Stock Returns and Inflation: Some New Evidence

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Abstract

Using aggregate and industry-wise monthly UK data over a period of 44 years we examine the long run relationship between stock return index (S_t) and retail price index (P_t) in a VAR framework. Univariate tests confirm P_t as I(2); nevertheless pairs of S_t and P_t are co-integrated and share common I(1) trend. There is no evidence of shared I(2) trend. We find evidence of shifts in the co-integrating ranks and parameters, and accounting for these shifts improved estimates' precision. The long run price elasticity of return index is consistently above unity, a finding that stands in sharp contrast to the existing ones. Overall our results suggest that tax-paying stock investors are fully insulated against inflation in the long run.

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I. INTRODUCTION

Fisher hypothesis, in its strict sense, predicts a positive homogeneous relationship of degree one between stock return and inflation[1]. However, the empirical evidence is far from conclusive. Event studies, which look at the inflation announcement effects, report a negative relationship between inflation and stock returns (e.g. Amihud, 1996). Short horizon studies that use monthly or annual data covering typically a period of ten to fifteen years also report either negative or insignificant relationship between stock return and inflation (e.g. Jaffe and Mandelker, 1976). In contrast, long-horizon studies – which estimate stock return-inflation relationship using period average data (e.g. Boudoukh and Richardson, 1993; Boudoukh *et al.*, 1994) and a co-integration framework (e.g. Ely and Robinson, 1997) - find a positive and significant relationship between stock returns and inflation with a price elasticity of less than unity[2]. Hence, the consensus that the Fisher hypothesis holds in the long-run, albeit in a weaker sense in that the null of unit elasticity is widely rejected by the data.

Given that stock return and commodity price indices are both known to be integrated processes, by taking their first difference one may loose the long-run information contained in their levels; hence a motivation for the co-integration analysis. In this paper we examine the long-run relationship between stock return index (S_t) and retail price index (P_t) in a co-intgrating framework using both aggregate and industry-wise data

[1] Fisher hypothesis is formulated on an *ex-ante* basis however we examine the long-run relationship between stock return and retail price indices on an *ex-post* basis, which is equivalent to assuming perfect foresight.

[2] If one were to allow for taxation then the net compensation to inflation received by investors would be even smaller.

obtained from the London Stock Exchange. In so doing we extend the analysis into three important directions.

First, co-integration tests require long-span data in order to preserve the power of the test. However, the long span data is exposed to various policy changes and several economic shocks with potentially serious structural shifts in the relationships. During the last forty years or so the world economy has witnessed events like oil price shocks, the demise of Bretton Woods, the stock market crash of 1987, globalisation etc., which may induce structural shifts in any economic relationship. Hence a natural enquiry would be to examine whether the co-integrating relationship between S_t and P_t is structurally stable. Interestingly, this issue remains un-addressed in the existing empirical literature that tests co-integration between S_t and P_t, which we address.

Second, we evaluate whether a bi-variate VAR of S_t and P_t contains an I(2) component. This is important because stock return indices are known to be first order integrated (I(1)) whereas price indices are widely reported as second order intergated (I(2)). Nelson and Schwert (1977), Baillie(1989) and Johansen(1992), amongst others, find evidence of two unit roots in price levels using different countries and time periods. If I(2) component is present in the system [a realistic possibility if P_t is I(2)] then the usual co-integrating vector, ($\beta' X_t$), becomes an I(1) relationship which invalidates the interpretation of the Fisher relationship as a long-run stationary relationship between S_t and P_t, a typical route taken by the existing standard I(1) co-integration tests. In fact, the presence of the I(2) component may induce a higher order polynomially co-integrated relationship between S_t and P_t, an issue which has remained outside the remit of I(1) analysis[3]. In this paper we evaluate whether systems of S_t and P_t contain an I(2) component and whether they are polynomially co-integrated. Third, an assessment of the long-run relationship between S_t and P_t in a co-integrating framework using industry-wide data is, to our knowledge, non-existent. Boudoukh *et al.*, 1994 emphasize the importance of the Fisher effect on

[3] Haldrup (1994) credits Yoo as the first person who invented the concept of polynomial co-integration.

industry-wide data. Thus, this study will extend the empirical literature to dis-aggregated data.

Our findings are quite unique and interesting. First, we find pairs of S_t and P_t to be co-integrated and the long-run price elasticity of stock returns to be significantly above unity. Of the eight pairs of S_t and P_t analysed, seven show price elasticity significantly above unity and the remaining one shows unit elasticity. This finding is theoretically plausible with the tax- augmented version of the Fisher hypothesis, i.e. nominal stock return must exceed the rate of inflation in order to fully insulate the tax-paying investors. This finding is also unique in the sense that it stands in sharp contrast to the existing estimates, most of which report elasticity of below unity. Second, we find evidence of significant structural shifts in both the co-integrating ranks and the co-integrating parameters in the majority of cases (five out of eight). Accounting for these shifts improved the precision of our estimates. Third, univariate unit root tests confirm P_t to be I(2) which is consistent with the existing literature in the time series tradition. Nevertheless, we do not find evidence of common I(2) trends in the system of S_t and P_t. An implication of this finding is that I(1) analysis is valid for a system consisting of S_t and P_t irrespective of the univariate findings that P_t is I(2). The rest of the paper is organised as follows. Section II discusses data; section III briefly outlines the univariate time series properties of data; section IV discusses tests of co-integration and common trends; and section V concludes.

II. SAMPLE AND DATA SOURCES

The data used in this study are obtained from London Business School Share Price database. The sample covers a period of January 1955 to December 1998 thus giving 528 monthly observations. The Retail Price Index (RPI) is used to measure the inflation. The FT All Share Index (FTA) is a market capitalisation weighted stock return index that we use to represent stock return in the London Stock Exchange [4]. We have constructed seven

[4] Stock return index is inclusive of dividend payments.

equally weighted industry portfolios from individual company returns[5]. The seven industry portfolios thus constructed are Mineral Extraction (MIN), General Manufacturing (MAN), Consumer Goods (CON), Services (SER), Utilities (UTL), Financial Institutions (FIN), and Investment Trusts (INV). In figure 1 we plot log levels of RPI, FTA, MIN and figures 2a and 2b show the rates of growth of RPI and FTA respectively. Since these plots closely mimic the movements of other return indices we do not present their plots in order to conserve space.

Figures 1- 2 about here

Plots reveal that the levels of all series are trending and appear clearly non-stationary whereas their logarithmic first differences appear stationary. An examination of these plots suggests two important shifts in the stock return indices during the sample period. The first one occurred in January 1975 when the FTA return jumped by 53.66 percent and the stock market remained buoyant throughout the February of 1975. Industry-wise stock returns also exhibited remarkable upsurges during this period. This positive growth in stock prices is known as the 'stock market rally' of 1975 in the UK. The second shift is the well-known October 1987 crash. The FTA return plunged by 26%; other industry groups plunged between 22.50% and 29.87%. As argued above, such jumps may induce structural breaks in the long-run relationship between S_t and P_t, and addressing this issue is one of the motivations of this paper.

Preliminary calculations show that during our sample period the average annual inflation has been 6.41 percent while the stock market (FTA) grew at the rate of 16.52 percent per annum. At the industry level the minimum average annual return is 7.71 percent recorded by MIN while the maximum return has been 20.45 percent, recorded by UTL. These growth rates indicate that the overall market return index grew by 2.58 times the inflation; at the industry level this growth has been a factor of 3.19 for the UTL (highest) and 1.20 for MIN (lowest). Thus, a casual observation indicates a positive mean real rate

[5] Boudoukh et al. (1994) also construct equally weighted industry portfolios for the US market while examining this relation; however their empirical approach is different from ours.

of return and hence a positive relationship between inflation and stock returns although the magnitude varies across the industries.

III. UNIVARIATE TIME SERIES PROPERTIES

A. *Unit Root Tests*

The uni-variate time series properties of retail price index (P_t) and stock return index (S_t) are examined through the standard unit root tests and the tests that help identify a structural break endogenously. Specifically, two unit-root tests, viz., the KPSS and the augmented Dickey-Fuller (ADF) tests are implemented first[6]. The ADF test tests the null of unit root whereas the KPSS test tests the null of stationarity. If the KPSS test rejects the null but the ADF test fails to reject, then both tests support the same conclusion, i.e., the series in question is a unit root process.

The ADF and KPSS tests are reported in table 1. As expected, the typical non-rejection of a unit root in all stock return indices is upheld. Considering the stock return indices, the ADF tests cannot reject the unit root null in any of the indices and, at the same time, KPSS tests reject the null of stationarity in all cases. At their first differences, the ADF tests reject the unit root in all the stock returns whereas KPSS tests cannot reject the null of stationarity. Thus, ADF and KPSS tests both confirm that all stock return indices are unit root processes. However, results suggest that P_t is an I(2) process. The ADF test cannot reject the unit root null in inflation (logarithmic first difference of P_t); the KPSS test supports this conclusion by rejecting the stationarity of inflation. Thus, as reported elsewhere, our results confirm that stock return indices are I(1) processes whereas price indices appear to be I(2).

Table 1 about here

B. *Tests of Structural Break*

[6] These tests are well known; nevertheless readers interested for references should consult Dickey and Fuller (1979) and Kwiatkowski et al. (1992).

It is well known that stationary series with structural breaks may appear non-stationary. Failure to allow for structural shifts could bias the unit root tests in favour of non-stationarity. In order to test for the trend break stationarity of S_t and P_t we employ Perron's (1997) sequential unit root tests, which identify the date of structural break endogenously. Consider the following models:

$$y_t = m_1 + q_1 DU_t + b_1 t + d_1 D(T_b)_t + a_1 y_{t-1} + \sum_{i=1}^{k} c_{1i} \Delta y_{t-i} + e_{1t} \qquad (1)$$

$$y_t = m_2 + q_2 DU_t + b_2 t + g_2 DT_t + d_2 D(T_b)_t + a_2 y_{t-1} + \sum_{i=1}^{k} c_{2i} \Delta y_{t-i} + e_{2t} \qquad (2)$$

$$y_t = m_3 + q_3 DU_t + b_3 t + g_3 DT^*{}_t + \tilde{y} \qquad (3a)$$

$$\tilde{y} = a_3 \tilde{y}_{t-1} + \sum c_{3i} \Delta \tilde{y}_{t-i} + e_{3t} \qquad (3b)$$

where y_t is a time series of interest, $DU_t = 1(t>T_b)$; $D(T_b)1(t=T_b+1)$; $DT_t = 1(t>T_b)t$ and T_b denotes the time of change in the trend function. Model (1), the *innovative outlier model,* allows for only a change in the intercept. Model (2) allows for both changes in the intercept and the slope. Model (3), the *additive outlier model*, is estimated in two steps and allows for a change in the slope but both segments of the trend are joined at the time of the break.

The sequential approach involves estimating equations 1-3(b) using the full T observations for each possible break date. Inference on break date is derived in the following three ways. First, select the T_b which minimises the sequential ADF t statistic, *min* t_a= min $t_{ADF}\{T_b \in (k+1, T)\}$ for testing $\alpha_i=1$ (i=1,2,3) over the T-(k-1) regressions (k is the number of parameters). This test is similar to that suggested by Banerjee *et al.* (1992) and Zivot and Andrews (1992) except that Perron shows that there is no need to apply arbitrary trimming at each end of the sample[7]. Second, T_b is chosen such that $t\hat{q}$, the t-statistics on the parameter associated with the change in the intercept (Model 1), or $t\hat{g}$, the t-statistics on the change in slope (Models 2 and 3), are minimised. The test statistic is the

[7] For an application of Banerjee *et al*. (1992) and Zivot and Andrews (1992) tests see, among others, Luintel (2000).

corresponding t-statistic on α obtained under these procedures. Finally, T_b is also identified on the basis of the t-statistics on α corresponding to the maximum of the absolute value of $t\hat{q}$ and $t\hat{g}$. In all cases the no-break unit root null is rejected if the *min* t_a exceeds (in absolute value) the corresponding critical value. The estimated break point corresponds to the date for which the t-statistic is minimised under the null[8]. We follow the data dependent 't-sig' method, which sets statistically suitable k for each regression in the sequence[9]. We set an upper bound, k_{max}, on k =12; if the last lag is significant (at 10%) then the lag length (k) = k_{max}. Otherwise, we reduce k until a significant lag is found. If no lags are found to be significant then k is set to zero, i.e., k=0.

Table 2 about here

Table 2 contains results of sequential unit root tests that identify break dates endogenously. Alternative approaches (based on the minimisation/maximisation of $t\hat{q}$ and $t\hat{g}$) to identify break dates, T_b, produced qualitatively similar results; therefore only those results pertaining to *min* t_a= min $t_{ADF}\{T_b \in (k+1, T)\}$ for testing α_i=1 (i=1,2,3) over the T-(k-1) regressions are reported for all three models (Models: 1-3(b)). The lag lengths (k) are selected following the data dependent 't-sig' approach as suggested by Perron (1997). Results suggest that there is no evidence of structural shift in any of the stock return indices. Sequential tests cannot reject the null of no-break unit root in favour of broken trend stationarity at 10% or better. However, P_t shows a shift in its intercept in July 1973, which is significant at 10%. Interestingly, the statistically determined optimal lag length is identical (nine) for all stock return indices and 12 for the P_t and inflation. The sharp rise of stock return of 1975 and the crash of 1987, discussed in section II, do not appear to be statistically significant turning points. Sequential tests are unable to identify these shifts as significant breaks, which is puzzling but not unique. It has been

[8] It should be noted that trimming at each end of the sample is necessary for models (2) and (3).

[9] Perron (1997) shows that the data dependent method of selecting the lag parameter, k, is superior (stable size and higher power) to a fixed k chosen *a priori*.

shown elsewhere in the literature (Raj, 1992; Luintel, 2000) that visually identified break dates do not necessarily turn out to be the statistically significant. The main points of our results are as follows. Standard unit root tests suggest that all stock return indices are I(1) whereas the retail price index is I(2). Although there are economic events that lead us to believe that structural shifts in stock prices occur, sequential tests do not identify them as significant turning points. We find the retail price index exhibiting structural shift in mean at 10 percent; however the null of the no-break unit root cannot be rejected for inflation. We pursue the issue of structural shift more rigorously in a co-integrating framework in the next section.

IV. CO-INTEGRATION, COMMON TREND AND STABILITY TESTS

A. *Econometric Methodology*

We apply Johansen's (1992a, 1995) multivariate method in order to test for the co-integration and common trend between S_t and P_t. Evidence of common I(2) trend between S_t and P_t would suggest that P_t is I(2). This is because univariate tests (both in the literature as well as our own findings) suggest that S_t is I(1) but P_t is I(2); therefore co-integration and a common I(2) trend between S_t and P_t is possible if and only if P_t is I(2). We examine this issue by conducting a joint test of co-integration and common I(2) trend in pairs of S_t and P_t and assess whether a different approach of polynomial co-integration is required in order to estimate the long-run relationship between them. The presence of an I(2) common stochastic trend would imply that S_t and P_t would be polynomially co-integrated. A system of VAR(p) I(1) variables can be re-parameterised as (see, among others, Johansen, 1992a):

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + ...+ \Gamma_{p-1} \Delta X_{t-p+1} + \Pi X_{t-p} + \varphi D_t + u_t \tag{4}$$

where X_t is a (px1) vector; Γ_i and Π_i are (pxp) coefficient matrices; D_t are usual deterministic components such as seasonal and impulse dummies; μ is a constant term and u_t is a vector of normally and independently distributed error terms. A co-integrated I(1) system implies that: (i) $\Pi = \alpha \beta'$ has reduced rank, r, for r < p; and (ii) $\{\alpha_\perp \Gamma \beta_\perp\}$ has full rank, (p-r), where $\alpha_\perp$ and $\beta_\perp$ are (p x (p-r)) orthogonal matrices to α and β. If conditions (i) and (ii) are satisfied then X_t is an I(1) system and it contains at most r stationary co-integrating vectors and (p-r) common I(1) trends. If however $\{\alpha_\perp\Gamma\beta_\perp\}$ is rank deficient, s_1 < (p-r), that implies X_t is integrated of second order or higher and the VAR will have at most r co-integrating I(1) relationships, s_1 common I(1) trends and s_2 = (p-r-s_1) common I(2) trends. These r co-integrating I(1) relationships can however be transformed into polynomially co-integrated stationary relationships. The rank of $\{\alpha_\perp\Gamma\beta_\perp\}$, conditional on the estimates of r, α and β, is determined by trace tests computed on the basis of the following auxiliary regression (see Johansen, 1995):

$$\alpha_\perp\Delta^2X_t = \alpha_\perp\mu + \alpha_\perp\Gamma_1\,\Delta X_{t-1} + \alpha_\perp\Gamma^*_2\,\Delta^2X_{t-2} + ...+ \alpha_\perp\Gamma^*_{p-2}\Delta^2X_{t-p+2} + \alpha_\perp u_t \qquad (5)$$

Equation (5) is a ((p-r)x(p-r)) system of equations in only first and second differences. The reduced rank regressions of (5) will provide estimates of (p-r) eigenvalues, $\hat{l}_1 > \hat{l}_2 > \ldots > \hat{l}_{p-r}$, which are used to test the null that there are at most s_1 common I(1) trends. The relevant trace statistic is:

$$Q_{r,s_1} = -T\sum_{i=s_1+1}^{p-r}\log(1-\hat{l}_i) \qquad s_1 = 0,1,2,..p\text{-}r\text{-}1. \qquad (6)$$

B. ***Empirical Results***

As a precursor to co-integration tests, we report the results in table 3 obtained from regressing the logarithmic first difference of stock return index on inflation in our data set. This is one of the common approaches followed in the literature while testing the Fisher hypothesis. Our results are broadly consistent with findings reported elsewhere (Chen *et al.*, 1986; James *et al.*, 1985). All estimated coefficients are positive but statistically insignificant at any meaningful level. Given these results, the nature of our data may not be very different from those used by other researchers.

Table 3 about here

In order to estimate the VAR models (4-6) we identify the lag-lengths following Sims' (1980) Likelihood Ratio (LR) tests and multivariate Akaike Information Criteria (AIC). Under the LR tests we began with a maximum lag length (k-max) of 20 and sequentially tested down deleting one VAR-lag at a time until the deleted lags are jointly significant. Information criteria normally choose shorter lag length, which are not always sufficient to flush serial correlation from the VAR residuals. However, it is important to render VAR residuals un-correlated (Johansen, 1992). To circumvent this, we restricted the AIC search between k-max = 20 to k-min = 10. The VAR lengths specified following both methods are reported in table 4. As expected, in all but one case (FTA) the LR test selects longer VAR lengths of either 16 or 17. The AIC selects VAR lengths of 13 or 14. Given these two sets of lag lengths, we estimate VAR through 13 to 17 lag lengths and choose the one that shows no serial correlation in the VAR residuals. Evidence (Cheung and Lai, 1993) suggests that estimates of the Johansen method are more robust to over-parameterisation. Therefore, in the event of more than one VAR length (13-17) producing serially uncorrelated residuals we opted for the longer lag length. The precise VAR lengths thus adopted are also reported in the last column of table 4. Trace statistics in (6) are estimated by incorporating a drift term in the unrestricted part of the model which would allow for a non-zero mean in the I(0) component and linear trend in the I(1) and I(2) components. We preclude those deterministic components that generate quadratic trend in the presence of an I(2) component.

Table 4 about here

Table 5 shows trace tests for the co-integration ranks r and s_1. The joint null hypotheses tested are: (i) H_0: $r=1 \cap s_1=0$; and (ii) H_0: $r=1 \cap s_1=1$. Results in the first row reject the null H_0: $r=1 \cap s_1=0$ for all pairs of S_t and P_t. This implies that the joint hypothesis of co-integration and common I(2) trend between S_t and P_t is rejected. Results in the second row cannot reject the joint null of co-integration and common I(1) trend between pairs of S_t and P_t. Thus, the pairs of stock return and retail price indices analysed here are co-integrated and share common I(1) trend, irrespective of the evidence from univariate unit root tests that P_t is I(2). This seemingly contradictory finding can be attributed to the low power of uni-variate unit root tests. Papell (1997) finds similar results: while employing uni-variate time series tests he cannot reject that P_t is I(2) whereas with more powerful panel unit root tests he finds P_t to be I(1). Thus our joint tests of co-integration and shared trend between S_t and P_t support the findings such as those of Papell (1997) albeit from a different perspective. However, our main concern here is to determine whether S_t and P_t share a common I(2) trend. Our results show that this is not the case, which precludes us from the requirement of estimating polynomial co-integration. Hence, in what follows, we set out to estimate the long-run relationship between S_t and P_t treating the system as I(1).

Table 5 about here

The specification of I(1) system of S_t and P_t is as follows. A constant term is restricted in the co-integrating space which allows for a non-zero mean; the VAR lengths are adopted as before which are reported in the last column of table 4; centred seasonal dummies are used in the unrestricted part of the model in order to account for seasonality. We use this specification as our benchmark model in order to investigate the stability of co-integrating ranks and co-integrating parameters.

Table 6 about here

Table 6 (panel A) contains the trace statistics between pairs of S_t and P_t estimated using the benchmark model. Results show that all but one pair of stock return and retail price indices is co-integrated. Trace tests reject the null of non-co-integration at 5% or better for all pairs except for the Mineral Industry. The return index of the mineral industry does not appear to be co-integrated with the retail price index in this benchmark model; this non-co-integration may however be attributed to the rank and/or the parameter instability (see below). Thus, the weight of evidence is that the pairs of stock return and retail price indices are co-integrated both at aggregate and industry levels. Each loading factor associated with the normalised variable (i.e., stock return indices) is negatively signed and significant which validates the normalisation followed and indicates that an error correcting behaviour ensues when S_t and P_t deviate from their long-run equilibrium value. One interesting aspect of our results is that the estimated long-run price elasticity of stock return is well above unity in all cases. This finding is in sharp contrast to the existing ones, most of which suggest a price elasticity of below unity. In fact, our finding that price elasticity of stock return is above unity is fully consistent with the tax-augmented Fisher hypothesis. However, before we dwell further on this set of results we evaluate their stability.

Stability Tests:

Stability of co-integrating ranks and co-integrating parameters are implemented following Hansen and Johansen (1993, 1998)[10]. The test compares the recursively-computed sub-sample ranks of Π matrix with its full sample estimate. If the recursively computed ranks of Π differ significantly from those of the full sample then that signifies a structural break. Likewise, conditional on the identified number of co-integrating vectors, a stability test of β vector is carried out by imposing the full sample parametric values on recursively computed sub-sample estimates. The LR statistic for these hypotheses is asymptotically $\chi 2$, with pr-r^2 degrees of freedom. Stability tests are carried out in two frameworks: the *Z-model* (which allows for both the short-run and the long-run parameters to vary) and the *R-model* (where short-run parameters are fixed to their full sample values and only the long-run parameters are allowed to vary). In the event of conflicting results, Hansen and Johansen (1998) recommend the *R-model*, which is in fact the relevant model for long-run stability tests.

We analyse the stability of co-integrating ranks and β vector over a period of more than 29 years (August 1968 - December 1998)[11]. This period covers episodes like the UK stock market rise of 1975, oil price shocks of 1973 and 1981, the stock market crash of October 1987 etc., events that may induce structural shifts in the co-integrating relation between S_t and P_t. The seven pairs of S_t and P_t, which are found to be co-integrated using

[10] Quintos (1995) also analyses this issue and derives a testing framework which is similar to the one used here.

[11] We allow the first 175 observations for recursion. Starting with initial 150 or 200 observations does not alter the qualitative nature of our results.

our benchmark model can be assembled into two groups based on the results of stability tests. The first group consists of MAN, INV and UTL, which show no evidence of structural break in the co-integrating rank and co-integrating parameters. The second group includes CON, FTA, FIN and SER, which show shifts in co-integrating ranks and / or parameters.

Plots of scaled LR statistics (LR statistic divided by its 5% critical value) that tests the stability of co-integrating ranks as well as the co-integrating parameters for the selected pairs of S_t and P_t are shown in figures 3-5. Tests are computed using both the Z- and the R-models. Since the test statistics are normalised, empirical values of greater than unity imply rejection of the null of stability and *vice versa*. Two line graphs are plotted in each figure: one tests the null of r=0 and the other of r=1. In order to have a stable co-integrating rank we require the rejection of the null of r=0 but non-rejection of r=1.

Figures 3 – 5 about here

The normalised LR statistics for INV are shown in figure 3. The R-model, which is suitable for evaluating the stability of long-run relationships, consistently rejects the null of non-co-integration (i.e. r=0) but cannot reject that of at least one co-integration (i.e. r=1). Since the scaled LR statistics are always below the unity threshold there is no evidence of structural shift in the co-integrating ranks. Tests also indicate stability of co-integrating parameters. The Z-model provides support to the rank stability; however the co-integrating parameters do show episodes of significant shifts. The parameter instability shown by the Z-model is not unexpected given the short-run volatility of stock markets. Plots of MAN and UTL, the other two pairs that show stable co-integrating ranks and parameters, are qualitatively similar to that of INV and are therefore not shown here to conserve the space[12].

[12] Results on the stability tests of all eight pairs of S_t and P_t analysed in this paper are ready and available on request.

Recursive results for FIN are plotted in figure 4 (panel A) which show episodes of rank instability, whether we use R- or Z-model. The R-model indicates that rank instability occurred around the October crash of 1987 but co-integrating parameters remain constant throughout the sample period. The Z-model corroborates the finding of rank instability but shows parameter instability, an outcome which may be due to the short-run volatility. Recursive results for CON and SER are broadly similar to those shown for FIN and hence are not reported to conserve space[13].

The normalised LR tests for FTA (figure 5; panel A) show stability of co-integrating rank but not the parameters. Parameter instability is evident during 1974-1975, which coincides with the 'stock market rally' of 1975. The Z-model confirms these results except that it shows larger magnitude of parameter instability. The overall conclusions from the benchmark model are as follows: (i) evidence is overwhelming that the pairs of S_t and P_t are co-integrated (seven out of eight cases); (ii) the retail price elasticity of stock return is well above unity; and (iii) structural breaks in co-integrating rank and / or parameters are not uncommon (four out of seven co-integrated pairs).

In order to account for the shifts in the co-integrating relations (the ranks and the parameters) we incorporated shift dummies for the identified break periods. Significant shifts in the co-integrating ranks were found with CON, FIN and SER around October 1987; FTA showed significant breaks in the co-integrating parameters around 1974-75, and the benchmark model did not find co-integration for MIN. In panel B of table 6 we report co-integration tests, which include the impulse (crash) dummy for the crash of 1987 in the unrestricted part of the model for all but the pair of FTA and P_t. The crash dummy takes a value of 1 for the months of October and November of 1987 and zero otherwise. For FTA, the impulse dummy takes a value of 1 for the months of January and February of 1975 and zero otherwise. Although INV, UTL and MAN showed stable long-run (co-integrating) relationships nevertheless we report results including the 1987 impulse dummy in order to assess their sensitivity.

[13] A slight difference to be noted is that CON and SER show a slightly longer period of rank instability compared to FIN while co-integrating parameters appear stable.

Results show that incorporation of the shift dummy improves the results in several respects[14]. First, all trace tests for the null of non-co-integration show huge improvement in their precision. The null is rejected at very high precision indeed (0.01 percent or better). The results back-up a single unique co-integrating vector between the pairs of S_t and P_t found earlier. LM statistics, which test the seventh order serial correlation in the VAR residuals, show the absence of serial correlation at 5% significance level. Second, plots of normalised recursive LR statistics (Panel B of figures 4 and 5) show that both the co-integrating ranks and the co-integrating parameters become stable once impulse dummies are incorporated[15]. Stability of co-integrating ranks and parameters is achieved through all five pairs of S_t and P_t, which had shown structural breaks in the benchmark model. Since recursive plots shown in figures 4-5 closely mimic the patterns of other stock indices they are not reported to conserve space. Third, the pair of MIN and P_t, shown non-co-integrated by the benchmark model, is now highly significantly co-integrated. This implies that their non-co-integration may be attributed to the instability caused by the crash of 1987.

The impulse dummy produces stable co-integrating ranks and co-integrating parameters; the magnitude of the intercept term in the co-integrating space changes but the slope co-efficient remains very close (statistically identical). All loading factors are correctly (negatively) signed and significant which validates the normalisation. The magnitude of the loading factors indicates the speed of adjustment towards equilibrium, which appears to be very similar across industries of around 2.0% each month. This adjustment rate is somewhat slower than expected. It may be that stock prices adjust rather sluggishly vis-à-vis the retail price index compared to other company specific news. We find retail price elasticity of above unity which is, as argued above, in sharp contrast to the existing

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[14] We also tried slope dummies but they did not improve on the instability problem as shown by the impulse dummy. This may indicate that the identified shift was in mean rather than on the price elasticity.

[15] The 0,1 impulse dummy generates a null matrix for the recursive tests which we circumvent by the 1, 2 impulse dummy. In some cases it alters the magnitude of the constant term in a marginal way but the slope parameters remain unaffected.

literature. We report the tests of homogeneity restriction between S_t and P_t in column H-R in table 6. The restriction is rejected in favour of price elasticity of above unity for seven of the eight pairs. For the three pairs (i.e., MAN, INV and UTL) that were structurally stable the incorporation of shift dummies makes hardly any difference to their slope parameters, but the magnitudes of the constant terms are affected. LR tests cannot reject the null of equality of slope parameters across models with and without the impulse dummies. In fact, this pattern is similar across the board. The rejection of homogeneity in favour of price elasticity of above unity is consistent with the tax-augmented Fisher hypothesis.

V. CONCLUSION AND IMPLICATIONS

Utilising the aggregate and industry-wise monthly UK data that spans 44 years we have examined the long-run (co-integrating) relationship between stock return index (S_t) and retail price index (P_t). The economic motivation is to assess whether stock investments hedge inflation using 'one of the oldest and most respected financial models' (Boudoukh and Richardson, 1993; p 1346), well known as the Fisher hypothesis. Hitherto the empirical support for this hypothesis is rather lacklustre (see the discussion in section I). In this paper we extend the existing empirical literature in two important directions. First, we address whether the long-run relationship between pairs of stock return index (S_t) and price index (P_t) remain stable during a sufficiently long sample period. Given various economic shocks (e.g., oil price shocks, demise of Bretton Woods, stock market crash of 1987 etc.) witnessed during the last forty years or so it is possible that this relationship may have been through structural shifts. To our knowledge, existing empirical studies that test the long-run relationship between S_t and P_t by employing co-integration techniques have not addressed this issue.

Second, we test whether S_t and P_t are polynomially co-integrated. This emanates from the fact that univariate tests (including our own results in this paper) indicate that the price index is I (2) and the stock return index is I (1). A system that contains both the I(2) and the I(1) variables may share common I (2) trend. If it does then it invalidates the I(1) analysis. In particular, the estimated co-integrating vectors become non-stationary I(1) processes if the I(2) component is present in the system. This jeopardises the usual interpretation of co-integrating vector as a stationary long-run relationship. Thus a valid inference about the long-run relationship between S_t and P_t requires a joint test of co-integration and common I(2) trend. Interestingly this is yet another unaddressed issue in the literature, which we have addressed in this paper.

Our results are interesting and unique. First, we find overwhelming evidence of co-integration between S_t and P_t. All eight pairs of S_t and P_t we analyse are co-integrated when we allow for the structural shifts in the long-run relationship. The trace tests could not reject the joint null of one co-integration and one common I(1) trend. Thus, pairs of S_t and P_t share one co-integrating relation and one common I(1) trend. We do not find evidence of I(2) common trend although the univariate unit root tests did suggest that P_t was I(2). These seemingly contradictory findings can be attributed to the low power of univariate tests. Our results on the order of integration of P_t echo the findings of Papell (1997). One implication of our finding is that a system consisting of stock return index and retail price index may be treated as a I(1) system even if univariate tests report P_t as I(2).

Second, we find that instability of co-integrating ranks and co-integrating parameters is quite pervasive. Instability of the long-run relationship was evident in five of the eight pairs of S_t and P_t analysed. Accounting for these structural shifts through impulse dummies improved the precision of our estimates. Moreover, allowing for structural shift reversed the non-cointegrating result shown by the benchmark model for the stock return of Mineral Extractions (MIN) into a co-integrating one. Overall it appears that the long-run relationships between S_t and P_t have been through periods of instability and such shifts in the relationship need addressing for a reliable inference. Third, we find the retail price elasticity of stock return to be significantly above unity in seven of the eight cases; the remaining (MIN) also shows unit elasticity. This finding is theoretically plausible in that the nominal stock return must exceed the rate of inflation in order to fully insulate the tax-paying investors. These findings are also unique in that they are in sharp contrast to the existing ones, which largely report an elasticity of below unity. Overall our results suggest that investment in stocks over a long period of time fully insulates a tax paying investor from inflation, a finding hitherto quite hard to come by.

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Table 1. Augmented Dickey-Fuller and KPSS tests

	Levels				First Differences[1]	
	ADFτ_μ	ADFτ_τ	KPSSη_μ	KPSSη_τ	ADFτ_μ	KPSSη_μ
Retail Price Index	-0.689[19]	-2.059[19]	2.721[a]	0.329[a]	-2.222 [18]	1.050[a]
FTALL Share Index	0.419[13]	-2.151[13]	3.832[a]	0.672[a]	-6.393[a] [20]	0.103
Mineral Extraction	-1.579[17]	-1.071[17]	2.912[a]	0.300[a]	-5.139[a] [16]	0.299
Manufacturing	-1.473[17]	-2.054[9]	2.957[a]	0.308[a]	-5.253[a] [16]	0.170
Consumer Goods	-1.656[17]	-1.837[9]	2.986[a]	0.368[a]	-5.375[a] [16]	0.201
Services	-1.484[17]	-2.158[17]	2.970[a]	0.169[b]	-5.351[a] [16]	0.146
Utilities	-1.341[16]	-2.845[16]	3.088[a]	0.318[a]	-5.887[a][15]	0.104
Financial	-1.157[17]	-2.528[13]	2.969[a]	0.178[b]	-5.736[a] [16]	0.082
Investment Trusts	-0.992[13]	-2.357[13]	3.776[a]	0.292[a]	-6.246[a] [20]	0.067

1. ADF tests are conducted setting a lag-length of 20 and testing down as suggested by Campbell & Perron (1991). However, the results are robust to any lag-length of 0-20. The KPSS tests for the first difference series are reported for k = 5. Results remain qualitatively the same as other lag lengths; a & b indicates significant at 1% & 5%. For the ADF tests 'τ_μ' indicates constant term only in the estimating equation whereas 'τ_τ' indicates constant term and linear time trend both inclusive. Likewise ' η_μ' denotes constant term only whereas 'η_τ' denotes both constant term and a linear trend in the KPSS model.

Critical Values:

	ADFτ_μ	ADFτ_τ	KPSSη_μ	KPSSη_τ
1%	-3.83	-3.93	0.739	0.216
5%	-2.86	-3.41	0.463	0.146

Table 2 The Sequential Break tests

	Innovative outlier: Model-1			Change in Intercept & slopes: Model-2			Additive Outlier Model: Model-3		
	T_α	K	T_b	T_α	k	T_b	T_α	K	T_b
Retail Price Index	-4.957	12	73:07	-3.971	12	73:08	-3.079	12	57:04
FTALL Share	-3.841	9	82:05	-4.584	9	76:09	-3.878	9	73:07
Mineral Ext.	-2.984	9	91:03	-3.630	9	86:08	-3.586	9	89:02
Manufacturing	-3.187	9	76:09	-3.839	9	85:06	-3.221	9	98:04
Consumer Goods	-3.085	9	94:01	-4.065	9	85:06	-3.685	9	90:11
Services	-3.257	9	89:06	-3.856	9	84:06	-3.302	9	91:06
Utilities	-3.793	9	72:07	-3.659	9	72:03	-2.892	9	57:11
Financial	-3.610	9	91:02	-4.169	9	84:06	-3.493	9	91:06
Investment Trusts	-3.883	9	82:07	-3.779	9	82:07	-2.936	9	75:11
Δrpi	-3.970	12	80:3	-4.757	12	80:3	-3.79	12	76:2
Critical Values									
100 obs									
5%	-5.10			-5.55			-4.65		
10%	-4.82			-5.25			-4.38		
$\propto$									
5%	-4.80			-5.08			-4.36		
10%	-4.58			-4.82			-4.07		

Results of sequential unit root tests that identify break date endogenously (Perron, 1997). Alternative models produced qualitatively similar results so only those results pertaining to model (1) are reported. Critical values are adopted from Perron (1997). Reported Results for all models (Models 1-3) pertain to the method which selects the break date, T_b, by minimising the sequential ADF-t-statistic, $min\ t_a$= min $t_{ADF}\{T_b \in (k+1,\ T)\}$ for testing the null: α_i=1 (i=1,2,3). Other T_b selection methods produced qualitatively similar results and hence are not reported (for details please refer to the text).

Table 3. Stock Return and Inflation (OLS Estimation)

Indices	Constant	T-Stat	Inflation	T-Stat	R-barSQ	DW
FTALL Share Index (FTALL)	0.007	2.44	1.022	1.65	0.012	1.75
Mineral Extraction (Min)	0.001	0.37	0.945	1.77	0.008	1.66
Manufacturing (Man)	0.005	1.70	0.717	1.65	0.007	1.37
Consumer Goods (Con)	0.006	2.77	0.408	1.18	0.002	1.33
Services (Ser)	0.008	3.08	0.508	1.30	0.003	1.34
Utilities (Utl)	0.017	3.80	-0.222	-0.41	0.000	1.87
Financial (Fin)	0.007	2.60	0.501	1.07	0.002	1.51
Investment Trusts (Inv)	0.008	2.80	0.593	1.17	0.003	1.53

The sample period is January 1955 to December 1998. T-ratios are heteroscedasticity consistent. The estimates based on the Robust regression procedure that controls for the effects of outliers are qualitatively similar to those reported here with the exception of the Mineral Extraction industry which resumes a positive coefficient which is not significantly different from unity. As an alternate specification we included impulse dummy in order to allow for the crash of October 1987; however results remain the same.

Table 4: LR Statistics & AIC for VAR Lengths Specification

	LR: (T-c)(log $\|\Sigma_R\|$ - log $\|\Sigma_U\|$)		AIC: T Log $\|\Sigma\|$ + 2N	Lags adopted[1]
FT-ALL	13	χ^2 (28) = 49.40[0.007]	14	13
Consumption	17	χ^2 (12) = 21.19[0.047]	13	17
Financial	17	χ^2 (12) = 26.34[0.009]	13	16
Manufacturing	16	χ^2 (16) = 26.80[0.044]	13	15
Investment	17	χ^2 (12) = 23.66[0.023]	13	16
Mineral	17	χ^2 (12) = 28.63[0.005]	13	17
Services	17	χ^2 (12) = 23.45[0.024]	13	17
Utility	16	χ^2 (16) = 30.08[0.017]	13	14

A VAR length of twenty (20) is specified as the most general model. LR test statistics and AIC are computed by sequentially reducing one VAR length at a time.The Multivariate Generalisation of AIC is give by: AIC = T Log $|\Sigma|$ + 2N; where $|\Sigma|$ is the determinant of variance/covariance matrix of the residuals and N is the total number of parameters estimated in all equations. For example, if each equation in an n-variable VAR has p lags and an intercept then: $N=n^2p+n$; each of the regressors has np lagged regressors and an intercept. Sims' (1980) likelihood ratio test is given by: LR = (T-c)(log $|\Sigma_R|$ - log $|\Sigma_U|$) ~ χ^2 (R).

1. Adopted lag lengths as discussed in the test.

Table: 5 Trace Tests for the Cointegration Ranks(r, s_1)

r	S_1	S_2	CON	FTA	FIN	MAN	INV	MIN	SER	UTL
1	0	1	12.00[c] (12.9)	15.33[b] (12.9)	13.01[b] (12.9)	11.15[c] (12.9)	14.11[b] (12.9)	28.56[a] (12.9)	12.52[c] (12.9)	14.36[b] (12.9)
1	1	0	2.11 (3.8)	0.00 (3.8)	0.52 (3.8)	0.90 (3.8)	0.31 (3.8)	1.41 (3.8)	1.33 (3.8)	0.51 (3.8)

r= number of co-integrating vectors; s_1 = number of common I (1) trends; and s_2 = number of common I(2) trends. The null is defined in the first column vis-à-vis r and s_1. For example, in the first row the null is H0: r = 1 and s_1=0. Likewise, in the last row the null tested is H0: r=1 and s_1=1. Note that $s_2 = p-r-s_1$. The 5% critical values of Trace Statistics are reported within parentheses. a, b, and c indicate significance at 1%, 5% and 10%, respectively. For the adopted VAR length refer to table 4.

Table 6 : Co-integration Tests

Panel A: Tests without Dummy				Panel b: Tests inclusive of impulse Dummy					LM
Trace Statistics H0: rank=r				Trace Statistics					
r=0	r≤1	Co-integrating vectors	α	r=0	r≤1	Co-integrating Vectors	α	H-R	
22.56^{b}	8.63	CON_t = -0.664 + 1.554^{a} rpi (0.451) (0.083)	-0.015^{a} (0.006)	54.47^{a}	6.06	CON_t = 11.344 +1.496^{a} rpi_t (1.917) (0.116)	-0.016^{a} (0.002)	0.049	0.246
20.43^{b}	4.76	FTA_t = -1.957 + 2.191^{a} rpi (1.065) (0.204)	-0.010^{a} (0.003)	30.42^{a}	0.08	FTA_t = -7.527^{a} + 2.226^{a} rpi_t (2.329) (0.142)	-0.018$_{a}$ (0.003)	0.000	0.816
21.5^{b}	8.72	FIN_t = -0.347 + 1.608^{a} rpi (0.594) (0.111)	-0.022^{a} (0.006)	45.58^{a}	3.61	FIN_t= 12.269^{a} +1.631^{a}rpi_t (2.335) (0.140)	-0.017$_{a}$ (0.003)	0.023	0.082
20.17^{b}	6.08	MAN_t = -0.012 +1.480^{a} rpi (0.521) (0.096)	-0.021^{a} (0.006)	48.40^{a}	2.66	MAN_t= 11.319^{a} +1.488^{a}rpi_t (1.991) (0.119)	-0.016$_{a}$ (0.003)	0.035	0.400
27.95^{b}	6.08	INV_t = 0.561 + 1.618^{a} rpi (0.669) (0.125)	-0.018^{a} (0.004)	34.57^{a}	0.67	INV_t = 7.763^{a} +1.671^{a}rpi_t (1.862) (0.113)	-0.020$_{a}$ (0.003)	0.002	0.308
15.42	4.39	MIN_t = -	-	40.30^{a}	6.07	MIN_t= 16.473^{a} + 1.158^{a} rpi (2.940) (0.178)	-0.016$_{a}$ (0.002)	0.506	0.423
23.01^{b}	9.16	SER_t = -0.695 + 1.697^{a} rpi (0.578) (0.106)	-0.017^{a} (0.005)	56.55^{a}	4.15	SER_t = 13.442 + 1.699^{a}rpi_t (2.260) (0.136)	-0.014$_{a}$ (0.002)	0.019	0.137
23.02^{b}	5.19	UTL = 1.468 + 1.953^{a} rpi (1.553) (0.286)	-0.010^{a} (0.003)	26.19^{a}	0.89	UTL_t = -12.706 +2.109^{a} rpi_t (3.502) (0.214)	-0.015$_{a}$ (0.0031)	0.009	0.825

Finite sample corrections, as suggested by Reimers (1992), are implemented to the reported trace statistics. The 1% and 5% critical values for the trace statistics are 24.60 and 19.96 for H_0: r=0; they are 12.97 and 9.24 for r≤1. The lag lengths used are reported in table 3. The column H-R refers to the tests of homogeneity restriction between S_t and P_t which are $\chi^2(1)$ distributed; p-values are reported for this test. Standard errors are reported within the parentheses of loading factors and co-integrating parameters. LM tests refer to the F version of Lagrange Multiplier tests under the null that VAR residuals are uncorrelated. Serial correlation of seventh order is tested for the VAR residuals.

Figure: 1
Natural Log of Retail Price Index and Stock Return Indices

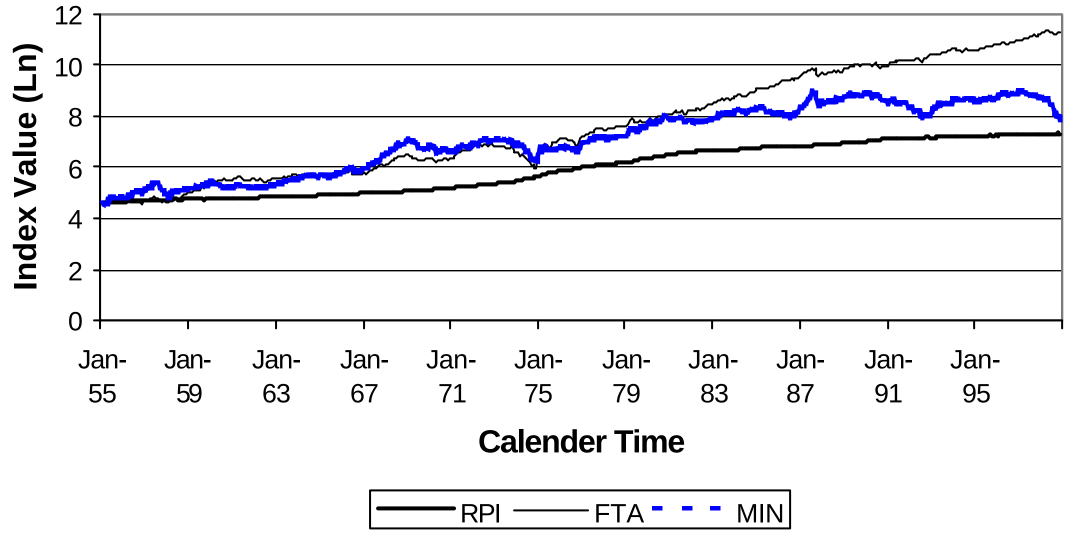


Figure: 2a
Rate of Inflation (%)

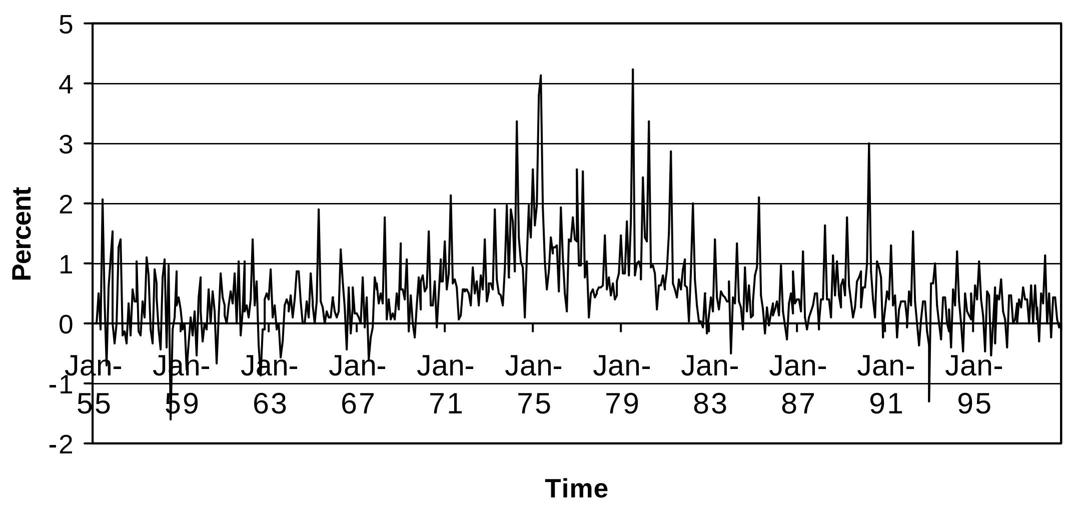


Figure: 2b
Stock Returns (FTA) (in %)

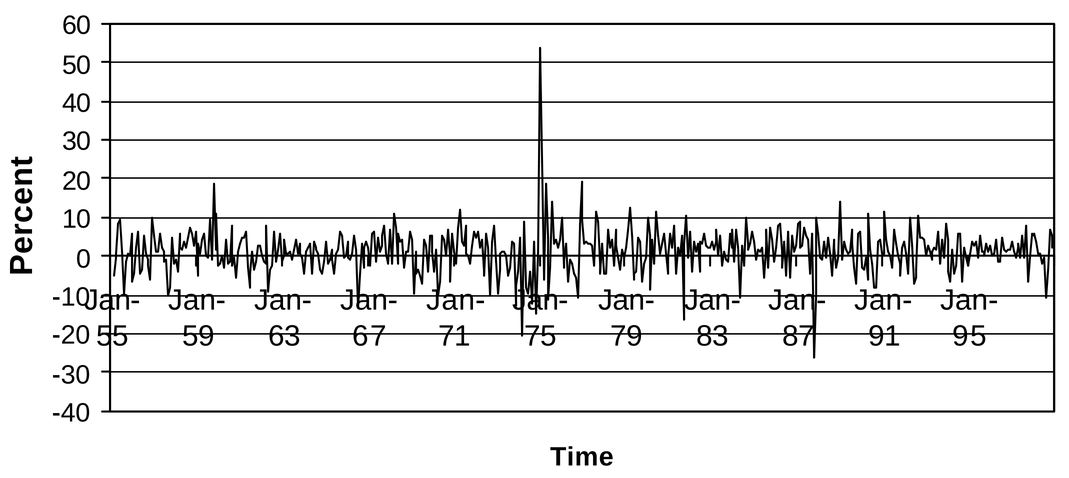
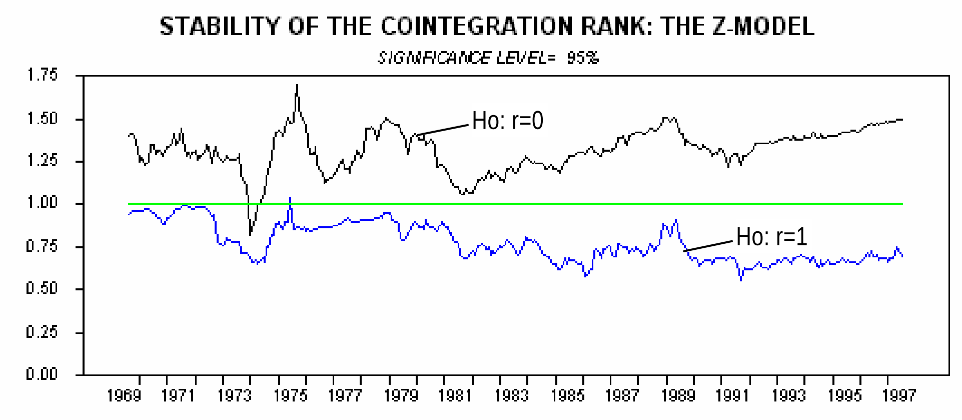
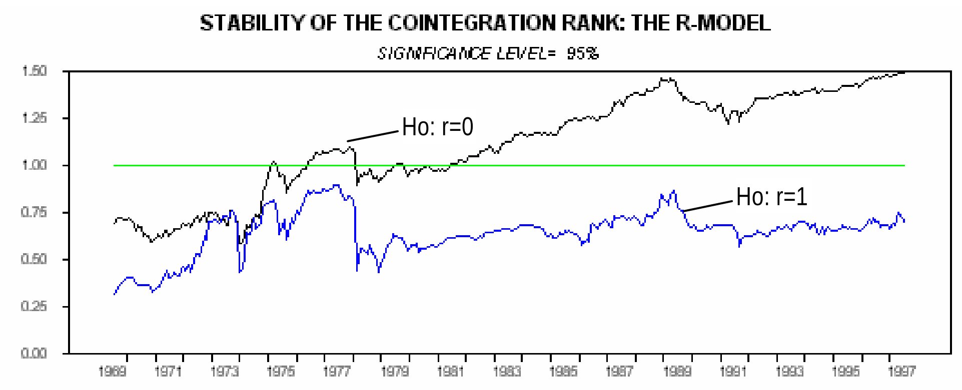


Figure: 3
Investment Trusts

Recursive normalised LR statistics for Investment Trust industry (INV) are depicted in the graphs below. Stability of cointegration rank is tested using the Z and R models.





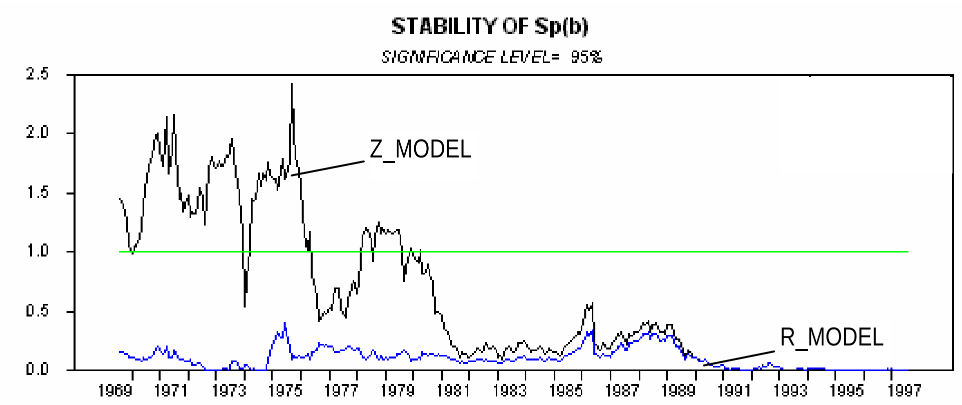


Figure: 4
Financial Sector

Recursive results of the financial sector (FIN) are depicted in the diagrams below. Stability of cointegration rank is tested using the Z and R models. The plots in panel A indicate that rank instability occurred around the stock market crash of 1987. Diagrams in Panel B show the results that include dummy to represent the stock market crash of 1987.

Panel: A Panel: B

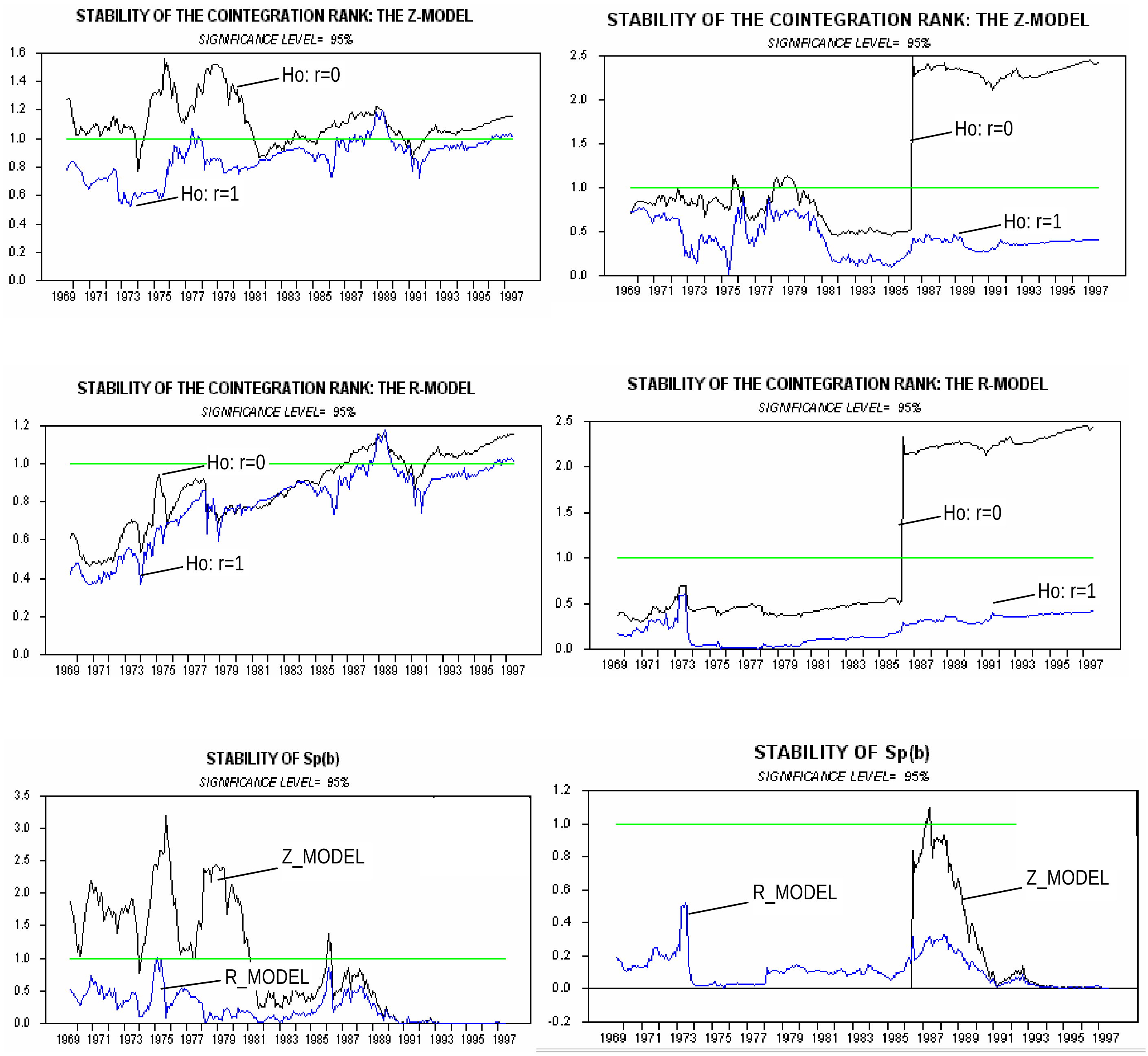


Figure: 5
FT-All Share Index

Recursive results of the FT-All Share Return Index (FTA) are depicted in the diagrams below. Stability of cointegration rank is tested using the Z and R models. The plots in panel A show rank stability, however parameters appear unstable around 1975. Panel B of the figure plots results that include impulse dummies to capture the stock market rally of 1975.

Panel: A Panel: B