index model presents a problem because the two regressors exhibit collinearity with the market index. The examination of the inter–dependence between the exogenous variables reveals a correlation coefficient with the market portfolio of –0.31 and –0.22 for the three–month Treasury bill and the exchange rate respectively. Testing the hypothesis that the correlation is zero, the *t* statistic is found to be 3.02 and 2.29 respectively. The correlation is significant at a high level and to overcome this problem an orhogonalisation method is adopted[8]. That is, the residuals from the auxiliary regressions of interest and exchange rates on the market portfolio replace the two macroeconomic variables in equation (8).

5. Estimation results

The multivariate estimation for the interest and exchange rate sensitivity of the life and general insurance companies are presented in table 1 and table 2 respectively. To conserve space and due to the similarity of the regression results, by using either the one or three month Treasury bill rate, only the results of the three month Treasury bill are reported.

Table 1: *Multivariate Estimation of Life Insurance Companies*

	Beta of Market Portfolio	**Beta of Interest Rate**	**Beta of Exchange Rate**
Britannic Assurance	0.481	–0.939	3.301
	(6.02)	(–1.99)	(1.89)*
London & Manc. Group	0.515	–1.831	1.989
	(7.40)	(–2.02)	(0.75)

7 See Dinenis & Staikouras (1996).

8 See Dinenis & Staikouras (1998).

Legal & General	1.012	–1.011	5.010
	(5.88)	(–2.01)	(1.12)
Prudential Corporation	1.160	–0.636	–0.511
	(5.99)	(–1.96)	(–0.69)
Refuge Group	0.502	–2.345	5.261
	(6.10)	(–2.45)	(1.21)
Lloyds Abbey	0.989	–2.033	–3.211
	(6.81)	(–2.01)	(–0.99)
UTD Friendly	0.499	–1.369	–5.971
	(5.55)	(–1.93)	(–1.11)
AVERAGE	0.737	–1.452	0.838

t – values in parenthesis.

* Significant coefficient for the exchange rate factor.

It is clear from both tables that the returns of all the insurance firms are negatively related to the unanticipated changes in interest rates. The regression coefficients of the life and general insurance firms for the particular interest rate index indicate that on average, a 2% points increase (decrease) in interest rates would cause a 2.904% and 1.096% points decrease (increase) respectively, in the common stock returns of these insurance firms. The life insurance companies seem to be much more sensitive to interest rate shocks than the general insurance group. On the contrary to the previous discussion, the exchange rate variable seems to discriminate between the two kinds of insurance companies.

Table 2: *Multivariate Estimation of General Insurance Companies*

	Beta of Market Portfolio	**Beta of Interest Rate**	**Beta of Exchange Rate**
General Accident	1.011	–0.211	1.249 °
	(10.06)	(–1.95)	(1.98)
Heath	0.555	–0.654	–4.371
	(7.01)	(–1.99)	(–2.15)
Royal Insurance Holdings	1.262	–0.321	–8.546
	(9.11)	(–1.98)	(–2.17)
Sedgwick Group	0.712	–0.989	–4.314
	(6.31)	(–1.93)	(–2.31)
Sun Alliance Group	1.066	–0.861	0.741 °
	(9.01)	(–2.01)	(0.98)
Trade Indemnity	0.333	–0.256	–6.491
	(1.99)	(–2.03)	(–1.99)
Willis Corroon	0.569	–1.123	–5.128
	(4.88)	(–2.21)	(–2.03)
Winsdor	0.313	–0.765	–2.351
	(2.01)	(–2.15)	(–2.18)
Archer Group	0.398	–0.450	–3.129 °
	(2.69)	(–1.97)	(–0.97)
Bradstock Group	0.390	–0.667	–6.160
	(3.69)	(–1.90)	(–1.96)
Domestic & General	0.101	–0.471	–5.137

	(2.32)	(–1.96)	(–2.09)
Lloyd Thompson	0.364	–0.301	–4.245
	(3.12)	(–2.22)	(–1.97)
PWS Holdings	0.222	–0.212	3.012 °
	(1.98)	(–2.00)	(1.39)
Steel Burrill	0.399	–0.399	–3.690
	(2.99)	(–2.10)	(–2.02)
AVERAGE	0.549	–0.548	–3.574

t – values in parenthesis.

° Insignificant coefficients or "wrong" sign for the exchange rate factor.

By looking the third and forth column in table 1 and table 2 our earlier presumption is verified. Since general insurance companies invest a lot in foreign currency, while life insurance firms restrict themselves to the local market, their exchange rate beta coefficients are particularly higher and significant at a high level of confidence. Eighty percent of the sample available, and for the period under consideration, exhibit significant coefficients for the exchange rate variable. Only three companies show insignificant coefficients for the exchange rate factor and one is significant but with the "wrong" sign (positive). On the other hand, life insurance companies are not affected by the exchange rate variable, since six out of seven companies exhibit coefficients insignificantly different from zero. Although Britannic Assurance exhibit significant coefficient, this has a positive sign and so no strong inferences can be made about its sensitivity to this particular factor.

Table 3: *Risk Premia Estimates from the Multivariate Regression*

		Average insurance exposure⋆		**Contribution to insurance expected returns (%)**	
	Price of risk	***Life***	***General***	***Life***	***General***
$\hat{\gamma}_0$	1.987	–	–	1.987	1.987
	(1.91)				
$\hat{\gamma}_1$	1.189	0.737	0.549	0.876	0.653
	(1.99)				
$\hat{\gamma}_2$	–1.349	–1.452	–0.548	1.959	0.739
	(–1.97)				
$\hat{\gamma}_3$	–0.831	0.838*	–3.574	–	2.970
	(–1.98)				
		Average expected return		4.822	6.349

t – values in parenthesis.

⋆ Hedge ratios for every risk factor (i.e. market, interest rate and exchange rate).

* Insignificant value.

The advantage of using a system of equations (multivariate approach) is that you can obtain estimates of the assets' sensitivities to the factors employed and simultaneously estimate whether the market prices each of these risk factors. The prices per unit of sensitivity are called the price of risk. The fact that a group of companies exhibit a significant beta coefficient to a particular factor this does not necessarily indicate a pricing relationship. When the market prices a factor that means that the coefficient on the beta of the particular factor is statistically different from zero. If this is not the case then the market can be said to be inefficient and opportunities of arbitrage exist. The estimated prices of risk associated with each of our three factors are presented above in table 3. Both exchange and interest rate factors seem to be priced by the market since they exhibit significant coefficients. The negative signs for the price of risk seem to be consistent with the view that these institutions can be considered as hedges against other assets that are more fixed in nominal terms.

Finally, using the traditional two step methodology one would be able to estimate the ex–post sample risk premia plus the difference between the sample and population means. The fact that the premium on the market is not significant is not that worrying, since the estimated premia for the market using multivariate regressions cannot be estimated as the expected return on the market drops out from thegenerating function[9].

5. Conclusion

In the present study the risk–return relationship for individual securities from the UK insurance sector is investigated. Using the SURE nonlinear multivariate method, we employ a three factor return generating process to quantify the concept of the systematic exchange and interest rate risk. Listed general and life insurance companies were used to check the relationship between the surprises in these factors and their equity returns. Apart from the academic importance of such research, there is a number of practical applications concerning fund and money managers, such as portfolio selection, performance measurement, asset pricing, investing and financing decisions, and asset allocation.

The empirical results show a significant negative relationship between the shocks in interest rate variable and the sample's stock returns. That is, when there is an unanticipated increase in interest rates portfolios consisting of these financial firms exhibits a decrease in their returns probably attributed to the duration gap between their assets and liabilities, or in other words by the intermediary's risk exposure. However, consistent with our original priors, the life insurance group is not affected by the changes in exchange rates. Exchange rate shocks seem to have a significant negative effect on general insurance firms, mainly through their investments in foreign currency.

The upshot of our research can be summarized as follows. First, both life and general insurance companies are equally affected by the interest rate variable employed. Second, the exchange rate risk factor seem to predominate the generating function of returns of only the general insurance firms leaving the life in-

[9] A pertinent discussion can be found in Sweeney & Warga (1986b).

surance sector unaffected, at least for the period under consideration. Third, the risk premia for both interest and exchange rates are priced by the market being consistent with the theory of arbitrage pricing.

References

Bae S.C., 1990, “Interest rate changes and common stock returns of financial institutions: Revisited” *Journal of Financial Research* 13, 71–79.

Burmeister E. – M.B.McElroy, 1988, “Joint estimation of factor sensitivities and risk premia for the arbitrage pricing theory” *Journal of Finance* 43, 721–735.

Dinenis E. – S.K.Staikouras, 1996, “Does the UK equity market compensate investors for bearing interest rate risk?” Paper presented at the V Tor Vergata Financial Conference, Financial Markets: Imprfect Information and Risk Management, Rome, Italy.

Dinenis E. – S.K.Staikouras, 1998, “Interest rate changes and common stock returns of financial institutions: Evidence from the UK” *European Journal of Finance* 4, pp.113–127.

Fama E.F. “Foundations of finance” NY 1976.

Flannery M.J. – C.M.James, 1984, “Market evidence on the effective maturity of bank assets and liabilities” *Journal of Money, Credit and Banking* 16, 435–445.

French K.R. – R.S.Ruback – G.W.Schwert, 1983, “Effects of nominal contracting on stock returns” *Journal of Political Economy* 91, 70–96.

Hicks J.R., 1946, “Value and capital” Oxford: Clarendon Press.

Kalman R.E., 1961, “A new approach to linear filtering and prediction problems” *Journal of Basic Engineering* 83D, 95–108.

Roll R. – S.A.Ross, 1980, “An empirical investigation of the arbitrage pricing theory” *Journal of Finance* 35, 1073–1103.

Ross S.A, 1976, “The arbitrage theory of capital asset pricing” *Journal of Economic Theory* 13, 341–360.

Samuelson P.A. “*The effect of interest rate increases on the banking system*” *The American Economic Review* 35, pp.16–27.

Saunders A. – P.Yourougou, 1990, “Are banks special? The seperation of banking from commerce and interest rate risk” *Journal of Economics and Business* 42, 171–182.

Sweeney R.J. – A.D.Warga, 1986b, “The possibility of estimating risk premia in asset pricing models” *The Financial Review* 21, 299–308.

Theil H. “Principles of econometrics” Wiley 1971.

Trzcinka C., 1986, “On the number of factors in the arbitrage pricing model” *Journal of Finance* 41, 347–368.

Zellner A., 1962, “An efficient method of estimating seemingly unrelated regressions and test for aggregation bias” *Journal of the American Statistical Association* 57, 348–368.

APPENDIX 1

Estimating Factor Sensitivities and Risk Premia using the Nonlinear SURE Approach.

The nonlinear multivariate estimating equation, incorporating the restrictions as mandated by the APT is:

$$R_{it} = \gamma_0 + \gamma_1\beta_{i1} + \gamma_2\beta_{i2} + \gamma_3\beta_{i3} + \beta_{i1}R_{mt} + \beta_{i2}SI_t + \beta_{i3}XR_t + \varepsilon_{it}\, i = 1,2,\ldots\ldots,N. \quad (12)$$

The above equation can be viewed as either a system of N regressions by restricting the $\gamma's$ to be the same across equations, or simply can be presented into a single regression of the following form:

$$\begin{bmatrix} R_1 \\ R_2 \\ \vdots \\ \vdots \\ R_N \end{bmatrix} = \begin{bmatrix} (i_T : R_m) & 0 & \ldots & 0 \\ 0 & (i_T : R_m) & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & (i_T : R_m) \end{bmatrix} \times \begin{bmatrix} \gamma_0 + \gamma_1\beta_{11} \\ \beta_{11} \\ \gamma_0 + \gamma_1\beta_{12} \\ \beta_{12} \\ \vdots \\ \gamma_0 + \gamma_1\beta_{1N} \\ \beta_{1N} \end{bmatrix} +$$

$$+ \begin{bmatrix} (i_T : SI) & 0 & \ldots & 0 \\ 0 & (i_T : SI) & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & (i_T : SI) \end{bmatrix} \times \begin{bmatrix} \gamma_2\beta_{21} \\ \beta_{21} \\ \gamma_2\beta_{22} \\ \beta_{22} \\ \vdots \\ \gamma_2\beta_{2N} \\ \beta_{2N} \end{bmatrix} +$$

$$+ \begin{bmatrix} (i_T : XR) & 0 & \ldots & 0 \\ 0 & (i_T : XR) & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & (i_T : XR) \end{bmatrix} \times \begin{bmatrix} \gamma_3\beta_{31} \\ \beta_{31} \\ \gamma_3\beta_{32} \\ \beta_{32} \\ \vdots \\ \gamma_3\beta_{3N} \\ \beta_{3N} \end{bmatrix} + \begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \vdots \\ \vdots \\ \varepsilon_N \end{bmatrix}$$

The previous system can be written more compactly [10] as follows:

$$\underset{TN\times 1}{R} = \left[\underset{TN\times N}{I_N \otimes (i_T : R_m)} \right] \left[\gamma_0 + \gamma_1 \underset{N\times 1}{\beta_m} : \underset{N\times 1}{\beta_m} \right] + \left[\underset{TN\times N}{I_N \otimes (i_T : SI)} \right] (\gamma_2 \underset{N\times 1}{\beta_I} : \underset{N\times 1}{\beta_I}) + \left[\underset{TN\times N}{I_N \otimes (i_T : XR)} \right] (\gamma_3 \underset{N\times 1}{\beta_{XR}} : \underset{N\times 1}{\beta_{XR}}) + \underset{TN\times 1}{\mathrm{E}} \quad (13)$$

The above equation in order to look more familiar with equation (8) can be written taking the following form:

$$\underset{TN\times 1}{R} = \underset{TN\times N}{(I_N \otimes i_T)} \left[\gamma_0 + \gamma_1 \underset{N\times 1}{\beta_m} + \gamma_2 \underset{N\times 1}{\beta_I} + \gamma_{32} \underset{N\times 1}{\beta_{XR}} \right] + \underset{TN\times N}{(I_N \otimes R_m)} \underset{N\times 1}{\beta_m} + \underset{TN\times N}{(I_N \otimes SI)} \underset{N\times 1}{\beta_I} + \underset{TN\times N}{(I_N \otimes XR)} \underset{N\times 1}{\beta_{XR}} + \underset{TN\times 1}{\mathrm{E}} \quad (14)$$

where: $R' = (R_{i1}, R_{i2}, \ldots\ldots, R_{iT})\, 1\times TN$ *vector.*

$i_T = T \times 1$ *unit vector.*

$I_N = N \times N$ *identity matrix.*

$R'_m = (R_{m1}, R_{m2}, \ldots\ldots, R_{mT})\, 1\times T$ *vector.*

$SI' = (SI_1, SI_2, \ldots\ldots, SI_T)\, 1\times T$ *vector.*

$XR^{©} = (XR_1, XR_2, \ldots\ldots, XR_T)\, 1\times T$ *vector.*

$\mathrm{E}' = (\varepsilon_{i1}, \varepsilon_{i2}, \ldots\ldots, \varepsilon_{iT})\, 1\times TN$ *vector.*

$\mathrm{E}\,´\, MVN(0, \sigma_{ij} \otimes I_T)\, i = 1,2,\ldots\ldots, N.$

The variance–covariance matrix of E is a block diagonal matrix with σ_{ij} appearing along the diagonal. This is an implication based on the assumption of serially independent but contemporaneously correlated returns.

[10] The symbol '$\otimes$' indicates the Kronecker or direct product operator of two matrices [Theil (1971), pp.303–306].