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INTERNATIONAL CAPM: WHY HAS IT FAILED?

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ABSTRACT

International CAPM: Why Has It Failed?*

Previous empirical studies of international CAPM models have not found much supporting evidence. In this paper we suggest reasons why this might have happened and perform new tests using improved models and data.

A range of monthly CAPM models are estimated for 1973-87 for aggregate equities and bonds in Germany, Japan, the United States and United Kingdom. The models are an improvement on earlier work in that we integrate equity markets into the analysis instead of focusing exclusively on government bond stocks, and we carefully measure the rates of return for both bonds and equities. In particular, bond returns reflect changes in the price of bonds as well as coupons. Despite this wider portfolio and the introduction of ARCH effects in the conditional covariance matrix of errors, our model still yields unlikely estimates of the coefficient of relative risk aversion and provides very little explanatory power for expected relative rates of return. Correcting the ICAPM for these major deficiencies does not reverse earlier conclusions in the literature.

JEL classification: 430
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NON-TECHNICAL SUMMARY

Risky assets offer a higher expected rate of return than the safe (or riskless) asset, otherwise risk-averse investors would not hold risky assets. The difference between this expected return and the return on the safe asset is called the risk premium. In our paper we look critically at the literature which seeks to explain the *foreign exchange* risk premium within a simple economic framework known as the International Capital Asset Pricing Model (ICAPM). This risk premium is intended to refer to the expected reward for holding assets denominated in one currency relative to the return on the safe asset. The ICAPM approach relates the expected relative return over, say, a one-month period on a given asset to the stocks of assets at the beginning of that month in the investor's portfolio. Typically researchers have considered the outstanding bond stocks of the major industrialized countries as the constituents of their investment portfolios, including Germany, Japan, Italy, France, Canada, Switzerland, the United States and United Kingdom. In other words, they have omitted equities and other financial and physical assets. Also they have used Eurodeposit interest rates to reflect the return on the various government bonds.

The major empirical conclusion of these studies is that they do not provide a good explanation of foreign exchange risk premia. We suggest that researchers should make a number of simple empirical improvements before rejecting completely these ICAPM models:

(i) A more complete portfolio should be specified, including equities, and focusing on the major markets alone (say United Kingdom, United States, Japan and Germany), to make difficult estimation problems more tractable.

(ii) One should exercise greater care when constructing rates of return for the various assets considered; e.g. if government bonds are the assets considered for investment purposes then their rates of return should reflect bond price changes and coupons as well as exchange rate changes. Researchers have thus far used Eurodeposit rates adjusted for exchange rate changes, which are much less variable than those which include bond price changes.

(iii) As part of a progressive research strategy one should check for signs of statistical misspecification in individual equations as a guide for further improving our models. A basic statistical problem with this empirical area is that we are trying to explain highly variable relative returns by stocks of financial assets which change little. Researchers have introduced time-varying variances and covariances (ARCH) into models in an attempt to rectify this, but with little success. We suggest from misspecification analysis that such added sophistication may be inappropriate.

(iv) We should consider what we mean by the risk premium in the context of the representative international investor considering an international portfolio and consuming a 'representative' basket of international goods. Is there such a thing as a safe asset in this context? If there is no safe asset, what is it we are trying to explain? In this paper we refer to this as the *relative* risk premium, in so far as it represents the expected return for holding one risky asset relative to another.

In our paper we estimate a series of monthly ICAPM models for government bonds and equities from 1973-87 for Germany, Japan, UK and US using a new, carefully constructed data set with special emphasis on calculating appropriate bond returns. We still find little explanatory power for relative risk premia in our models, even when we introduce ARCH processes in the equation errors: we also find that CAPM models for different assets and within separate countries yield conflicting inferences. The estimates of the coefficient of relative risk aversion are also highly variable and frequently ill-determined, as in earlier studies. Overall we would have to conclude that risk neutrality seems to be a better description of the data in this context.

We conclude that it is unlikely that the variation in risk premia can be explained by the ICAPM approach.

Introduction

To say that estimates of the international CAPM (ICAPM) have been disastrous for the model would only be a slight exaggeration. Introduced initially as a way of filling the gap in the fundamentals' explanation of the exchange rate left by the failure of models based on uncovered interest parity (UIP), the CAPM has not translated to the international context in the way hoped. However, just as most of the empirical work on models of the exchange rate are fatally flawed[1], that on the ICAPM[2], whilst of higher quality, is sufficiently negligent in some respects to leave reasonable doubt about whether rejecting the model at this stage would be premature. The purpose of this paper is to make a critical re-examination of the empirical content of the ICAPM and to provide some new estimates based on more appropriate data.

One criticism of previous studies is the data used. The problems in obtaining appropriate date are enormous and many are insurmountable. Nevertheless, we believe that we have constructed a data set which incorporates a number of improvements. A second criticism is the choice of assets and countries included in the ICAPM. In his work Frankel (1982) has six countries (France, Italy, Japan, UK, US and West Germany) but only one asset (government bonds) for which he uses euro-rates as the rate of return; Engel and Rodrigues (1989) and Giovannini and Jorion (1989) use the same data. As we show later, there is little or no

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[1]See Backus (1984) and Smith and Wickens (1986, 1989).

[2]See Engel and Rodrigues (1989), Frankel (1982, 1986) and Giovannini and Jorion (1989).

correlation between bond and euro-deposit returns. In an attempt to rectify the omission of a large share of the world's financial wealth, Giovannini and Jorion (1989) include US equities together with government bonds for four countries, but exclude Japan, which has one of the world's largest stock markets, and include Switzerland, which has a tiny share of the world's bonds. Also they use euro-rates instead of proper bond prices.

Roll (1977) observed that an unambiguous test of the CAPM is not possible unless researchers use the true market portfolio. However, the addition of bonds, real estate and consumer durables to a portfolio of stocks or the addition of prior information does not affect inferences in the US context (Stambaugh (1982), Shanken (1986, 1987), Kandel and Stambaugh (1987), though Kandel (1984) suggests that mean variance efficiency is not really testable on a subset of assets even if bounds can be placed on the market share and expected return of the missing asset. No such empirical results have been noted in an international setting for CAPM (though see Cho, Eun, and Senbet (1986) for the international APT). Nevertheless, our ICAPM results, even when based on a much wider market portfolio than used by Frankel (1982), Engel and Rodrigues (1989), Giovannini and Jorion (1989), do not contradict their main conclusions.

In his initial, and pioneering, work on ICAPM Frankel used a static model. Later[3] he admitted this restricted the ability of the CAPM to explain the well-documented volatility of excess returns. One way of generalising the static CAPM is to assume that the conditional covariance matrix of returns is generated by an ARCH in mean model[4]. This is the

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[3]Frankel and Meese (1988).

[4]See Engle and Bollerslev (1987), Engle, Lilien and Robbins (1987) and Bollerslev and Engle and Wooldridge (1988).

approach adopted by Engel and Rodrigues, and Giovannini and Jorion. There is a similarity in the results obtained; the estimates of the coefficient of relative risk aversion (CRRA) are frequently large, negative and insignificant, the ARCH effects are not pronounced and, as we suggest below, could be spurious, while the overall explanatory power of the model is very low.

An insignificant CRRA implies that there is no well determined risk premium in these models of ICAPM and hence, apparently, no risk aversion. This result may be compared with those of Wickens and Thomas (1989) who derive non-parametric estimates of the foreign exchange risk premium and of the equity risk premium. It is found that although the average risk premium is quite small over the period 1973-1988 as a whole, in some years it can be substantial and can explain a large part of the variation in the excess return. Also, they find that the estimate of the CRRA for the whole period is close to risk neutrality, which is consistent with the findings on the average risk premium. These results suggest that although the attempt to allow the ICAPM to exhibit more short-run variability has not succeeded so far, this is the most promising research direction to take.

The present paper is set out as follows. In section 2 we present the mean-variance optimisation theory, while in section 3 we discuss the data in detail. Section 4 contains our empirical results.

2. ICAPM: Mean-Variance Optimisation for Multicurrency Assets

In keeping with earlier studies (e.g. Kouri (1977), Dornbusch (1980), Frankel (1982)), we assume that the representative investor, consuming a representative basket of goods, maximises a function of the mean and variance of end-of-period real wealth, given information available at the beginning of the period.

$$(1) \qquad \text{Max } U[E_t((W_{t-1}), V_t(W_{t+1})]$$

where E_t is the conditional expectation of end-of-period real wealth, W_{t+1}, and V_t is, similarly, the conditional variance. The investor must decide how to allocate initial real wealth among the vector of assets available in the different currencies. As explained above, with the exception of Giovannini and Jorion who include US equities, researchers have followed Frankel (1982) and only considered the outstanding stock of government bonds as wealth. End-of-period real wealth is defined as:

$$(2) \qquad W_{t+1} = W_t + W_t \underset{\sim}{x}_t{}' \underset{\sim}{r}_{t+1} + W_t(1-\underset{\sim}{x}_t{}'\underset{\sim}{\iota})r^n_{t+1}$$

Here $\underset{\sim}{x}_t$ is a vector of shares of the portfolio allocated to (n-1) assets, and $\underset{\sim}{r}_{t+1}$ is the vector of real rates of return realised for the (n-1) assets. The n-th asset has portfolio share $(1 - \underset{\sim}{x}_t{}'\iota)$ and rate of return r^n_{t+1} where ι is an (n-1) vector of ones. This is usually taken as 'U.S. bonds' and interpreted as the 'safe asset', though at this stage we only require that it is an asset in the currency chosen to value our basket of consumption goods. Nevertheless, for reasons of familiarity, we shall use U.S. bonds too. Researchers, again following Frankel (1982), have typically chosen euro-deposit rates to reflect bond returns, or perhaps alternatively have considered euro-deposit rates as appropriate returns to isolate the foreign exchange risk premium. It should be noted, firstly, that euro-deposit rates are inappropriate as measures of the return on government bonds, and, secondly, that eurodollar deposit rates can attract a risk premium over US Treasury securities, e.g. during the Bank Herstatt crisis of 1974 the differential was three full percentage points. Euro-dollar deposits are <u>not</u> risk free.

The first -order condition for the maximum value of (1) subject to (2) is given by:

(3) $$\frac{dU}{d\underset{\sim}{x}_t} = U_1 \frac{d\, E_t W_{t+1}}{d\underset{\sim}{x}_t} + U_2 \frac{dV_t W_{t+1}}{d\underset{\sim}{x}_t} = 0$$

where $U_1 = \dfrac{dU}{dE_t\, W_{t+1}}$ and $U_2 = \dfrac{dU}{dV_t W_{t+1}}$

If we write the return of the (n-1) assets relative to the numeraire asset as:

(4) $$\underset{\sim}{z}_{t+1} = \underset{\sim}{r}_{t+1} - r^n_{t+1}\,\underset{\sim}{\iota}$$

then the first-order condition reduces to

(5) $$U_1 W_t [E_t z_{t+1}] + U_2 W_t^2 [2\Omega_t \underset{\sim}{x}_t] = 0$$

where $\Omega_t \equiv [E_t z_{t+1} - z_{t+1}]' \, [E_t z_{t+1} - z_{t+1}]$

Solving for excess returns gives:

(6) $$E_t\, z_{t+1} = \rho_t \Omega_t \underset{\sim}{x}_t$$

where $\rho_t = -W_t\, 2\,\dfrac{U_2}{U_1}$, is the coefficient of relative risk aversion (CRRA). If we assume that expectations are formed rationally, then

(7) $$\underset{\sim}{z}_{t+1} = \rho_t \Omega_t \underset{\sim}{x}_t + \underset{\sim}{\varepsilon}_{t+1}$$

where $$\underset{\sim}{\varepsilon}_{t+1} = z_{t+1} - E_t z_{t+1}$$

with $E_t\, \underset{\sim}{\varepsilon}_{t+1} = 0$.

From the definition of Ω_t and $\underset{\sim}{\varepsilon}_{t+1}$ we see that there is a powerful restriction on (7): that the variance-covariance matrix of $\underset{\sim}{\varepsilon}_{t+1}$ is none other than Ω_t. This is the key observation of Frankel (1982). The imposition of this particular restriction is the basis of tests of the ICAPM in the succeeding literature.

In Frankel (1982), and Frankel and Froot (1986) it is assumed that ρ_t and Ω_t are constant over time. (7) is then called the static ICAPM. More generally, they can vary over time. Although ρ_t is usually assumed constant, a number of studies have used the ARCH-M model of Engle et al

(1987) to formulate a time series process to describe the way Ω_t evolves. See, for example, Bollerslev, Engle, and Wooldridge (1988) for US domestic financial assets and Engel and Rodrigues (1989) and Giovannini and Jorion (1989) for the ICAPM.

In applying the theory to international financial investment it may be noted that to the international investor risk arises from three sources: uncertainty in

(i) changes in the price (domestic currency) of the asset

(ii) changes in the exchange rate between the currency in which consumption will take place and the currency of denomination of the asset

(iii) changes in goods' prices.

In common with other authors using mean-variance models, we ignore (iii). This is taken into account in the Lucas (1978) arbitrage pricing model, (see also Hodrick and Srivastava (1986) and Backus and Gregory (1989)). However, for short time horizons of around one month this is clearly quite reasonable. The international asset pricing literature follows Frankel (1982) in focussing on (ii). Hence the claim to identify the 'exchange risk premia'. However, as we show below, in a truly international portfolio of equities and bonds, before one even considers translating for exchange rate effects, account has to be taken of the fact that asset prices measured in domestic currency are also highly volatile, and hence a truly general formulation of return and risk should refer to (i) as well. This would be particularly important for Frankel's example where wealth is defined to include only 'outside' assets, namely the (issued) value of government bonds. If the international investor was choosing between fixed nominal (say, euro-) deposits then indeed (i) would be irrelevant and the only stochastic component would be due to (ii). Early in this literature Frankel (1982) suggested a fuller set of assets to better reflect international portfolio choice but until Giovannini and Jorion (1989) this drew little attention. Indeed, even these authors conclude that no conclusive test of the ICAPM is possible without a much fuller asset set. However, they refer to the portfolio as involving the 'foreign exchange market' and US equities, whereas the former is actually foreign currency denominated bonds, which, as we emphasise below, involve quite different rates of return. Unfortunately, given this tradition in the empirical work of using government debt for wealth and euro-rates in the derivation of excess returns, the introduction of equities aggravates the problem still further.

In Wickens and Thomas (1989) non-parametric estimates of the pure equity and foreign exchange risk premia are obtained. The results show that in some years there are substantial risk premia of both types. Over the whole sample period, however, neither risk premium was found to be large and correspondingly, estimates of the CRRA were found to support risk neutrality and not risk-averse behaviour. These results indicate that there is sizeable time variation in the risk premia to be explained and provides prima facie evidence in support of both (i) and (ii).

3. Variable definitions and their measurement

Empirical work on the ICAPM suffers from using inappropriate variable definitions and sources perhaps because attention has focussed instead on the specification and estimation of the model. We wish to highlight those key data issues which must be addressed before such models can be tested sensibly:

(i) given that only a few countries can be included, which should they be?

(ii) what is the appropriate choice of assets and how should they be measured?

(iii) how can rates of return be measured to properly reflect holding period returns on bonds and equities?

3.1 Country coverage

We would prefer to include those countries which contribute most to world wealth. Thus we believe it is essential to include Germany, Japan and the U.S. in any ICAPM not only to reflect their importance for total wealth but also because their currencies occupy a pre-eminent role in the international exchanges. As remarked earlier, it seems inappropriate to include, say, Swiss government bonds, but to exclude both Japanese bonds and equities, (as is done by Giovannini and Jorion (1989)). In the Salomon Brothers World Government Bond Index for 1986, Switzerland has a weight of only 0.4%, while Japan has a weight of 26.4%. In the dataset used by Giovannini and Jorion (1989) Swiss bonds account for only 1.1% of the wealth on average. We also include the U.K. for our empirical work.

3.2. Asset stocks

Examination of the time-path of the shares of bonds and equities in total financial wealth from 1973(6) to 1987(12) (not reproduced here) shows clearly that Japan's equity market becomes of great importance after 1985 (with all assets expressed in $'s), while both the US bond and equity markets fall substantially in importance. Indeed, the Japanese government bond market is also very large. By the end of 1987 Japanese equity accounts for around 26% of total wealth in our sample, and its bonds account for 18%. The US bond share has fallen from 30% in 1984 to around 17% by end -1987, and its equity market share falls from 47% in 1974 to under 23% by the end of our period. These figures reflect both asset price changes in domestic currency and exchange rate changes. It is clear, therefore, that Japanese equity wealth cannot be ignored. In contrast the share of UK and German equities never rise above 8.5% and 4% of total wealth respectively, and their government bond wealth is only slightly higher at 9% and 7.3%. Table 1 contains summary descriptive statistics of country shares by the two asset classes. The US and Japan clearly dominate both equity and bond wealth.[5]

Asset pricing models refer to the allocation of existing wealth measured at current market prices, not of wealth measured at a fixed issue value. Thus Frankel's (1982) careful methodology for calculating outstanding government bond stocks in the hands of the private sector is incorrect since it omits any reference to revaluation effects due to interest rate changes. A measure, like Frankel's, which cumulates historical government borrowing needs or adds the issue value of bonds will have much less variability than one which takes into account day-to-day capital revaluations due to interest rate changes. The familiar complaint in this literature[6] that there is not enough variability in asset shares (referring to government bonds) to explain volatility of excess returns may not be correct if such revaluations were taken into account. Indeed one may conjecture that the combined variability of the asset shares and the conditional covariance matrix of

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[5]The equity market wealth is taken from Datastream. Alberto Giovannini kindly provided bond market data to end-1984 and we constructed the data to end-1987.

[6]Giovannini and Jorion suggest that, under certain not implausible assumptions, the fluctuation in asset supplies can only predict a standard error of the D.M. risk premium that is 1/200 of its standard error obtained from unrestricted projection equations.

returns - $\hat{u}_t$ in (7) - will be roughly of the same order of magnitude as the variability of the returns themselves since interest changes will be important for each. In adding U.S. stock market wealth to the portfolio, Giovannini and Jorion (1989) are presumably using updated share values for this asset but issue values for their bond assets. In addition, they interpolate their monthly bond stocks to obtain weekly data: this would presumably lead to even smoother asset shares. No investor would seriously consider using issue or book prices to value current equity wealth. Unfortunately, a time series of revalued government bond stock data are not available to our knowledge. However, we do know from Salomon Brothers (1986) that principal values and market values can diverge considerably. For the sample of bonds which enter their World Government Bond Index, the principal and market values were, (1st October 1986): US, \$841.0 bn and \$971.3 bn; Japan, \$473.3 bn and \$516.8 bn; UK \$153.1 bn and \$153.6 bn; West Germany, \$57.3 bn and \$87.1 bn. Engel and Rodrigues (1989) consider measurement error in the relative rates of return due to preference shocks; including such effects does not greatly improve their estimates and does not seem as intuitively plausible as allowing for measurement error in the asset shares.

3.3. Rates of Return

Basic finance tells us how to measure ex post <u>nominal</u> rates of return on bonds and equities. The ex-post nominal return on equities (holding period return) is

$$\frac{D_t + S_{t+1} - S_t}{S_t} \tag{8}$$

where D_t is the dividend at the beginning of the period (assumed known here) and S_t is the share price at t

The ex-post nominal return on bonds is:

$$\frac{C_t + B_{t+1} - B_t}{B_t} \tag{9}$$

where C_t = coupon at time t and B_t = bond price at time t

Using Frankel's (1982, p.257) notation, ex-post real rates of return on the j-th asset are obtained as follows:

$$1 + r_{t+1}^{j} \equiv \frac{1 + i_t^{j}}{(1+\pi_{t+1}^{\$})(1+\Delta e_{t+1}^{j})} \simeq 1 + i_t^{j} - \pi_{t+1}^{\$} - \Delta e_{t+1}^{j} \tag{10}$$

where i_t^j is the one-period interest rate and $\pi_{t+1}^{\$}$ is the rate of inflation during the period, converted into a dollar price index for a common consumption bundle. Δe_{t+1}^{j} is the rate of depreciation of currency j against the dollar. Equation (10) is a sensible measure of the real return from holding, say, a one-month bank deposit in currency j: the only change in 'capital' value arises from changes in e_t, the spot exchange rate. This is the rate of return employed in tests of international CAPM. The ex-post real rate of return on dollar assets is given by:

$$1 + r_{t+1}^{\$} \equiv \frac{1 + i_t^{\$}}{1 + \pi_{t+1}^{\$}} \simeq 1 + i_t^{\$} - \pi_{t+1}^{\$} \tag{11}$$

Corresponding expressions can be obtained for equities and bonds. These allow for changes in capital value other than through exchange rate changes. Let d_t^j be the j-th country's dividend rate at the end of period t, and let ΔS_{t+1}^{j} be the % change in equity prices in the course of period t+1 (the first difference of the logs). The real one-period return on holding equity in currency j is then

$$(12) \qquad 1 + r^{sj}_{t+1} = \frac{1 + d_t^{\,j} + \Delta S^{\,j}_{t+1}}{(1+\pi^{\$}_{t+1})\ (1+\Delta\ e^{\,j}_{t+1})}$$

$$\simeq 1 + d_t^{\,j} + \Delta\ S^{\,j}_{t+1} - \pi^{\$}_{t+1} - \Delta e^{\,j}_{t+1}$$

Similarly, the real return from holding bonds in currency j is given by

$$(13) \qquad 1 + r_{t+1}^{\,bj} \equiv \frac{1 + c_t^{\,bj} + \Delta b^{\,j}_{t+1}}{(1+\pi^{\$}_{t+1})\ (1+\Delta e^{\,j}_{t+1})}$$

$$\simeq 1 + c_t^{bj} + \Delta b_t^j - \pi^{\$}_{t+1} - \Delta e^j_{t+1}$$

where $c_t^{\,bj}$ refers to a coupon rate of interest.

Next we consider relative (or excess) rates of return. A vector of ex-post relative rates of return on foreign assets versus a U.S. dollar asset can be written:

$$(14) \qquad rr_{\sim t+1} \equiv r_{\sim t+1} - \underset{\sim}{\iota}\, r^{\$}_{t+1} \simeq i_{\sim t} - \underset{\sim}{\iota} i_t^{\$} - \Delta e_{\sim t+1}$$

where i_t is an n-1 vector of nominal interest rates, and Δe_{t+1} is an n-1 vector of exchange rate changes. Clearly relative real and relative nominal rates of return are the same since inflation is common and cancels out. Formulae (9)-(11) apply to bank deposits and similar fixed nominal assets, and indeed the data used in empirical applications are eurocurrency deposit rate data, available from DRI, and which are selected for consistency with spot and forward rate data. We can also construct analogous relative rates of return for equities and bonds. At this point we diverge further from Frankel (1982) since now any one of two dollar assets will be the numeraire. In our empirical work we shall choose the U.S. government securities.

From (12) the real rate of return for U.S. equities relative to U.S. bonds (which has no exchange rate effect) is:

$$(15) \qquad 1 + r^{\$b\$}_{t+1} \simeq 1 + d^{j}_{t} + \Delta\, s^{j}_{t+1} - c_{t}^{\$} - \Delta b^{\$}_{t+1} - \Delta e^{j}_{t+1}$$

In general, other real returns relative to US bonds are given by:

$$(16) \qquad \underset{\sim}{r}_{t+1} - \underset{\sim}{\iota} r^{b\$}_{t+1} \simeq \underset{\sim}{R}_{t} + \Delta \underset{\sim}{P}_{t+1} - \underset{\sim}{\iota}(c_{t}^{\$} + \Delta b^{\$}_{t}) - \Delta\, e_{t+1}$$

where $\underset{\sim}{R}_t$ is c^j_t or d^j_t depending on whether the asset is a bond or equity, and ΔP_{t+1} is the change in the price of the bond or equity.

Tables 2a–2e contain descriptive statistics for the major rates of return in our dataset. Table 2a refers to the simple nominal interest rates available on eurodeposits. These can be compared with the highly variable monthly equity returns in Table 2b. We use Datastream dividend yields and market indices to work out equity returns. These are in domestic currency. The UK actually had a 40% return on one-month (1974(1)). The Japanese market is noticeably less volatile than the rest, but still much more so than euro-deposit rate returns. Table 2c shows Table 2a returns adjusted for exchange rate changes, and naturally the range of values rises as do the coefficients of variation. However, including exchange rate adjustments does not alter the range of values of equity returns in the same way.

We suggested earlier that euro-deposit rates may have little to do with government bond returns and hence may be inappropriate measures of the rate of return on those assets. There are many types of bond and they have different returns, but there is no data comparable with that for equity markets to gauge overall bond price changes and coupon flows. This may be one reason why previous authors have relied on euro-rates. Salomon Brothers, however, have produced an index since 1978 for the major bond markets which we believe to be more suitable as a measure of bond market returns. This index is described in detail in a pamphlet entitled 'Introducing the Salomon Brothers World Government Bond Index' (November, 1986).

Salomon Brothers construct a performance index which includes accrued interest and price changes weighted by amounts outstanding at par and is based on a wide variety of the bonds they trade. Table 2e shows the descriptive statistics for rates of change of this index, i.e. monthly total returns. Clearly the change in capital values is now an important aspect of the rate of return for both the range and the coefficient of variation is much higher than for Table 2a. We suggest that such rates of return are appropriate for use with government bond wealth data. These data are also available from Datastream.

Table 3 compares the rate of return using the Salomon figures with euro-rates over the period 1978(2)-1988(1). Although the average rates of return are similar, their standard deviations are approximately 10 times greater than those of the euro-rates. The simple correlations between the same country returns are very low: .132 for West Germany, .169 for Japan, .217 for the UK and .048 for the US. In other words, it is most unlikely that using eurodeposit rates is remotely appropriate for measuring rates of return on the bond portfolios which have been so extensively explored recently in the mean-variance strand of the literature. The cross country correlations are also generally slightly lower for the Salomon data. It should be noted that these rates of return do not include exchange rate changes.

4. ICAPM Estimates for Four Countries and Two Assets

4.1. Specification and Estimation

The full system of equations defined across both different assets and countries is given in matrix form as equations (7) which can be rewritten as

(17) $$\underset{\sim}{z}_{t+1} = \underset{\sim}{\alpha} + \underset{\sim}{\beta}_t \underset{\sim}{x}_t + \underset{\sim}{\varepsilon}_{t+1}$$

where $\underset{\sim}{\beta}_t = \rho_t \Omega_t$ and $\underset{\sim}{\alpha}$ is an intercept which has been added to (7).

If we assume that all errors are distributed normally, then the log-likelihood for observation t is given in general, with time-varying covariance matrix, as

(18) $$\log L = -(N/2)\log(2\pi) - \tfrac{1}{2}\log|\Omega_t| - \tfrac{1}{2}\varepsilon_t'\Omega_t^{-1}\varepsilon_t$$

where N is the number of assets in the portfolio. We can write Ω_t as $P_t'P_t$ where P_t is an upper triangular matrix.

For static CAPM we can write $\underset{\sim}{\beta} = \rho\Omega$. Recently, however, researchers have focussed on relaxing the restrictions that Ω is fixed, largely, perhaps, due to the arrival of ARCH technology. This gives the model of Engel and Rodrigues (1989). In addition, they introduce macro-surprises into the specifications of the covariance matrix. Giovannini and Jorion (1989) also try a range of ARCH specifications. None of the authors carry out misspecification tests on the static CAPM as a prelude to ARCH or related estimation. It is interesting, therefore, that Wickens and Thomas (1989) find that there is little indication of ARCH effects in the relative rates of return. It may well be that the ARCH terms or macro-surprise parameters are picking up 'one-off' events which could equally well be treated (and indeed interpreted) by a simple dummy variable.

Our estimation period involves 1973(6)-1987(12) though the Salomon Brothers bond returns are only available since 1978(1). Other studies have three important empirical differences:

(i) their estimation periods are shorter and exclude the rise of Japanese financial importance and the crash of October, 1987.

(ii) they exclude German, UK, and Japanese equities, and indeed also Japanese bonds.

(iii) they use inappropriate bond rates of return.

All calculations were carried out in GAUSS using standard optimisation algorithms, (see Dennis and Schnabel (1983)).

4.2. Static ICAPM Results

For the purposes of comparison and before reporting our results we initially we took the six countries and same data period as studied by Engel and Rodrigues (1989) and replicated their results. Since we believe their model uses inappropriate assets and countries, we estimate four alternative versions of the static ICAPM (see Tables 4a-4d). In Table 4a we present the results for the same data period as Engel-Rodrigues (1989), i.e. 1973(6)-1984(12), using aggregate equities and bonds for the US, UK, Germany and Japan. We obtain an estimate of ρ, the CRRA, of -95.97 with a t-value of 2.20. This compares with a value of -19.29 with a standard error of 42.68 for Engel and Rodrigues (1989) and -142.91 with a t-value of 2.97 for Giovannini and Jorion (1989). If we continue the data period to end-1987 (Table 4b) ρ rises to -30.4 with a t-value of 1.21. Other parameter values change substantially. In particular none of the intercepts is now significant.

Salomon Brothers bond return data is available from 1973(1) and hence we estimate our model using both the euro-rates (Table 4c) and the Salomon Brothers' rates (Table 4d) over this period. Clearly Tables 4b and 4c are very similar but Table 4d is very different. It has a poorly determined estimate of ρ of 2.15 and no significant intercepts, though the P parameters are surprisingly similar for Tables 4c and 4d.

Usually the only diagnostic statistics reported in empirical studies of ICAPM are t-statistics on individual coefficients and the log-likelihood values. Although most of the coefficients are significant, this is a misleading indicator of the explanatory power of the ICAPM because these significant estimates are of the residual covariance matrix. In contrast, the estimates of ρ are poorly determined. Without a significant and meaningful estimate of ρ the relative returns remain unexplained and the explanatory power of ICAPM will be weak. A similar interpretation applies to the finding of a significant likelihood ratio statistic for the hypothesis of a zero risk premium for the model as a whole against the alternative of a variable risk premium. The poor explanatory power of the model is brought out most clearly by single equation R^2 values. For the model of Table 4b (the standard data for the whole period) the maximum R^2 value is 0.015, and for the model of Table 4d(i.e. using the Salomon bond prices) the maximum R^2 is 0.20×10^{-4}.

One possible explanation for the poorness of fit of the ICAPM relates to the order of integrability of the relative rates of return and the asset shares. If, for example, the relative rates of return were I(1), i.e. non-stationary, but the asset shares were I(0), i.e. stationary, then even though a risk premium is present it will not be revealed by estimating equation (17) which would fit very badly.

Moreover, the residuals would be integrated.[7] In Table 5 we provide summary statistics relating to the order of integrability of the relative rates of return. Although the test statistics give conflicting results, the estimates of the dominant root are always close to zero and very far from unity. We are reasonably safe in concluding, therefore, that the relative rates of return are stationary and hence this is not a reason for the poor explanatory power of the model.

Turning to the residuals themselves, in Table 6 we report the correlation matrix of the estimates of Tables 4b and 4d. The large and significant positive correlations suggest that the foreign relative rates of return have been subject to a common shock associated with the US dollar.

In addition to measures of goodness of fit we have carried out tests of misspecification for serially correlated errors and ARCH errors. Rather than assuming that ARCH effects are present it seems sensible to carry out an ARCH test of misspecification first. Surprisingly this has not usually been done. The results of carrying out both types of tests on the residuals of each equation using the standard data and the Salomon data are reported in Table 7. There is relatively little evidence of serially correlated errors and only a little more of ARCH effects. Certainly there is no systematic evidence of ARCH effects at any particular lag, such as lag 1 or at seasonal frequencies. This is a major blow for the ARCH in mean version of ICAPM and calls into question the necessity for the estimates of Engel and Rodrigues, apart from the

[7] See Engle and Granger (1987) for a discussion of integrability and Wickens and Thomas (1989) for a discussion of the use of the concept in analysing models of excess (or relative) rates of return.

issue of the power of LM and Wald type tests. Partly for completeness, however, in section 4.4 we present ARCH-in-mean estimates of the 'dynamic' ICAPM.

Another issue in misspecification relates to the mis-measurement of asset stocks, particularly the stock of bonds. As remarked earlier, the stock of bonds is valued at the issue price and not at the current market price. It is shown in the appendix that this is likely to lead to downward biased estimates of ρ. These results suggest that the bias is likely to be greater, the greater the difference between the market price and the issue price.

Finally, and not too surprisingly, we may note that the restrictions implied by the ICAPM for Tables 4b and 4d are overwhelmingly rejected by the data.

4.3. Restricted Static ICAPM formulations

As the ICAPM above finds little support in the data, one question to ask is whether the CAPM fits better to restricted versions of the full model. It may be that a different CAPM explains the international bond data from that which explains the equity data. Or, a CAPM may fit national data on bonds and equity but may not carry over to an international CAPM. By estimating these restricted CAPMs it is possible to determine if there are national differences in attitudes to risk or if the international degree of risk aversion on bonds is different from that on equity. In a sense this is equivalent to carrying out another form of misspecification test for the whole model.

The results of estimating the restricted CAPMs are reported in Table 8. The principle way of comparing the models is through the estimates of ρ. These estimates show considerable variation between the different equations. The international equity CAPM has a positive and marginally significant ρ but the corresponding bond ρ is negative but insignificant. The estimate of ρ for the UK is positive but insignificant; the other estimates of ρ are all negative with that for the US highly significant and that for Japan marginally significant. Virtually all of the covariance terms are significant. The main conclusion to follow from these results is that there is sufficient disparity to cast doubt on the existence of a well defined general ICAPM of the sort estimated earlier. We note that when Giovannini and Jorion (1989) add the US equity market to their CAPM the estimate of ρ falls from -27.86 to -142.91, which is not too different from our estimate for US ρ of -186.07. Simple misspecification evidence is presented in Table 9 for the restricted ICAPM estimates of Table 8. While the bond market model, in the spirit of Frankel (1982), reveals little evidence of autoregressive or ARCH effects, the same is not true for either the equity market model or the individual country models, with much more noticeable ARCH effects present. This suggests that the range and type of assets considered has a major impact on the level of misspecification present, and calls into question the reliability of earlier results which involve only a small portion of the world portfolio. Earlier studies considering the mean-variance efficiency of the U.S. market portfolio (e.g. Stambaugh (1982)) do not consider the sensitivity of equation misspecification to alternative market portfolios; rather they focus only on the test of market efficiency in this context.

4.4 ICAPM with ARCH

Thus far we have assumed constant variances in the process generating excess returns (eq(17)), and indeed this assumption is supported by our misspecification tests in Table 7. However, it has been observed in various empirical contexts that such variances vary systematically over time (e.g. Giovannini and Jorian (1989)). The ARCH model of Engle (1982), which is agnostic about the economic structure postulates a relationships through time for the conditional variances. The multivariate ARCH(1) can be written (see also Engel and Rodrigues (1989)):

(19) $$\Omega_t = P'P + G\varepsilon_t\varepsilon_t'G$$

P is a constant upper triangular matrix as before; G is a constant symmetric matrix and ε_t is the lagged forecast error (i.e. the error made in predicting the relative returns between t-1 and t). This specification ensures that Ω_t is positive semi-definite.

Tables 10a and 10b present the results from estimating (19) on the full sample period ('old' data) and on the shorter sample period data using the Salomon Brothers returns. Initially G is chosen to be a diagonal matrix, hence 'own' lagged forecast errors only affect the 'own' conditional covariance. Clearly there is little substantive change between Tables 4b and 10a. The CRRA is still negative and insignificant, though 6 of the 7 G parameters have t-values greater than 1; together they are jointly significant with a $\chi^2(7)$ value of 19.2. The ARCH parameters range in (absolute) value from .07 to .36 and hence the variance process is stationary over time. Thus, although there is evidence that introducing ARCH effects (via a diagonal G matrix) has some added value, this way of formulating of time-varying second moments is not the panacea for the ills of the static ICAPM. Comparison of Tables 4c and 10b (the Salomon Brothers' data) yields similar conclusions.

A more general model is obtained by specifying G to be a full symmetric matrix (containing 28 parameters). The estimates for this specification are presented in Tables 10c and 10d. The G parameters are jointly significant, with a $\chi^2(28)$ of 69.4 for table 10c (the full period, 'old' data) and a $\chi^2(28)$ of 120 for the shorter period with the Salomon returns. Using the 'old' data, the estimates of the CRRA is negative and highly significant while using the 'new' data it is negative but, as earlier, not statistically different from zero.

We have to conclude , therefore, that the introduction of first-order ARCH terms to facilitate time-varying covariances seems to provide relatively little empirical or economic improvement over the static ICAPM. Nor are the estimates of the CRRA more sensible or stable.

4.5 The Behaviour of Risk Premia and Conditional Variances

In this section we examine the time path of both the 'risk premia' and the conditional standard errors obtained from the estimates of Table 10b and 10d, i.e., the models with the full G matrix in the ARCH processes for time-varying second moments. We have already suggested that the use of the eurodollar interest rate to represent the return on US bonds is inappropriate and hence the interpretation of the fitted values of equation (17) as 'risk premia' is somewhat dubious. Nevertheless, in the spirit of empirical work in this area we present the risk premia on US equities and West German bonds in Figures 1a and 1b as representative of the estimated risk premia. Compared with similar risk premia in Giovannini and Jorion (1989, Figures 3 and 4, pages 320-321) our results are much more volatile and larger in absolute value.

In addition and in contrast to the estimates of Giovannini and Jorion our risk premia can take both positive and negative values. (see also Wickens and Thomas (1989)). In fact our results are much closer in this respect to the risk premia estimates in their Figures 1 and 2 (pages 318-319) which are derived from 'unrestricted' single-equation projections. It can be seen from our Figure 1, that even on US equities the risk premium is not always positive. (Indeed see Bollerslev, Engle and Wooldridge (1988) where it is always negative in a domestic US version of the CAPM.) Giovannini and Jorion make no attempt to reconcile the results in their two sets of figures. It may be noted, however, that CAPM theory assumes the existence of a risk free asset against which to define the risk premium. In the ICAPM model as estimated above there is no risk free asset, only relative degrees of risk. For example, US bonds are not risk free except in the sense of default risk: their capital value is as volatile as equity prices, and hence a positive risk premium relative to US bonds is not really the same as finance theory's risk premium with respect to a safe asset. In Figures 2a and 2b we present the fitted values from Table 10d for West German bonds and US equities: in this case the bond returns are calculated from the Salomon Brothers data and, as one would expect, the estimated risk premia are much more volatile than for the earlier cases. However, we still clearly have both positive and negative values for the risk premia.

In Figures 3a and 3b we present the conditional variances for West German bonds and US equities, for the 'old' data, and in Figures 4a and 4b we present conditional variances for the same equations using Salomon's data. Clearly the latter are more volatile as we would now expect, and indeed are greater in absolute value than the results presented in Giovannini and Jorion (1989) for a similar data period.

5. Conclusions

Fundamentals explanations of exchange rate movements based on models that assume uncovered interest parity are widely, though possibly mistakenly, thought to have failed. As a result attempts have been made to test whether or not a better fundamentals explanation is given by the assumption of imperfect substitutability due to risk aversion. This has led to the assumption of the representative international investor and to the development of an alternative empirical model, namely, the international CAPM. Unfortunately, estimates of the ICAPM have been less than encouraging. The purpose of this paper has been to try to identify whether this unsatisfactory performance is due to deficiencies in the econometric analysis or to the model itself.

We have identified a number of possible reasons why previous estimates of the ICAPM may have caused the model to perform badly: the omission of important assets or countries from the world portfolio, the use of inappropriate rates of return, the valuation of bonds and hence the calculation of asset shares using the issue price instead of the market price, non-stationarity of the relative rates of return but stationarity of the asset shares, the misinterpretation of temporary shocks as systematic ARCH effects and different attitudes to risk between the investors of one country and another and between asset types. This list could easily be extended but it is long enough already and contains several powerful reasons for the failure of the ICAPM. We have tried to assess the empirical importance of each of these possible causes of the failure of the ICAPM.

Our results confirms the findings of Wickens and Thomas (1989) that there is little evidence of a large or even systematic risk premium over the whole data period. On average, risk neutrality seems to be a better description of the data. However, there are periods when relative rates of return are clearly out of line with risk neutrality and this seems to be associated with "news" about the US dollar. Removing some of the deficiencies in the data does not have a decisive impact on the estimates though it does bring about improvements. There is a suggestion that the assumption of a single ICAPM covering all countries and all assets is incorrect. In short, our results confirm those in the literature without indicating where the solution lies. They suggest that the main problem which remains is to find an explanation for the apparent short-run variation in risk premia that is identified by Wickens and Thomas. But in our judgement the answer to this is unlikely to lie in the use of the ARCH in mean CAPM.

Appendix

The effects of mis-measuring bond stocks

For simplicity consider a two asset world with a safe asset and a risky asset (a bond) whose value is measured incorrectly at the issue price instead of the market value at the beginning of the period (or end of last period).

The static CAPM is the single equation

$$E_t(r^b_{t+1} - r_{t+1}) = \alpha + \beta x_t \tag{A1}$$

where r^b_t and r_t are the ex-post rates of return on the risky and safe assets, respectively, x_t is the share of the risky asset in the portfolio and expectations are assumed to be rational. Suppose r^b_t is the holding period return on the bond then

$$r^b_{t+1} = (P^b_{t+1} - P_t{}^b + c_t)/P_t{}^b \tag{A2}$$

wherer P^b_t is the bond price and c_t is a fixed coupon. Assume that there is a fixed stock of bonds, for convenience say one bond, then the market value of bonds at the beginning of the current period is P^b_t where P^b_t is the bond price and c_t is a fixed coupon. Assume that there is a fixed stock of bonds, for convenience say one bond, then the market value of bonds at the beginning of the period is also P^b_t. Also let the rate of return on the safe asset be constant ($r_t = \overline{r}$ for all t). This implies that market value is also constant at $\overline{p}$, say. The market value of bond shares is therefore

$$x_t = p^b_t/(p^b_t + \overline{p}) \tag{A3}$$

If

$$\log P_{t+1} = E_t \log P_{t+1} + e_{t+1} \tag{A4}$$

where e_t is iid $(0, \sigma^2)$ then

$$y_{t+1} = r_{t+1}{}^b - r_{t+1} = E_t(r^b_{t+1} - r_{t+1}) + e_{t+1} \tag{A5}$$

and so

$$y_{t-1} = \beta x_t + e_{t+1} \tag{A6}$$

If, for the purposes of measuring x_t the bond is valued at its issue price, say $\tilde{p}_t$, then the assumed value of x_t is

$$\tilde{x}_t = \tilde{p}_t/(\tilde{p}_t + \bar{p}) \tag{A7}$$

and is a constant. It follows that

$$y_{t+1} = \alpha+\beta\tilde{x}_t + u_{t+1} \tag{A8}$$

with

$$u_{t+1} = e_{t+1} + \beta(x_t - \tilde{x}_t) \tag{A9}$$

where

$$x_t - \tilde{x}_t = \frac{\bar{p}(p_t^b - \tilde{p}_t)}{(p_t^b + \bar{p})(\tilde{p}_t + \bar{p})} \tag{A10}$$

The asymptotic bias in b, the OLS estimator of β from (A8) is given by

$$\text{plim } b - \beta = \frac{\beta \text{cov}(\tilde{x}_t, x_t - \tilde{x}_t)}{\text{var}(\tilde{x}_t)}$$

$$= \beta\{\theta\ \rho[\text{var}(p^b)/\text{var}(\tilde{p})]^{\frac{1}{2}} - 1\} \tag{A11}$$

where θ = (1-mean x)/(1-mean $\tilde{x}$) and ρ is the correlation between p_t^b and $\tilde{p}_t$. In general the sign of the bias is indeterminate, but in practice, since the stocks of bonds is constantly changing with some being redeemed and new issues being made, it may be assumed that $\tilde{p}_t$ is a sort of moving average of current and past p_t^b and hence $\rho>0$. Assuming that $\theta\approx1$ and $\text{var}(p^b) \approx \text{var}(\tilde{p})$ we find therefore that the bias will be negative. In other words, the coefficient of relative risk aversion is likely to be biased downwards. Although in practice the formula for the bias will be more complicated, this result may go some way to account for the finding of a negative CRRA in the ICAPM estimates.

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Table 1: Countries' Share of total Assets - Bonds & Equity

Share of Bonds

Sample period: 73M6 to 87M12

Variable(s)	WGSHG	JASHG	UKSHG	USSHG
Maximum	.0734	.1788	.0906	.3049
Minimum	.0212	.0271	.0118	.2359
Mean	.0451	.1035	.0499	.2359
Std. Deviation	.0149	.0418	.0250	.0319
Coef. of Variation	.3306	.4042	.5018	.1352

Estimated Correlation Matrix of Variables

	WGSHS	JASHG	UKSHG	USSHG
WGSHG	1.0000	-.4799	.8646	-.2687
JASHG	-.4799	1.0000	-.6288	-.1181
UKSHG	.8646	-.6288	1.0000	-.2946
USSHG	-.2687	-.1181	-.2946	1.0000

Share of Equity

Sample period: 73M6 to 87M12

Variable(s)	WGSHE	JASHE	UKSHE	USSHE
Maximum	.0394	.2720	.0852	.4743
Minimum	.0181	.0876	.0279	.2258
Mean	.0287	.1316	.0540	.3514
Std. Deviation	.0059213	.0435	.0099587	.0511
Coef. of Variation	.2066	.3306	.1845	.1454

Estimated Correlation Matrix of Variables

	WGSHE	JASHE	UKSHE	USSHE
WGSHE	1.0000	-.0865	-.1533	.1358
JASHE	-.0865	1.0000	.6637	-.7583
UKSHE	-.1533	.6637	1.0000	-.5464
USSHE	.1358	-.7583	-.5464	1.0000

Key:

WGSHG, JASHG, UKSHG, USSHG: West Germany, Japan, UK and US shares of their respective government bond stocks in total financial wealth. Source: Alberto Giovannini and our own estimates.

Similarly for equities: WGSHE, JASHE, UKSHE, USSHE. Source: Datastream

Table 2a: Returns on Bonds - Domestic Currency

Sample period: 73M6 to 87M12

Variable(s)	WGINTL	JAINTL	UKINTL	USINTL
Maximum	.0250	.0414	.0200	.0154
Minimum	.0015441	-.0214	.0038137	.0039670
Mean	.0048319	.0054609	.0098420	.0076218
Std. Deviation	.0025976	.0051669	.0025050	.0026659
Coef. of Variation	.5736	.9462	.2545	.3498

Estimated Correlation Matrix of Variables

	WGINTL	JAINTL	UKINTL	USINTL
WGINTL	1.0000	.3812	.4041	.5459
JAINTL		1.0000	.4117	.1315
UKINTL			1.0000	.3040
USINTL				1.0000

Table 2b: Returns on Equity - Domestic Currency

Sample period: 73M6 to 87M12

Variable(s)	WGREQ	JAREQ	UKREQ	USREQ
Maximum	.1363	.1211	.4163	.1646
Minimum	-.2431	-.1354	-.3030	-.2402
Mean	.0059337	.0113	.0137	.0088828
Std. Deviation	.0471	.0402	.0743	.0515
Coef. of Variation	7.9323	3.5505	5.4186	5.7955

Estimated Correlation Matrix of Variables

	WGREQ	JAREQ	UKREQ	USREQ
WGREQ	1.0000	.3085	.4169	.3964
JAREQ		1.0000	.3621	.3859
UKREQ			1.0000	.5869
USREQ				1.0000

Key:

WGINTL, JAINTL, UKINTL, WINTL, are one-month eurodeposit rates, end-of-month, for West Germany, Japan , the UK and the US respectively. Source: DRI.

WGREQ, JAREQ, UKREQ, USREQ are monthly equity returns for the same countries. Source: Datastream.

Table 2c: Returns on Bonds - Nominal - Dollar Terms

Sample period: 73M6 to 87M12

Variable(s)	: WGNOM	JANOM	UKNOM	USNOM
Maximum	: .0865	.1029	.1440	.0154
Minimum	: -.0873	-01170	-.0673	.0039670
Mean	: .0069422	.0095546	.0076836	.0076218
Std. Deviation	: .0342	.3040	.0324	.0026659
Coef. of Variation	: 4.9214	3.5606	4.2171	.3498

Estimated Correlation Matrix of Variables

	WGNOM	JANOM	UKNOM	USNOM
WGNOM	1.0000	.5924	.6026	-.1472
JANOM		1.0000	.5080	-.1591
UKNOM			1.0000	-.0722
USNOM				1.0000

Table 2d: Returns on Equity - Nominal - Dollar Terms

Sample period: 73M6 to 87M12

Variable(s)	: WGRREQ	JARREQ	UKRREQ	USRREQ
Maximum	: .1664	.1714	.4284	.1646
Minimum	: -.1817	-.1279	-.2510	-.2402
Mean	: .0080440	.0154	.0116	.0088828
Std. Deviation	: .0578	.0566	.0830	.0515
Coef. of Variation	: 7.1827	3.6704	7.1767	5.7955

Estimated Correlation Matrix of Variables

	WGRREQ	JARREQ	UKRREQ	USRREQ
WGRREQ	1.0000	.4401	.4479	.3853
JARREQ		1.0000	.3612	.2907
UKRREQ			1.0000	.5315
USRREQ				1.0000

Table 2e: Rates of Change - Salomon Bros. Data

Sample period: 78M2 to 89M2#

Variable(s)	RCWGSB	RCJASB	RCUKSB	RCUSSB
Maximum	.0559	.0559	.0902	.1213
Minimum	-.0543	-.0617	-.0766	-.0817
Mean	.0055275	.0098904	.0093094	.0075833
Std. Deviation	.0166	.0179	.0301	.0320
Coef. of Variation	3.0024	2.7574	3.0653	4.2164

Estimated Correlation Matrix of Variables

	RCWGSB	RCJASB	RCUKSB	RCUSSB
RCWGSB	1.0000	.5877	.3435	.5210
RCJASB		1.0000	.3674	.4036
RCUKSB			1.0000	.3799
RCUSSB				1.0000

Key:

RCWGSB, RCJASB, RCUKSB, and RCUSSB are the monthly total returns on government bonds in domestic currency using the Salomon Brothers indices for West Germany, Japan, the UK and US respectively. Source: Datastream.

Table 3: Descriptive Statistics & Correlations

Sample period: 78M2 to 88M1#

Variable(s)	RCWGSB	WGINTL	RCJASB	JAINTL	RCUKSB	UKINTL	RCUSSB	USINTL
Maximum	.0559	.0119	.0686	.0114	.0902	.0149	.1213	.0154
Minimum	-.0543	.0015441	-.0617	-.0013747	.0054800	.0054800	-.0817	.0047082
Mean	.0058945	.0047145	.0068329	.0045269	.0095930	.0095930	.0081505	.0083407
Std. Deviation	.0170	.0019967	.0183	.0020852	.0080704	.0080704	.0331	.0086473
Coef. of Variation	2.8904	.4235	2.6768	.4606	.8158	.2158	4.0564	.3174

Estimated Correlation Matrix of Variables

	RCWGSB	WGINTL	RCJASB	JAINTL	RCUUSB	UKINTL	RCUSSB	USINTL
RCWGSB	1.0000	.1316	.5911	.1738	.3391	-.0037	.5137	-.0165
WGINTL		1.0000	.0517	.5625	.0467	.5469	-.0183	.7732
RCJASB			1.0000	.1691	.3577	-.0156	.4108	-.0090
JAINTL				1.0000	.1271	.5603	.0651	.3466
RCUBSB					1.0000	.2174	.3774	.0137
UKINTL						1.0000	-.0585	.4795
RCUSSB							1.0000	.0480
USINTL								1.0000

Key: As Tables 2a-2e

Table 4a: 1973(6)-1984(12)

CAPM Estimation: Constant Omega

$$z_{t+1} = c + \rho(P'P)_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = P'P$$

Log Likelihood: 2129.482678

Estimate of the vector c':

BONDS W.GERMANY	JAPAN	UK	EQUITY W.GERMANY	JAPAN	UK	US
0.034804 (1.63)	0.039890 (1.77)	0.026324 (1.48)	0.079603 (2.06)	0.094028 (2.11)	0.153192 (2.22)	0.109982 (2.29)

Estimate of the coefficient of ρ:

-95.972849
(-2.20)

Estimate of the upper triangular matrix P:

0.031703 (14.59)	0.016491 (5.37)	0.015802 (5.55)	0.034153 (7.83)	0.023115 (4.06)	0.019879 (1.98)	0.011548 (2.42)
0	0.027814 (16.04)	0.006433 (2.83)	0.005522 (1.06)	0.033773 (5.79)	0.012992 (1.12)	0.000812 (0.15)
0	0	0.024073 (15.52)	0.002693 (0.61)	-0.003675 (-0.87)	0.033804 (3.14)	-0.001151 (-0.18)
0	0	0	0.036897 (14.51)	0.012101 (3.16)	0.030462 (4.04)	0.015719 (3.02)
0	0	0	0	0.032381 (15.16)	0.019107 (2.24)	0.010870 (2.90)
0	0	0	0	0	0.067794 (24.96)	0.017153 (3.59)
0	0	0	0	0	0	-0.039410 (-14.38)

(t-statistics in parentheses)

Sample Size: 139
Degrees of Freedom: 103

Table 4b: 1973(6)-1987(12)

CAPM Estimation: Constant Omega

$$z_{t+1} = c + \rho(P'P)_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = P'P$$

Log Likelihood: 2580.035173

Estimate of the vector c':

BONDS W.GERMANY	JAPAN	UK	EQUITY W.GERMANY	JAPAN	UK	US
0.012799	0.016762	0.010991	0.031991	0.043316	0.058630	0.039759
(1.04)	(1.26)	(1.09)	(1.25)	(1.44)	(1.42)	(1.32)

Estimate of the coefficient of ρ:

-30.439128
(-1.21)

Estimate of the upper triangular matrix P:

0.034793 (15.23)	0.021000 (7.44)	0.020383 (8.41)	0.033731 (8.08)	0.024489 (4.99)	0.020269 (2.42)	0.006099 (1.48)
0	0.027824 (18.63)	0.005983 (3.01)	0.003567 (0.58)	0.033543 (7.65)	0.009444 (0.99)	-0.001686 (-0.33)
0	0	0.025377 (20.77)	0.001648 (0.32)	-0.003354 (-0.75)	0.032993 (3.37)	-0.002175 (-0.36)
0	0	0	0.047298 (19.60)	0.011913 (3.73)	0.030678 (4.98)	0.019530 (4.53)
0	0	0	0	0.038542 (18.63)	0.018549 (2.63)	0.014254 (4.23)
0	0	0	0	0	0.067798 (28.09)	0.020748 (4.58)
0	0	0	0	0	0	-0.039748 (-16.89)

(t-statistics in parentheses)

Sample Size: 175
Degrees of Freedom: 139

Table 4c: 1978(1)-1987(12)

CAPM Estimation: Constant Omega

$$z_{t+1} = c + \rho(P'P)_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = P'P$$

Log Likelihood: 1795.992468

Estimate of the vector c':

BONDS W.GERMANY	JAPAN	UK	EQUITY W.GERMANY	JAPAN	UK	US
0.009176	0.017114	0.010602	0.024466	0.042847	0.040589	0.030398
(0.53)	(0.72)	(0.63)	(0.68)	(0.89)	(0.85)	(0.79)

Estimate of the coefficient of ρ:

-24.047845
(-0.65)

Estimate of the upper triangular matrix P:

0.037183	0.025004	0.024277	0.034994	0.024340	0.018124	0.000334
(9.94)	(5.29)	(6.44)	(5.46)	(3.00)	(2.24)	(0.06)
0	0.031414	0.004524	0.002584	0.040148	0.006541	-0.000241
	(12.66)	(1.74)	(0.26)	(5.66)	(0.81)	(-0.04)
0	0	0.026375	-0.001939	-0.003921	0.029849	0.001511
		(14.87)	(-0.30)	(-0.62)	(3.86)	(0.19)
0	0	0	0.048824	0.009803	0.023982	0.020345
			(12.35)	(2.35)	(4.25)	(3.68)
0	0	0	0	0.037135	0.015882	0.014555
				(12.76)	(2.76)	(3.03)
0	0	0	0	0	0.047997	0.022272
					(10.31)	(3.51)
0	0	0	0	0	0	-0.035998
						(-13.92)

(t-statistics in parentheses)

Sample Size: 120
Degrees of Freedom: 84

Table 4d: 1978(1)-1987(12), Salomon Bros. Data

CAPM Estimation: Constant Omega

$$z_{t+1} = c + \rho(P'P)_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = P'P$$

Log Likelihood: 1682.047086

Estimate of the vector c':

BONDS W.GERMANY	JAPAN	UK	EQUITY W.GERMANY	JAPAN	UK	US
-0.001338 (-0.05)	0.002618 (0.08)	0.000554 (0.02)	-0.000977 (-0.03)	0.010198 (0.20)	0.006543 (0.13)	0.002847 (0.07)

Estimate of the coefficient of ρ:

2.150216
(0.06)

Estimate of the upper triangular matrix P:

0.046632 (10.66)	0.031947 (4.95)	0.026051 (4.92)	0.037076 (5.72)	0.028532 (3.55)	0.020327 (2.49)	0.004264 (0.66)
0	0.039718 (13.21)	0.008547 (1.86)	0.000635 (0.07)	0.041169 (6.41)	0.006895 (0.80)	0.000048 (0.01)
0	0	0.043833 (12.62)	-0.000834 (-0.13)	-0.001210 (-0.22)	0.035078 (4.93)	0.004890 (0.79)
0	0	0	0.048113 (11.10)	0.012376 (3.03)	0.024440 (5.05)	0.020102 (3.80)
0	0	0	0	0.038115 (13.70)	0.015592 (3.24)	0.017021 (3.35)
0	0	0	0	0	0.045527 (10.15)	0.022234 (4.03)
0	0	0	0	0	0	0.036889 (14.17)

(t-statistics in parentheses)

Sample Size: 120
Degrees of Freedom: 84

Table 5 Order of Integrability of relative rates of return

Constant	-0.000 (0.02)	0.003 (0.99)	0.001 (0.41)	0.003 (0.62)	0.008 (1.70*)	0.007 (1.10)	0.002 (0.52)	-0.002 (0.05)	0.005 (1.01)	0.002 (0.42)	0.001 (0.13)	0.012 (2.01*)	0.010 (1.60)	0.05 (1.12)
$\hat{\alpha}-1$	-1.015	-0.883	-0.900	-0.908	-0.840	-0.860	-0.949	-1.105	-1.025	-0.870	-0.984	-0.920	-0.988	0.939
DF	-13.3*	-11.42*	-11.84*	-11.93*	-11.15*	-11.31*	-12.39*	-11.83*	-10.94*	-9.38*	-10.33*	-9.98*	-10.65*	-10.19*
ADF(12)	-2.36	-2.95*	-2.26	-2.82	-2.55	-4.54*	-3.99*	-2.64	-3.03*	-2.89*	-2.04	-2.48	-3.04*	-3.14*
Z_α	-14.85*	-8.68	-8.33	-8.59	-10.78	-9.17	-10.43	-8.82	-9.22	-6.97	-7.49	-8.51	-7.10	-6.41
Z_t	-2.79	-1.79	-2.06	-2.08	-2.35	-2.07	-2.23	-1.91	-1.98	-1.70	-1.59	-2.07	-1.83	-1.81

Notes

(i) The equation estimated is $\Delta x_{t+1} = \mu + (\alpha-1)\, x_t + e_{t+1}$.

(ii) The numbers in parentheses are the absolute values of t-statistics.

(iii) Z_α is the Phillips (1987) test for H_o: $T(\alpha-1) = 0$
Z_t is the Phillips (1987) test for H_o: $\alpha=1$ based on the t-statistic.

(iv) * denotes a critical value at the 5* level.

These are DF, ADF and $2_t = -2.88$, $Z_\alpha = -14.0$.

Table 6

Correlation Matrix for the Residuals, Table 4b

WG Bonds	Japan Bonds	UK Bonds	WG Equity	Japan Equity	UK Equity	US Equity
1.00000000	0.60242070	0.61590036	0.57930708	0.42225561	0.23450152	0.11870721
	1.00000000	0.51522953	0.39788365	0.71601871	0.22848798	0.04531923
		1.00000000	0.38957343	0.32028233	0.45689667	0.03471843
			1.00000000	0.44526862	0.44167455	0.37433518
				1.00000000	0.35567592	0.29604627
					1.00000000	0.51932118
						1.00000000

Correlation Matrix for the Residuals, Table 4d

WG Bonds	Japan Bonds	UK Bonds	WG Equity	Japan Equity	UK Equity	US Equity
1.00000000	0.62675761	0.50387389	0.61939231	0.44470946	0.29956444	0.08376822
	1.00000000	0.44462228	0.39065646	0.77872662	0.26693295	0.05323716
		1.00000000	0.29760434	0.31416631	0.60601968	0.12381048
			1.00000000	0.43114364	0.46204474	0.36257838
				1.00000000	0.39465788	0.31087259
					1.00000000	0.58699332
						1.00000000

Table 7 Residual diagnostics for Tables 4b and 4d

	Table 4b ('old' rates, full period)		Table 4d (Salomon Brothers)	
Equation	Residuals	Residuals Squared	Residuals	Residuals Squared
WG Bonds	-	1(1)	-	-
Japan Bonds	-	2(4,5)	1(7th)	2(11,12)
UK Bonds	-	-	1(1st)	-
WG Equities	1(6)	1(3)	-	1(1)
Japan Equities	-	-	-	-
UK Equities	1(1)	1(1)	-	-
US Equities	-	1(9)	-	1(9)

Table 8

1972(6)-1984(12)

	SINGLE ASSET INTERNATIONAL CAPM		SINGLE COUNTRY CAPM, EQUITIES AND BONDS			
	Bonds	Equity	W.Germany	Japan	UK	US
obs	139	139	139	139	139	139
L.L.	646.505832	-362.632221	-214.424490	-246.244751	-260.244751	-220.936013
CONSTANTS						
BONDS						
W. Germany	-0.000701					
t.	-0.059095					
Japan	0.002641					
t	0.180141					
UK	-0.000224					
t	-0.020571					
EQUITY						
W. Germany		-0.245287	0.026862			
t		-0.085246	0.781860			
Japan		-11.332667		0.045506		
t		-1.768101		2.059615		
UK		-11.120293			-0.009568	
t		-1.624693			-0.703438	
US						0.014532
t						2.722194
RHO						
RHO	-13.021552	23.007259	-80.095865	-52.196384	12.747611	-186.074258
t	-0.346418	1.877571	-0.713222	-1.655084	0.671150	-2.472875
P Matrix						
1,1	0.031783	0.134113	0.028665	0.036038	0.039981	0.030039
t	16.970490	2.241327	19.304332	31.629811	25.159383	19.650606
2,1	0.016652	0.172114				
	6.065315	0.061089				
2,2	0.028106	1.193651				
	17.823763	2.177062				
3,1	0.015939	0.228474				
	6.866585	0.083867				
3,2	0.006494	1.155674				
	3.673680	2.028798				
3,3	0.024108	0.128464				
	16.776296	10.347254				

Table 9

Misspecification in the Restricted CAPM, Table 3

Bond Markets Only

Equation	Residuals	Residuals Squared
WG Bonds	-	-
Japan Bonds	-	2(1,4)
UK Bonds	-	-

Equity Markets Only

	Residuals	Residuals Squared
WG Equities	3(2,6,8)	3(2,3,8)
Japan Equities	2(1,2)	2(1,2)
UK Equities	1(1)	2(1,5)

Individual Country Markets, 2 Assets

	Residuals	Residuals Squared
WG	-	3(1,4,5)
Japan	-	-
UK	-	3(1,5,10)
US	-	2(3,12)

Table 10a: 1973(6)-1987(12)

CAPM Estimation: 7 parameter ARCH

$$z_{t+1} = c + \rho\Omega_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = \Omega_t = P'P + G\varepsilon_t\varepsilon_t'G'$$

Log Likelihood: 2589.634027

Estimate of the vector c':

BONDS W.GERMANY	JAPAN	UK	EQUITY W.GERMANY	JAPAN	UK	US
0.004576 (0.44)	0.007631 (0.65)	0.004390 (0.50)	0.014379 (0.62)	0.024021 (0.86)	0.028428 (0.74)	0.017733 (0.65)

Estimate of the coefficient of ρ:

-12.520058
(-0.55)

Estimate of the upper triangular matrix P:

0.033996 (14.30)	0.019172 (6.05)	0.020468 (8.53)	0.033357 (7.52)	0.023879 (4.09)	0.019214 (2.33)	0.005992 (1.34)
0	0.025903 (14.47)	0.006330 (2.37)	0.003327 (0.49)	0.033063 (6.94)	0.009258 (0.90)	-0.002001 (-0.36)
0	0	0.025073 (20.11)	0.002400 (0.46)	-0.003687 (-0.80)	0.033376 (3.33)	-0.002769 (-0.42)
0	0	0	0.046166 (16.25)	0.011103 (3.45)	0.029049 (4.14)	0.020848 (4.88)
0	0	0	0	0.037715 (15.65)	0.018842 (2.70)	0.015634 (4.03)
0	0	0	0	0	0.066505 (24.57)	0.021277 (4.39)
0	0	0	0	0	0	-0.037712 (-14.14)

Estimate of the diagonal elements of G

0.213121 (1.72)	0.346172 (3.30)	0.066139 (0.48)	0.208750 (1.98)	0.218853 (1.95)	0.143087 (1.09)	-0.115483 (-1.40)

(t-statistics in parentheses)

Sample Size: 175
Degrees of Freedom: 132

Table 10b: 1978(1)-1987(12)

CAPM Estimation: 7 parameter ARCH

$$z_{t+1} = c + \rho\Omega_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = \Omega_t = P'P + G\varepsilon_t\varepsilon_t'G'$$

Log Likelihood: 1689.537521

Estimate of the vector c':

BONDS W.GERMANY	JAPAN	UK	EQUITY W.GERMANY	JAPAN	UK	US
-0.022656 (-1.20)	-0.024478 (-1.04)	-0.021596 (-1.12)	-0.034508 (-1.14)	-0.034000 (-0.86)	-0.036786 (-1.00)	-0.031136 (-1.06)

Estimate of the coefficient of ρ:

31.235813
(1.27)

Estimate of the upper triangular matrix P:

0.046605 (11.06)	0.031618 (5.03)	0.026062 (4.59)	0.036662 (4.96)	0.028069 (3.45)	0.020128 (2.28)	0.004564 (0.67)
0	0.039572 (12.50)	0.009591 (1.89)	-0.000267 (-0.03)	0.040797 (5.74)	0.005185 (0.65)	-0.002929 (-0.40)
0	0	0.043352 (12.24)	-0.001747 (-0.26)	-0.001658 (-0.29)	0.034248 (4.37)	0.005572 (0.90)
0	0	0	0.046719 (10.35)	0.014172 (3.19)	0.022907 (4.54)	0.021702 (3.72)
0	0	0	0	0.037354 (10.03)	0.018985 (3.36)	0.019392 (3.59)
0	0	0	0	0	0.038812 (8.39)	0.017636 (3.22)
0	0	0	0	0	0	0.034931 (12.60)

Estimate of the diagonal elements of G

-0.013761 (-0.05)	-0.070275 (-0.41)	0.127919 (0.73)	0.216688 (1.49)	-0.048734 (-0.19)	0.417263 (3.00)	0.345763 (1.64)

(t-statistics in parentheses)

Sample Size: 120
Degrees of Freedom: 77

Table 10c: 1973(6)–1987(12)

CAPM Estimation: 15 parameter ARCH

$$z_{t+1} = c + \rho\Omega_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = \Omega_t = P'P + G\varepsilon_t\varepsilon_t'G'$$

Log Likelihood: 2614.727398

Estimate of the vector c':

BONDS W.GERMANY	JAPAN	UK	EQUITY W.GERMANY	JAPAN	UK	US
0.020980 (3.09)	0.025480 (3.13)	0.019700 (3.21)	0.050614 (3.44)	0.060680 (3.54)	0.084495 (3.47)	0.060912 (3.44)

Estimate of the coefficient of ρ:
−47.199714
(−3.31)

Estimate of the upper triangular matrix P:

0.033454 (18.87)	0.021489 (8.86)	0.020820 (9.19)	0.034323 (9.45)	0.025136 (6.16)	0.023067 (3.92)	0.007688 (2.26)
0	0.026516 (17.17)	0.008090 (3.66)	0.002829 (0.73)	0.033254 (9.38)	0.009711 (1.44)	−0.000555 (−0.14)
0	0	0.021659 (13.54)	0.000534 (0.13)	−0.004975 (−1.50)	0.041216 (6.48)	−0.002415 (−0.66)
0	0	0	0.046592 (17.98)	0.010050 (3.19)	0.031610 (6.08)	0.019923 (5.90)
0	0	0	0	0.033325 (15.63)	0.029717 (5.53)	0.019092 (5.58)
0	0	0	0	0	0.053841 (11.19)	0.015307 (4.49)
0	0	0	0	0	0	−0.038032 (−19.50)

Estimate of the symmetric matrix G

0.025830 (0.31)	0.042393 (0.72)	−0.297097 (−4.89)	−0.052299 (−0.84)	0.054694 (1.21)	0.141002 (3.13)	−0.081971 (−1.73)
0.042393	0.284512 (3.31)	−0.208356 (−3.47)	0.077427 (1.55)	−0.052753 (−1.03)	−0.031107 (−1.04)	0.025534 (0.44)
−0.297097	−0.208356	0.193374 (1.92)	−0.145139 (−2.47)	0.187317 (4.41)	−0.065801 (−1.10)	0.113149 (2.58)
−0.052299	0.077427	−0.145139	−0.019253 (−0.20)	0.053159 (1.07)	−0.040474 (−0.62)	−0.068033 (−1.13)
0.054694	−0.052753	0.187317	0.053159	0.197648 (2.68)	−0.194327 (−4.42)	−0.192212 (−4.00)
0.141002	−0.031107	−0.065801	−0.040474	−0.194327	0.030454 (0.30)	0.082286 (1.30)
−0.081971	0.025534	0.113149	−0.068033	−0.192212	0.082286	0.082715 (1.26)

(t-statistics in parentheses)
Sample Size: 175
Degrees of Freedom: 111

Table 10d: 1978(1)-1987(12)

CAPM Estimation: 15 parameter ARCH

$$z_{t+1} = c + \rho\Omega_t\lambda_t + \varepsilon_{t+1}$$

$$Var_t(\varepsilon_{t+1}) = \Omega_t = P'P + G\varepsilon_t\varepsilon_t'G'$$

Log Likelihood: 1742.644641

Estimate of the vector c':

BONDS W.GERMANY	BONDS JAPAN	BONDS UK	EQUITY W.GERMANY	EQUITY JAPAN	EQUITY UK	EQUITY US
-0.008329	-0.006796	-0.007893	-0.011479	-0.005166	-0.007290	-0.007114
(-1.05)	(-0.61)	(-0.85)	(-0.95)	(-0.37)	(-0.49)	(-0.62)

Estimate of the coefficient ρ:
11.423238
(1.31)

Estimate of the upper triangular matrix P:

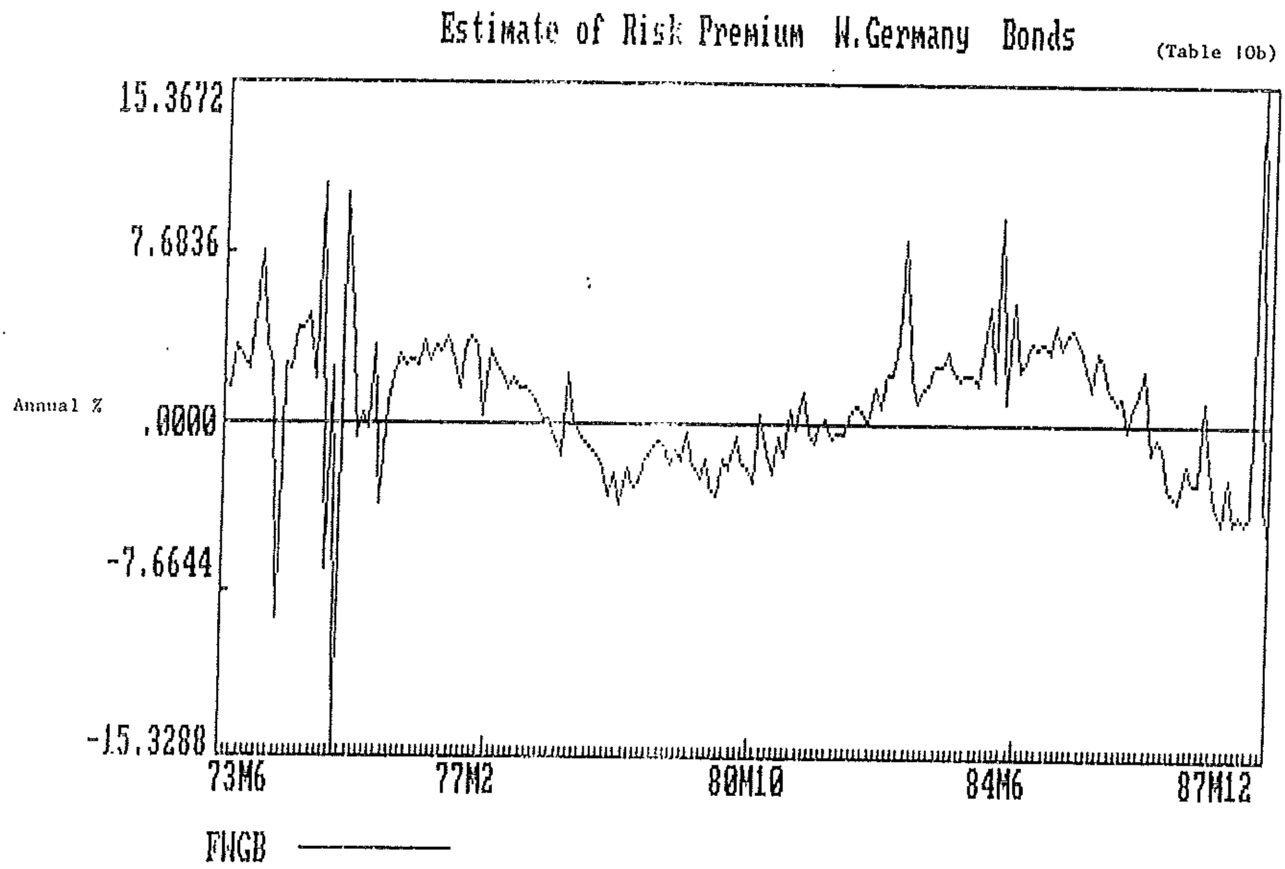
0.045495 (9.22)	0.031538 (3.81)	0.026945 (3.49)	0.037911 (4.55)	0.026822 (2.75)	0.018676 (1.81)	0.003334 (0.46)
0	0.037912 (7.69)	0.010517 (1.57)	0.004245 (0.50)	0.041850 (4.88)	0.014529 (1.56)	0.003037 (0.44)
0	0	0.042468 (9.13)	-0.004982 (-0.68)	-0.001275 (-0.21)	0.036223 (5.67)	0.007934 (1.11)
0	0	0	0.039033 (10.18)	0.012850 (2.45)	0.019116 (2.82)	0.018229 (2.30)
0	0	0	0	0.022033 (7.02)	0.013935 (1.99)	0.007732 (0.94)
0	0	0	0	0	0.034203 (7.02)	0.010459 (1.38)
0	0	0	0	0	0	0.031956 (9.47)

Estimate of the symmetric matrix G

-0.267157 (-1.26)	-0.073162 (-0.55)	0.083008 (0.53)	0.126763 (0.78)	0.357734 (2.64)	-0.086670 (-0.61)	-0.047017 (-0.29)
-0.073162	0.329600 (2.67)	-0.005878 (-0.05)	0.222600 (1.98)	-0.007072 (-0.07)	-0.147509 (-1.20)	0.012041 (0.11)
0.083008	-0.005878	-0.044871 (-0.25)	-0.219247 (-1.57)	0.234991 (1.98)	-0.104292 (-0.74)	0.298441 (2.35)
0.126763	0.222600	-0.219247	-0.175745 (-0.88)	-0.054200 (-0.53)	0.222191 (1.27)	0.442692 (2.53)
0.357734	-0.007072	0.234991	-0.054200	0.353725 (1.91)	-0.035671 (-0.27)	-0.297388 (-1.90)
-0.086670	-0.147509	-0.104292	0.222191	-0.035671	0.337463 (1.97)	0.198167 (1.46)
-0.047017	0.012041	0.298441	0.442692	-0.297388	0.198167	-0.084805 (-0.41)

(t-statistics in parentheses)
Sample Size: 120

FIGURE 1a



Estimate of Risk Premium W.Germany Bonds
(Table 10b)
15.3672
7.6836
.0000
-7.6644
-15.3288
Annual %
73M6
77M2
80M10
84M6
87M12
FWGB

FIGURE 1b

Estimate of Risk Premium U.S. Equities (Table 10b)

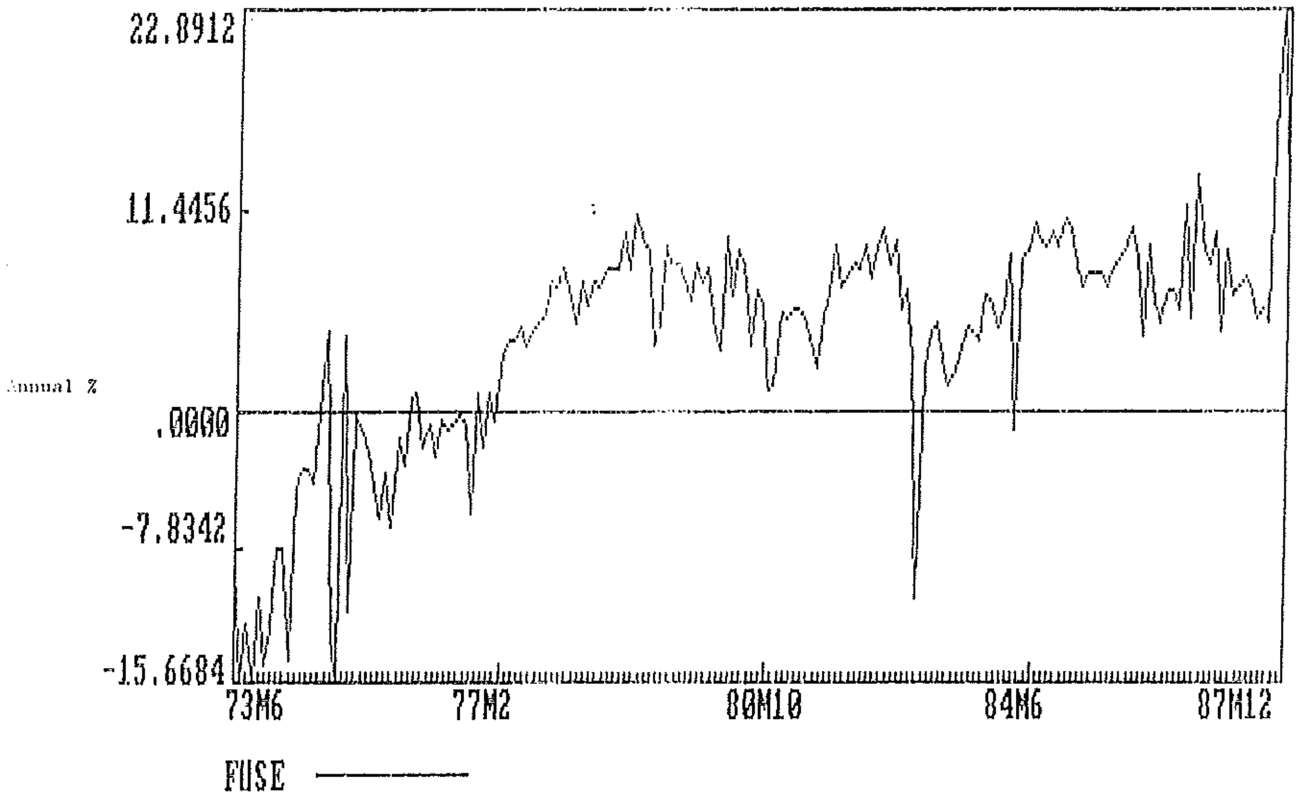


FIGURE 2a

Estimated Risk Premium U.S. Equities (Salomon Data) (Table 10d)

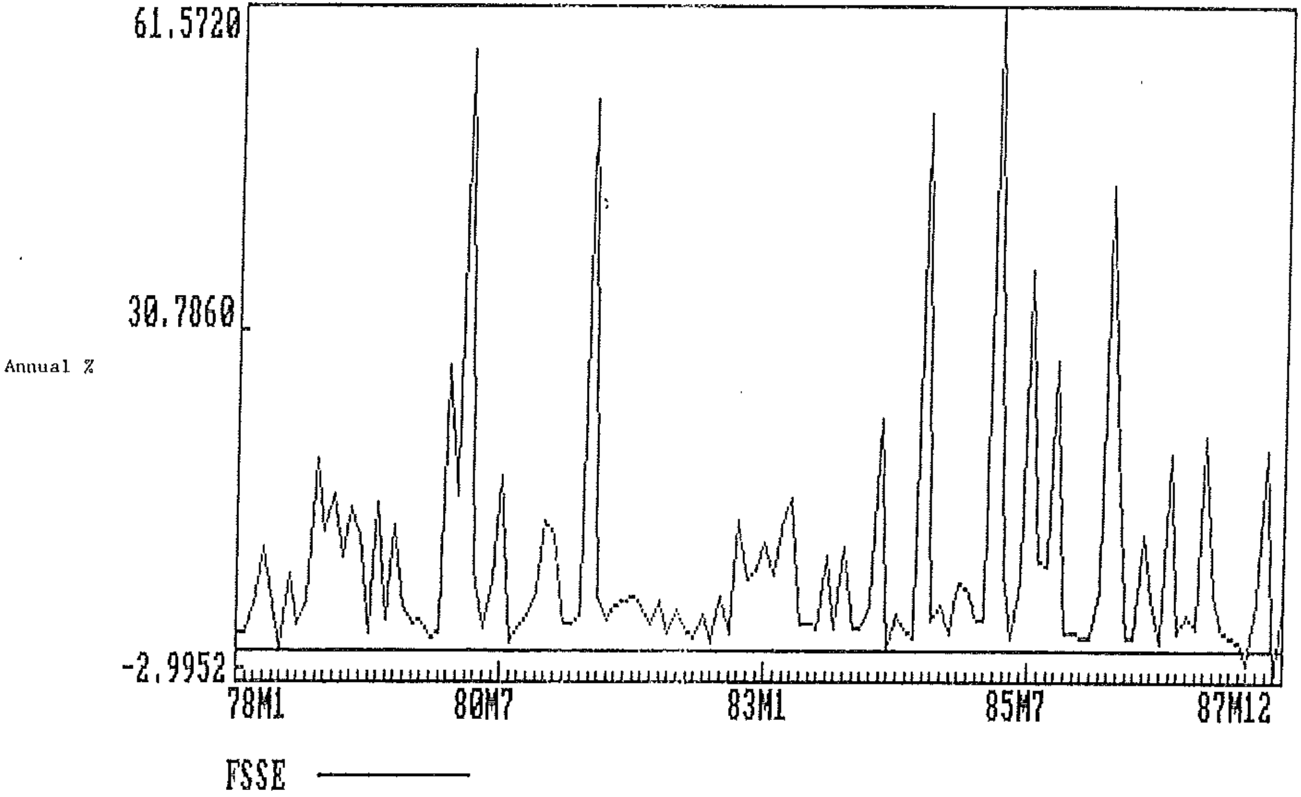


FIGURE 2b

Estimated Risk Premium W.Germany Bonds (Salomon Data) (Table 10d)

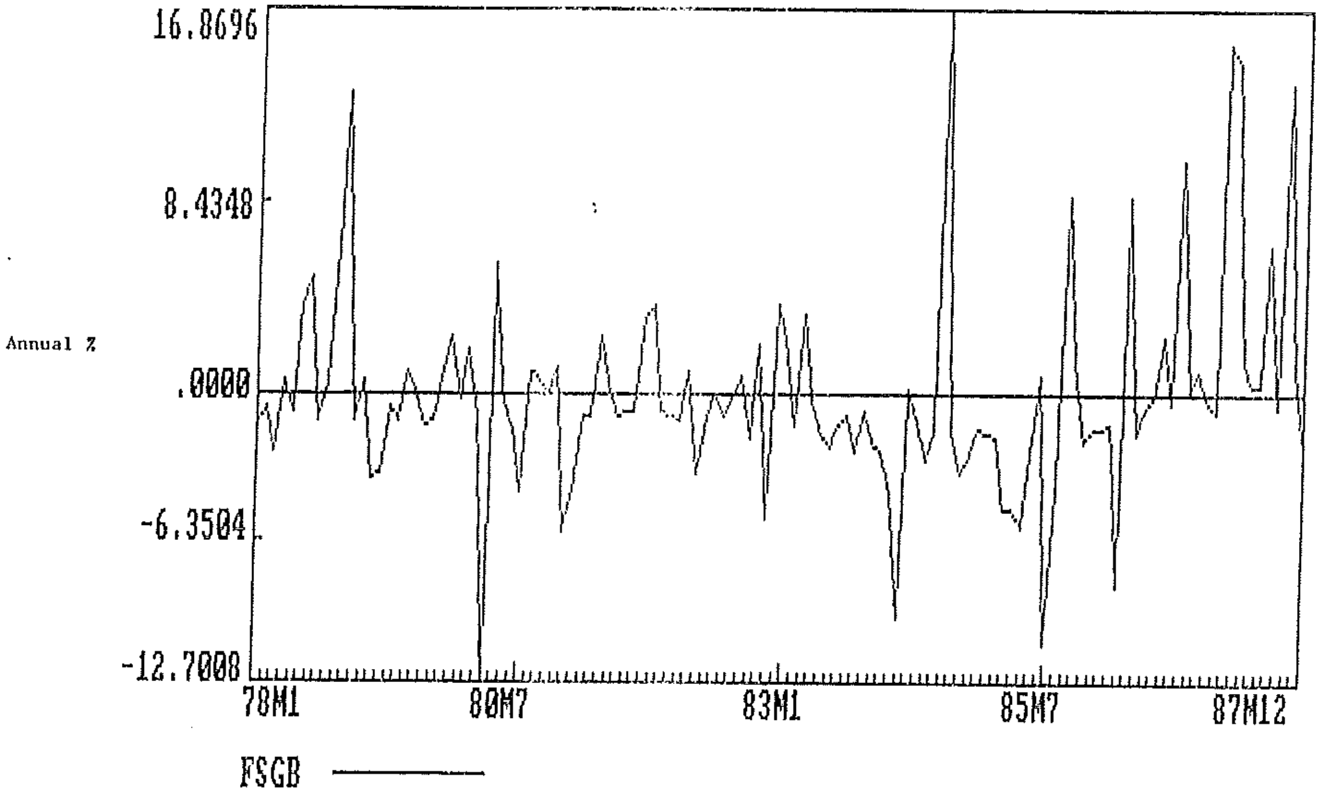


FIGURE 3a

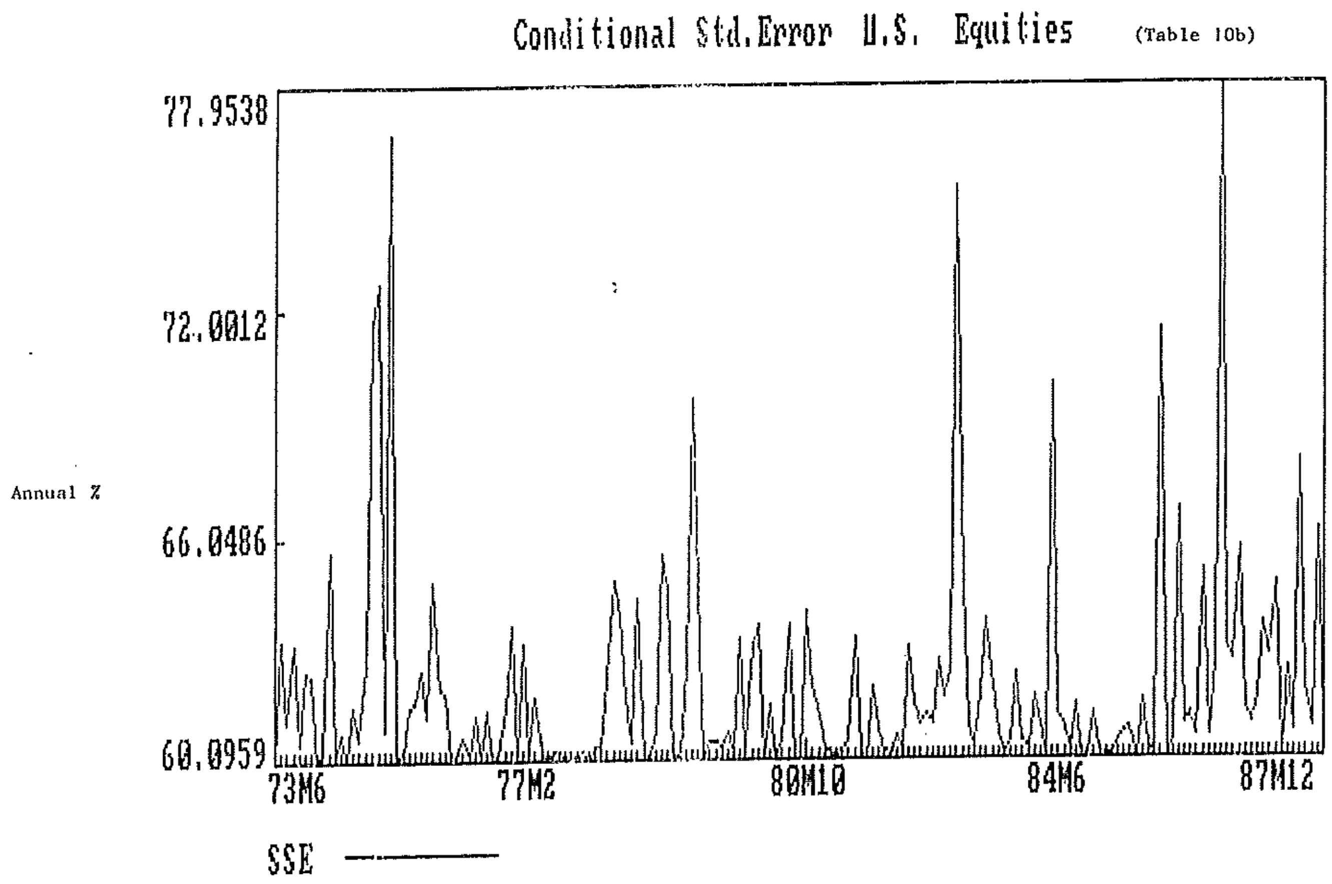
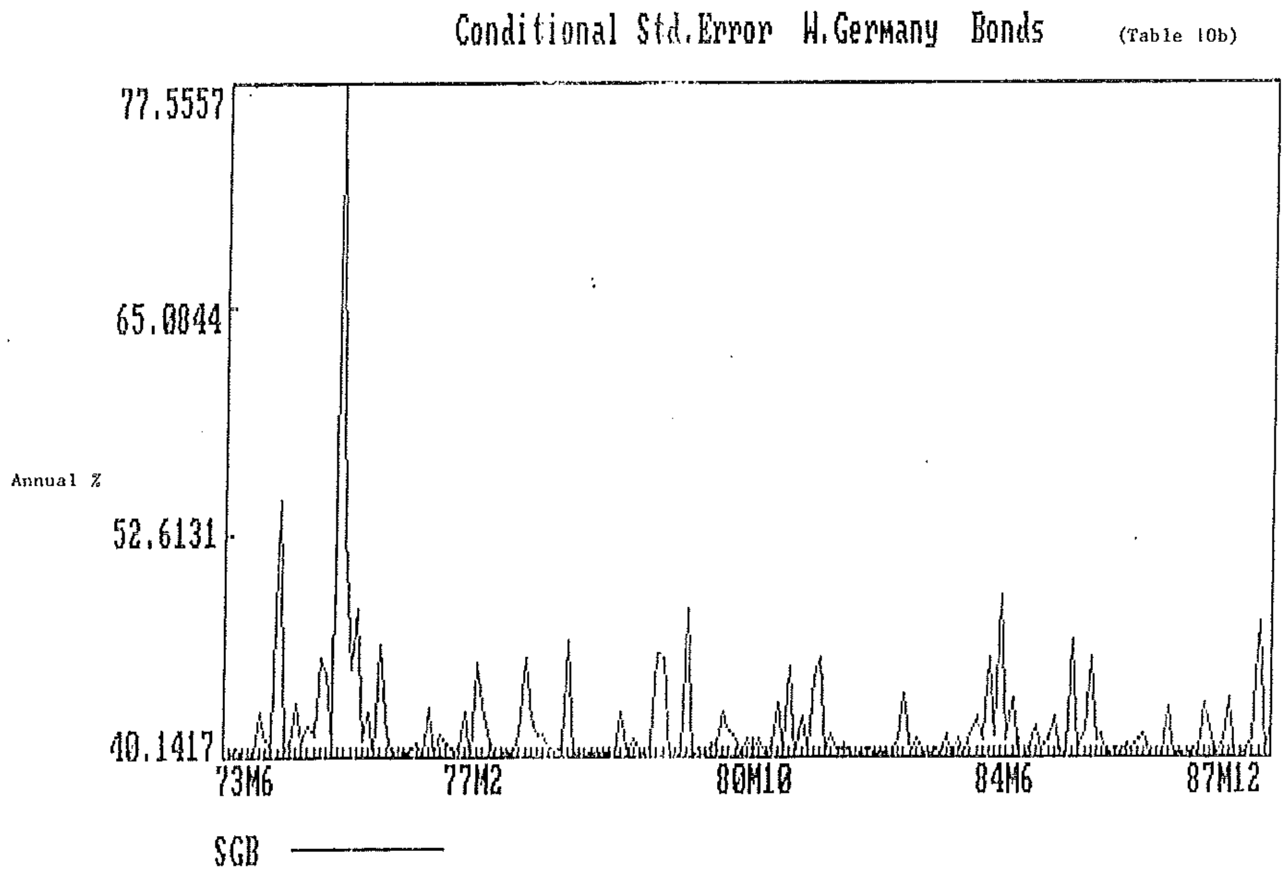

Conditional Std.Error U.S. Equities
(Table 10b)
Annual %
77.9538
72.0012
66.0486
60.0959
73M6
77M2
80M10
84M6
87M12
SSE

FIGURE 3b



Conditional Std.Error W.Germany Bonds
(Table 10b)
77.5557
65.0844
52.6131
40.1417
Annual %
73M6
77M2
80M10
84M6
87M12
SGB

FIGURE 4 a

Conditional Std.Error W.Germany Bonds (Salomon Data) (Table 10d)

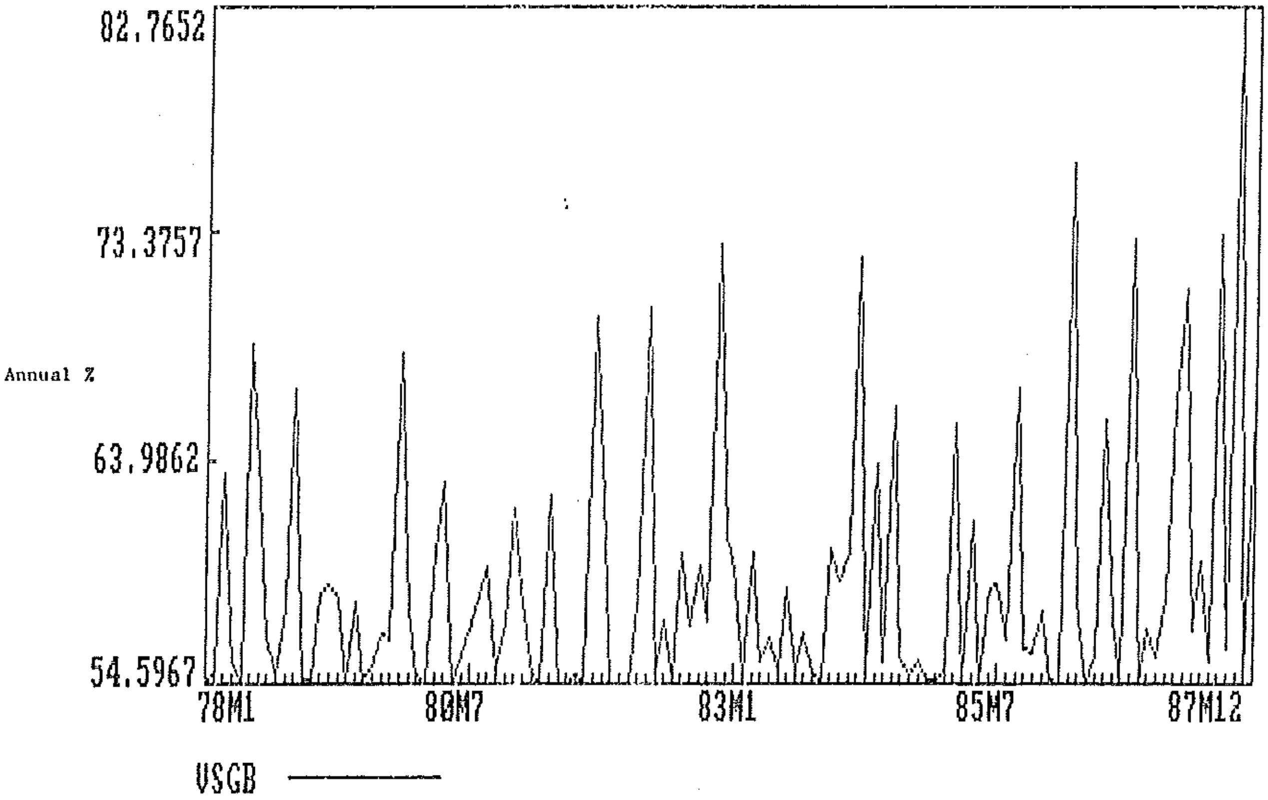


FIGURE 4b

Conditional Std.Error U.S. Equities (Salomon Data) (Table 10d)