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ABSTRACT

In this paper, we use the common structural break test suggested by Bai *et al*. (1998) to test for a common structural break in the stock prices of the US, the UK, and Japan. On the basis of the structural break, we divide each countries stock price series into sub-samples and investigate whether or not the structural break had slowed down the growth of stock markets. Our main findings are that when stock markets are modeled in a trivariate sense the common structural break turns out to be 1990:02, with the confidence interval including several episodes, such as the asset price bubble when housing prices and stock prices in Japan reached a peak in 1988/1989, the early 1990s recession in the UK, the business cycle peak of July 1990, the August 1990 Iraqi invasion of Kuwait and the March 1991 business cycle trough. Annual average growth rates suggest that the structural break has slowed down the growth rate of the UK and Japanese stock markets, while it has boosted the growth of the US stock market.

Keywords: Common Structural Break Test, Stock Markets

JEL: C22, G14, G15

1.   Introduction

There is a substantial literature (see, *inter alia*, Fama and French, 1988a, 1988b; Lo and MacKinlay, 1988; Poterba and Summers, 1988; Kim *et al*., 1991; Chaudhuri and Wu, 2003; Buguk and Brorsen, 2003; Richards 1995, 1997 and Balvers *et al*., 2000) that examines the behaviour of stock prices. This paper is an extension of this research agenda but differs from the extant literature in four ways: (a) it examines confidence intervals for the break date when stock prices in the USA, the UK and Japan are considered individually, (b) it examines whether or not the break in stock prices of different countries occurred at the same time, (c) it examines the interval estimate of the break date, when the date is modeled as common across the three countries, and (d) on the basis of the common structural break, it divides each countries stock price series into sub-samples and calculates annual average growth rates with the aim of investigation whether or not structural breaks have slowed down growth of stock markets.

The aims of this paper are achieved by using a technique for constructing asymptotically valid confidence intervals for the date of a single break in multivariate time series developed by Bai *et al*. (1998). There are two key advantages of using this technique: (1) an interval estimate of the break date, by virtue of providing sample certainty, is more useful in understanding the importance of shocks that create such breaks; and (2) there are many factors that may be crucial in explaining the existence of breaks.

In the case of stock markets, there are several episodes, such as the stock market crash in 1987, the oil crises in the 1970s, the Asian Financial crisis in 1997, among others, that

could result in breaks been simultaneous across countries; see Section 2 for a detailed discussion drawing on contagion effect. As a result, Bai *et al*. (1998) observe that gains in precision can be obtained by a multivariate treatment, where variables are modeled as breaking contemporaneously across series.

Briefly foreshadowing our main results, we find that when stock markets are modeled in a trivariate sense the common structural break turns out to be 1990:02. The associated confidence interval for this common break date includes several episodes, such as the asset price bubble when housing prices and stock prices in Japan reached a peak in 1988/1989, the early 1990s recession in the UK, the business cycle peak of July 1990, the August 1990 Iraqi invasion of Kuwait and the March 1991 business cycle trough. Our analysis of the annual average growth rates suggest that the structural break has slowed down the growth rate of the UK and Japanese stock markets, while it has boosted the growth of the US stock market.

The rest of the paper is organized as follows. Section 2 provides the motivation for the empirical analysis conducted in this paper. In this section, we draw on the "contagion effects" literature to provide an overview of how contagion effects can lead to a common structural break among stock markets. Section 3 includes a discussion of the methodology. Section 4 entails the empirical results, and the final section provides some concluding remarks.

2.   Contagion effect as a cause for a common structural break

In this section, we discuss the contagion effect that can be perceived as leading to a common structural break. Pericoli and Sbracia (2003: 574-575) explain the five definitions of contagion, and these include: (1) Contagion is a significant increase in the probability of a crisis in one country, conditional on a crisis occurring in another country; (2) Contagion occurs when volatility of asset process spills over from the crisis country to other countries; (3) Contagion occurs when cross-country comovements of asset prices cannot be explained by fundamentals; (4) Contagion is a significant increase in comovements of prices and quantities across markets, conditional on a crisis occurring in one market or group of markets; and (5) (Shift-) contagion occurs when the transmission channel intensifies or, more generally, changes after a shock in one market.

In general, however, the literature takes two lines of interpretation on the contagion effect. The first view originates from the fact that market economies are interdependent, which has been accentuated by trade liberalisations at the global scale, leading to macroeconomic similarities or dissimilarities, which create avenues for international trade as countries identify their areas of comparative advantage. Such integration of economies creates opportunities for offshore investment. Interdependence of this sort can lead to co-movement in financial asset prices, and is often referred to as "fundamentals-based contagion".

Several other factors, such as recessions or booms and oil price shocks, can trigger this kind of co-movement. It follows that, and as shown in Calvo and Reinhart (1996) shocks

regardless of whether they are of a global or local nature, are transmitted across countries through real and financial linkages. If countries share common or similar macroeconomic conditions, then a crisis, or shock, may spread contagiously among countries. Moser (2003: 159) explains this point more clearly by stating, "... several countries are hit by a common global or regional external shock ... . Candidates for such adverse common shocks with the potential for inflicting balance of payment difficulties, particularly in emerging market economies, are changes in global (US) interest rates, exchange rates between major currencies, commodity prices, or recessions in major industrial countries".

The second line of interpretation perceives contagion emerging from financial crises, which arise not from macroeconomic fundamentals but from the behaviour of investors or other financial agents. This is often referred to as "irrational contagion", associated with financial panic, herd behaviour, loss of confidence in increases in risk aversion (Karolyi, 2004).

Liquidity and other constraints on lenders or investors can be used to explain individual rational behaviour. If banks from a common creditor country, in the face of deteriorating quality of their loans, reduce the overall risk of the loan portfolio, the liquidity problems and the incidence of financial contagion might spread to those countries whose financial assets are widely traded in global markets and whose markets are more liquid (Karolyi, 2004; Kodres and Pritsker, 2002; Goldfajn and Vades, 1997). It follows that when financial institutions face a default in one country, they tend to withdraw capital not only from that country, but also from other countries so that they avoid further decline in their

asset values (see Kaminsky and Reinhart, 2000). This behaviour is commonly referred to as the common creditor hypothesis.

3. Methodology and theoretical model

This section draws heavily on the work of Bai *et al*. (1998), who consider the model which describes the systems of equations as

$$y_t = \mu + \sum_{j=1}^{p} A_j y_{t-j} + \Gamma X_{t-1} + d_t(k)\left(\lambda + \sum_{j=1}^{p} B_j y_{t-1} + \Pi X_{t-1}\right) + \varepsilon_t \tag{1}$$

where y_t, μ, λ, and ε_t are $n \times 1$ and $\{A_j\}$ and $\{B_j\}$ are $n \times n$; $d_t(k) = 0$ for $t \le k$ and $d_t(k) = 1$ for $t > k$; and X_t is a matrix of stationary variables. From equation (1), assuming that only a subset of coefficients such as the intercept has a possible break and because tests based on a partial model have more power than a full structural change model, Bai *et al*. (1998) derive the former in its stacked form as follows:

$$y_t = (V_t' \otimes I)\theta + d_t(k)(V_t' \otimes I)S'S\delta + \varepsilon_t \tag{2}$$

Here $V' = (1, y_{t-1}', ..., y_{t-p}', X_{t-1}')_t$, $\theta = \text{Vec}(\mu, A_1, ..., A_p, \Gamma)$, $\delta = \text{Vec}(\lambda, B_1, ..., B_p, \Pi)$, I is an $n \times n$ identity matrix, and S, whose rank is equal to the number of coefficients that are allowed to change, is a selection matrix containing 0's and 1's. Equation (2) can be written more compactly as follows:

$$y_t = Z_t'(k)\beta + \varepsilon_t$$

Here $Z'(k)_t = ((V_t' \otimes I), d_t(k)(V_t' \otimes I)S')$ and $\beta = (\theta', (S\delta)')'$. Bai *et al.* (1998) apply the sequence of F-statistics to test for a break in the coefficients: $S\delta = 0$ for $k = k_* + 1, ..., T - k_*$, where k_* represents some trimming value. Hence, the null hypothesis is that there is no structural break. For a given k, the estimator of $\hat{\beta}(k)$ is:

$$\hat{\beta}(k) = \left\{ \sum_{t=1}^{T} \left(Z_t \sum_{k}^{-1} Z_t' \right) \right\}^{-1} \sum_{t=1}^{T} Z_t \sum_{k}^{-1} y_t$$

and the F-statistic testing $S\delta = 0$ is:

$$\hat{F}(k) = T \left\{ R \hat{\beta}(k) \right\}' \left\{ R \left(T^{-1} \sum_{t=1}^{T} Z_t \sum_{k}^{-1} Z_t' \right)^{-1} R' \right\}^{-1} \left\{ R \hat{\beta}(k) \right\}$$

To test for a break, Bai *et al.* (1998) use two tests – the maximum Wald statistic and the logarithm of exponential Wald statistic which, respectively, have the form

$Sup - W:$ $\displaystyle \sup_{\tau \in (\tau*, 1-\tau*)} F_T(\tau),$

$Exp - W:$ $\displaystyle \ln \left\{ \int_{\tau*}^{1-\tau*} \exp \left\{ \frac{1}{2} F_T(\tau) \right\} d\tau \right\}.$

In these test statistics, τ is the trimming region.

There is a growing literature (see *inter alia*, Masih and Masih, 1997, 1999, 2002; Fernandez-Serrano and Sosvilla-Rivero, 2001) that examines the interdependence of stock markets using cointegration analysis of stock price indices of two or more countries. A finding of cointegration is taken as evidence in favour of stock market interdependence because it indicates a common force, such as arbitrage activity, which

brings the stock markets together in the long run. Therefore, testing for cointegration is tantamount to a test of the level of arbitrage activity in the long-run.

In theory, if stock markets are not cointegrated, this implies that arbitrage activity to bring the markets together in the long-run is zero (Masih and Masih, 1997, 1999, 2002). Given the theoretical and practical implications of testing for cointegration of stock markets, we investigate evidence for a cointegration relationship between the stock markets of the US, the UK and Japan.

Indeed the main goal of this exercise, as explained earlier, is to search for a common break in these stock markets. To achieve this, a test for cointegration is necessary since Bai *et al*. (1998) propose a test for the null hypothesis of a structural break in cointegrated models. For a cointegrated model, Bai *et al*. (1998) show that Equation (1) can be written in a vector error correction model (VECM) as follows:

$$\Delta Y_t = \mu + \lambda d_t(k_0) + A(L)\Delta Y_{t-1} + \gamma \alpha' Y_{t-1} + \varepsilon_t \qquad (3)$$

where Y_t can be perceived as the stock price series and $X_t = \alpha' Y_{t-1}$. Because our model encompasses three variables, Bai *et al*. (1998) show that the following triangular representation can be derived:

$$\Delta Y_t = \mu + \lambda d_t(k_0) + F(L)\Delta Y_{t-1} + B(L)(\alpha' Y_{t-1}) + \omega_t \qquad (4)$$

where $\mu = \{I - F(1)\}(D'D\tilde{\mu} - \tilde{B}(1)\alpha'\phi)$, $B(L) = \{I - F(L)L\}\tilde{B}(L)$, $\lambda = \{I - F(1)\}D'D\tilde{\lambda}$ and

$D_t(k) = F^*(L)D'D\Delta d_t(k)$, where $F_j^* = -\sum_{i=j}^{\infty} F_i$.

4.   Empirical results

4.1.   *Data and preliminary analysis of stock indices*

In this paper, we study the stock market price indices of the USA, the UK and Japan using monthly data spanning 1964:06 to 2003:04. The data are obtained from the *OECD Main Economic Indicators*. For the USA, we use the NYSE Common Stocks; for the UK, we use FT-SE-A Non-Financials; and for Japan, we use TSE TOPIX.

The stock price index for each of the three countries is plotted in Figure 1. Three features are worth noting here. First, stock price index rose, in general, in the post-1985 period. Second, a boom in stock prices occurred around the late 1980s to early 1990s period. We later perform the Lumsdaine and Papell (1997) endogenous structural break test to identify formally the exact break dates. Third, the behaviour of the USA and the UK stock price indices have followed a similar pattern throughout the entire period, and beginning in 1998 the Japanese stock market index has experienced a similar pattern of movement.

INSERT FIGURE 1

We explore the relationship among the stock price indices further through comparing the cyclical components of the three indices. We use the Hodrick-Prescott (1997) filter to extract the cycles. The cyclical components of stock price indices are plotted in Figure 2.

9

We can make two observations on the cyclical components. First, the cyclical behaviour of stock prices follows a consistent pattern; they are smoother in the pre-1985 period compared with the post-1985 period. Second, there are longer spikes in the cyclical components of the Japanese stock price index.

INSERT FIGURE 2

We attempt to gain further insights on the behaviour of stock prices through examining the stock market returns for each of three countries. We plot stock returns of each of the countries in Figures 3-5. A visual inspection suggests that the stock returns have mostly fluctuated within the ± 10 per cent band for the UK and within the ± 5 per cent band for the USA and Japan. Moreover, there seems to be more variability in stock returns for the USA and Japan. We explore the issue of volatility next through examining the conditional standard deviations for the UK, the USA, and Japan (see Figures 6-8). We make two observations here. First, there seems to be several spikes in volatility for all the three countries, and the spikes seem to be larger for the USA, particularly in the post-2000 period. Second, volatility clustering is evident for all the three stock price indices.

INSERT FIGURES 3-8

Some summary statistics for stock market returns of the USA, the UK and Japan are presented in Table 1. We notice that the mean growth rate is highest for the UK, followed by the USA and Japan, while volatility (as measured by the standard deviation) is the highest for the UK and lowest for Japan. The UK, which had the highest average return, also had the highest volatility. The kurtosis statistic is greater than 3 for all the three

10

countries stock returns, implying that the distribution is peaked (leptokurtic) relative to the normal. Because of excess kurtosis, it is not surprising that there is strong evidence of non-normality, as indicated by the Jargue-Bera test (see last row of Table 1).

INSERT TABLE 1

Skewness is a measure of asymmetry of the distribution of the series around its mean. The skewness of an asymmetric distribution, such as a normal distribution, is zero. The skweness is positive (has a right tail) for the UK and negative (has a left tail) for the USA and Japan.

We explore the dynamics of volatility for the three countries through estimating an exponential generalized autoregressive heteroskedasticity (EGARCH) mode1 proposed by Nelson (1991). The aim of this exercise is twofold: to examine volatility persistence and investigate whether shocks to volatility have asymmetric or symmetric effects. The variance model under the EGARCH framework is as follows:

$$\log(\sigma_t^2) = \omega + \alpha \left(\left| \frac{\varepsilon_{t-1}}{\sigma_{t-1}} \right| - \sqrt{\frac{2}{\delta}} \right) + \gamma \frac{\varepsilon_{t-1}}{\sigma_{t-1}} + \beta \log(\sigma_{t-1}^2) \tag{5}$$

The estimate of β allows one to evaluate whether shocks to the variance are persistent or not, while the parameter γ allows one to judge asymmetric volatility. If $\gamma > 0$, the implication is that positive shocks give rise to higher volatility than negative shocks, and vice versa.

The results are reported in Table 2. We find that the coefficient on γ is negative and statistically significant at the 5 per cent level or better for all the three countries, implying that negative shocks give rise to higher volatility than positive shocks. The coefficient on β, while statistically insignificant for the USA and the UK, is significant in the case of Japan but with a very small (0.4) magnitude. This implies that shocks to volatility are not persistent.

INSERT TABLE 2

In sum, our preliminary analysis suggests that the stock prices of the USA, the UK, and Japan share some common characteristics. It follows that modeling them together in an econometrics sense, as we do in the next section, is meaningful.

4.2.   Unit root tests

The starting point for our empirical analysis is an investigation of the integration properties of the data series. To achieve this, we apply the conventional Dickey and Fuller (1979) test, that is widely used hence we refrain from discussing the methodology here, which examines the null hypothesis of nonstationarity with the t-statistic.

Our finding from the ADF test is that the null hypothesis cannot be rejected given that the calculated t test statistics for the levels of all the three countries stock price series' are greater than the critical value at the 5 per cent level. When we take the first difference of the stock price series, we are able to reject the null hypothesis implying that stock price indices are integrated of order one.

12

However, following the work of Perron (1989) there is a caveat on the results obtained from the ADF test, for the failure to reject the unit root null hypothesis maybe due to the fact that the ADF test does not incorporate structural breaks in the data series. To circumvent this distortion, we apply the Lumsdaine and Papell (1997) model[1], which allows for two structural breaks in the intercept and two structural breaks in the slope, and the breaks are selected endogenously. The model takes the following form:

$$\Delta y_t = \kappa + \alpha y_{t-1} + \beta t + \theta DU1_t + \gamma DT1_t + \psi DU2_t + \omega DT2_t + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \qquad (6)$$

The null hypothesis is that $\alpha = 0$, which implies there is a unit root in y_t. The alternative hypothesis is that $\alpha < 0$, which implies that y_t is breakpoint stationary. $DU1_t$ and $DU2_t$ are indicator dummy variables for a mean shift occurring at $TB1$ and $TB2$ respectively, where $TB2 > TB1+2$ and $DT1_t$ and $DT2_t$ are the corresponding trend shift variables. $DU1_t = 1$ if $t > TB1$ and zero otherwise, $DU2_t = 1$ if $t > TB2$ and zero otherwise, and $DT1_t = t - TB1$ if $t > TB1$ and zero otherwise, and $DT2_t = t - TB2$ if $t > TB2$.

The lag length is selected using the Hall (1994) 't-sig' method, in that we begin with a maximum of 8 lags and use the 10 per cent significance level to select the optimal lag length. The critical values are calculated using the approach outlined in the Zivot and Andrews (1992) test.

The calculated t-statistic for the US stock price series turns out to be -2.07 while the critical value at the 5 per cent level of significance is -6.56, implying that the null hypothesis of a unit root cannot be rejected. The breaks suggested by the LP test are

13

1971:10 and 1972:12; however, both slope breaks are statistically insignificant at the 10 per cent level. For the UK stock price series, the calculated t-statistic turns out to be -2.82 and given the 5 per cent level critical value of -6.50, we are unable to reject the null hypothesis of a unit root.

The break dates (1979:03 and 1996:10) are statistically significant at the 10 per cent level for breaks in the slope. Meanwhile, for Japan's stock price series the t-statistic turns out to be -6.81; given the 5 per cent critical value of -6.48 and the 1 per cent critical value at the 1 per cent level of -6.92 we are able to reject the unit root null hypothesis at the 5 per cent but not at the 1 per cent level. The break dates are 1983:08 and 1989:11: the first break in the intercept and both breaks in the slope are statistically significant at the 1 per cent level.

4.3. *Cointegration*

To test for a cointegration relationship amongst the stock price indices, we use the bounds testing approach to cointegration, which is applicable irrespective of whether or not the variables are integrated of order zero or one. The model is based on the following unrestricted error correction model:

$$\Delta \ln SPJ_t = a_{0SPJ} + \sum_{i=1}^{n} b_{iSPJ} \Delta \ln SPJ_{t-i} + \sum_{i=0}^{n} c_{iSPJ} \Delta \ln SPUS_{t-i} + \sum_{i=0}^{n} c_{iSPJ} \Delta \ln SPUK_{t-i}$$
$$+ \lambda_{1SPJ} \ln SPJ_{t-1} + \lambda_{2SPJ} \ln SPUS_{t-1} + \lambda_{3SPJ} \ln SPUK_{t-1} + \varepsilon_{1t} \tag{7}$$

Here SPJ, $SPUS$ and $SPUK$ are the stock prices indexes for Japan, USA and the UK respectively. The lag length, *n*, is selected using the Schwarz Bayesian Criterion.

Equation (6) is estimated by taking each of the countries stock price series as the dependent variable. Hence, when a long-run relationship exists, the *F* test on the joint significance of the one period lagged level variables indicates which variable should be normalised.

The calculated *F* test statistic, which has a non-standard distribution, depends upon: (a) whether variables included in the model are $I(0)$ or $I(1)$, (b) the number of regressors, and (c) whether the model contains an intercept and/or a trend. The critical values are reported in Pesaran *et al*. (2001).

The calculated F-statistic when Japan's stock price series is the dependent variable is 5.06, when the US stock price series is the dependent variable the F-statistic is 2.32 and when the UK stock price is the dependent variable the F-statistic is 1.28. Given the 5 per cent level critical value of 4.05 the null hypothesis of no cointegration is rejected only when Japan's stock price series is treated as the dependent variable. In other words, there is a cointegration relationship between the three countries stock prices only when Japan's stock price series is the dependent variable. Bivariate test for cointegration was also undertaken and no evidence of a cointegration was found between any two markets. The detailed results are available from the author upon request.

4.4.   Common break test

The results on the break test statistic are reported in Table 3. Three set of results are estimated and presented: (1) Panel 1 consists of the univariate test statistics, (2) panel 2

reports the results for the bivariate systems, and (3) panel 3 reports the results for the multivariate cointegrated systems with cointegrating coefficients estimated using the autoregressive distributed lag estimator. Following Bai *et al*. (1998), we select the lag lengths using the Schwarz Bayesian Criterion: for the univariate model we adopt a maximum of 6 lags; for the bivariate model we adopt a maximum of 4 lags; and for the trivariate model we adopt a maximum of 3 lags.

The null hypothesis for a constant stock price series cannot be rejected for the USA and the UK at the 10 per cent level of significance. While the break dates are 82:10 and 97:7 for the USA and the UK respectively, the confidence interval is very imprecise – it is so wide that it includes the entire sample in the case of the USA and almost half the sample in the case of the UK. Meanwhile in the case of Japan's stock price, the null hypothesis is rejected at the 1 per cent level of significance; the break date is estimated to be 1990:02; and the 90 per cent confidence interval spans almost 13 years.

The results on the bivariate models are as follows. In the case of the USA-UK stock price series', the SupW test statistic rejects the null hypothesis of no break at the 1 per cent level of significance with the point estimate of the break date been 1994:04, which, however, has an imprecise 90 per cent confidence interval spanning 14 years. On the other hand, in the case of Japan-USA and the UK-Japan stock price series', the null hypothesis is rejected at the 1 per cent level of significance and the break date for both set of markets is 90:02. However, the 90 per cent confidence interval is wide, spanning over 5 years in the case of Japan-USA and over 9 years in the case of the UK-Japan stock

price series'. The break date confidence intervals for the case of the USUK-VAR (87:05, 01:03), USJP-VAR (87:01, 93:03) begin with the 1987 US stock market crash.

INSERT TABLE 3

Next we investigate the null hypothesis by treating all the three countries stock price series, earlier found to be cointegrated, as a VECM. The null hypothesis is rejected at the 1 per cent level implying that the countries share a common break date, 90:02, which has a 90 per cent confidence interval of 88:09-91:07, spanning less than 3 years. On the latter finding, notice that the confidence interval is much tighter compared to our earlier findings. A much tighter confidence interval is obtained when we consider the null hypothesis taking the stock price series as a VAR. In this case, the 90 per cent confidence interval declined from 3 years to around 1 year.

In terms of the location of the common break date (90:02) and the confidence interval, it can be associated with several events such as the asset price bubble when housing prices and stock prices in Japan reached a peak in 1988/1989, the early 1990s recession in the UK, the business cycle peak of July 1990, the August 1990 Iraqi invasion of Kuwait and the March 1991 business cycle trough.

4.5.   Growth rates in stock markets

In this section, we calculate the annual average growth rate of stock prices for each of the three markets on the basis of the common structural break of February 1990. Given this

common break, for each of the stock markets we divide the sample period into two: one period includes the structural break, while the other period excludes the structural break, allowing us to gauge whether or not the structural break slowed down stock market growth rate.

Our findings suggest that in the case of the UK's and Japan's, the structural break slowed down the growth rate of stock markets. For instance, in the case of the UK, for the period 1964:06 to 1990:02, the annual average growth rate in FT-SE-A Non-Financials was 0.9 per cent, while the corresponding growth rate in the period excluding the structural break was approximately 0.3 per cent. The slow down in Japan's stock market was more pronounced: while the annual average growth rate in TSE TOPIX in the period including the structural break was around 0.9 per cent, it had plummeted to an annual average rate of -0.3 per cent in the period excluding the structural break.

Meanwhile, the stock market development for the US was positive: the structural break actually boosted stock prices. The annual average growth rate of the NYSE Common Stocks increased from around 0.6 per cent over the 1964:06 to 1990:02 period to 0.9 per cent over the 1990:03 to 2003:04 period.

5. Conclusions

The literature on estimating structural breaks has attracted immense interest over the last couple of decades. In this regard, an important innovation has been Bai *et al*. (1998), who have developed a test that allows one to deduce evidence for a common structural break

in cointegrated series. In this paper, our goal was to revisit three of the largest global stock markets, namely the USA, the UK and Japan, and investigate whether these markets share a common break. A related contribution of Bai *et al*. (1998) is that their methodology allows one to derive a confidence interval for a given structural break.

Our results suggest that when the USA market is modeled with the UK and Japanese stock markets in a multivariate sense, the break date falls in the interval encompassing the US stock market crash of 1987. However, when the stock markets are modeled in a trivariate sense the confidence interval for a common break includes several episodes such as the asset price bubble when housing prices and stock prices in Japan reached a peak in 1988/1989, the early 1990s recession in the UK, the business cycle peak of July 1990, the August 1990 Iraqi invasion of Kuwait and the March 1991 business cycle trough.

Analyses of the common structural break matters because it allows one to gain more insights on the behaviour of stock prices. Our approach, on the basis of the obtained structural break, was to divide the sample into sub-samples, culminating into two periods. Period 1 included a sample including the structural break while period 2 was one excluding the structural break. We then calculate annual average growth rates in each countries stock price over the two sample periods. The division of sample in this way and the calculation of annual average growth rates allow one to gauge whether or not the structural break slowed down growth of stock markets. Our findings are interesting. We

find that while the structural break has slowed down the growth the UK and Japanese stock markets, it has boosted the growth of the USA stock market.

It closing, it is worth noting that the stock market analysis conducted in this paper is innovative and novel but not inclusive, meaning that there remains avenues for further research on stock market development. Future studies, for instance, might examine the presence of a common structural break in stock prices and gross domestic product (GDP). Using the structural break, the sample period can be divided into two for the two series, allowing one to deduce whether or not structural breaks caused a simultaneously slowdown in stock prices and GDP.

On the methodological front, work can be conducted to allow for two common breaks in bivariate and trivariate cases. Such a methodological innovation will allow one to draw further insights on the importance of structural changes on stock prices. For instance, with two common breaks, one will be able to obtain three sub-samples. An interesting question, among others, in this regard will be whether or not both breaks slowed down growth rates.

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Table 1: Some descriptive statistics of stock returns

	USA	UK	Japan
Mean	0.5146	0.5950	0.4359
Standard deviation	3.5840	4.7499	4.1609
Skewness	-0.6592	0.3209	-0.2394
Kurtosis	4.7379	11.4983	3.8011
Jarque-Bera (probability)	92.3999 (0.0000)	1410.29 (0.0000)	16.9105 (0.0000)

Table 2: Results from the variance equation of the EGARCH model

	USA	UK	Japan
ω	2.2546***	2.9403***	1.5576***
	(5.7003)	(8.8433)	(5.6224)
α	-0.0773	0.0433	0.0113
	(-1.1650)	(1.5542)	(0.4693)
γ	-0.2577***	-0.1679***	-0.1525**
	(-3.2329)	(-2.8974)	(-2.4475)
β	0.1163	-0.0175	0.4290***
	(0.8054)	(-0.1660)	(4.2432)

Note: ** (***) denote statistical significance at the 5 per cent and 1 per cent levels, respectively.

Table 3: US, UK and Japan stock market analysis of a common break

	p	$Sup-W$	$Exp-W$	$\hat{k}$	90% confidence interval
A. Univariate					
US	1	2.35**	0.73	82:10	(<70:04, >97:01)
		(0.05)	(0.45)		
UK	1	2.08***	0.20	97:07	(79:05, >96:01)
		(0.09)	(0.70)		
JP	1	9.65*	1.36	90:02	(83:06, 96:10)
		(0.00)	(0.85)		
B. Bivariate					
USUK-VAR	1	6.89*	1.40	94.04	(87:05,01:03)
		(0.00)	(0.11)		
USJP-VAR	1	17.79*	5.00*	90.02	(87:01-93:03)
		(0.00)	(0.00)		
UKJP-VAR	2	13.81*	2.74**	90:02	(85:08,94:08)
		(0.00)	(0.03)		
C. Multivariate, with estimated cointegrating coefficients					
USUKJP-VECM	1	33.76*	11.32*	90:02	(89:07,90:09)
		(0.00)	(0.01)		
USUKJP, Triangular form	1	10.80*	1.78	90:02	(88:09,91:07)
		(0.01)	(0.11)		

Notes: * (**) *** denote statistical significance at the 1 per cent, 5 per cent and 10 per cent levels respectively, with p-values reported in parentheses. The lag length p was selected using the Bayesian Information Criterion.

Figure 1: Stock price index
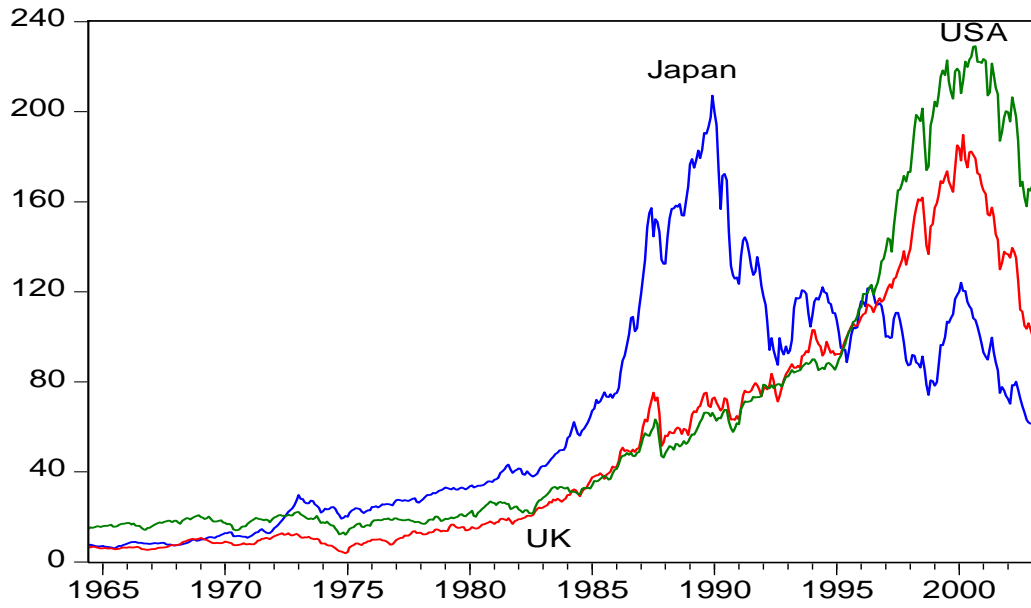
Figure 2: Cyclical components of stock price series for the USA, the UK and Japan
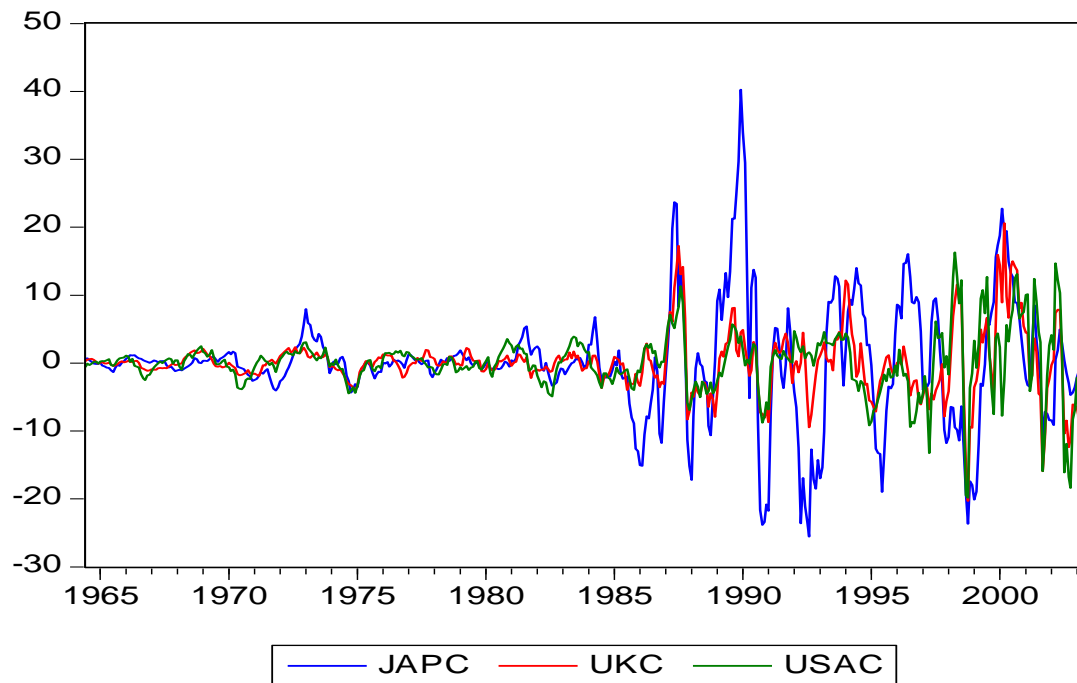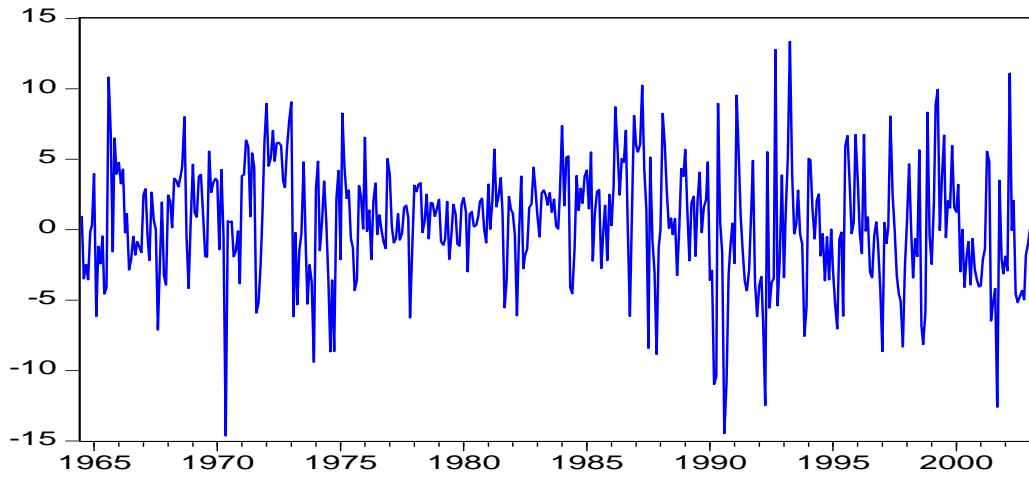
Figure 3: Growth rate in Japan's stock price

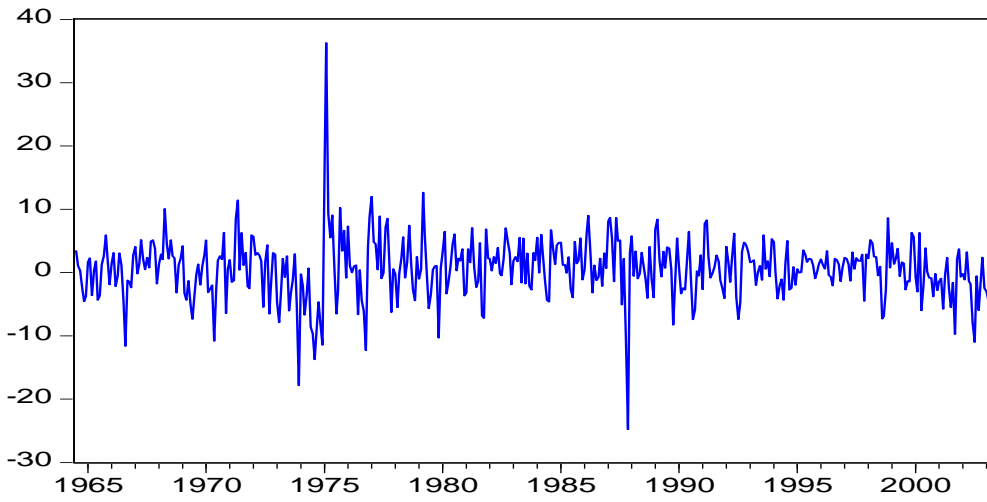Figure 4: Growth rate in UK's stock price

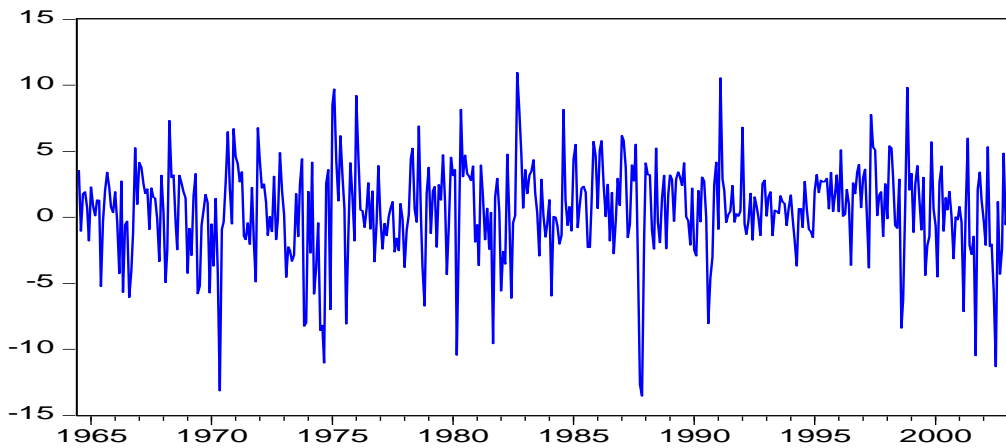Figure 5: Growth rate in USA's stock price

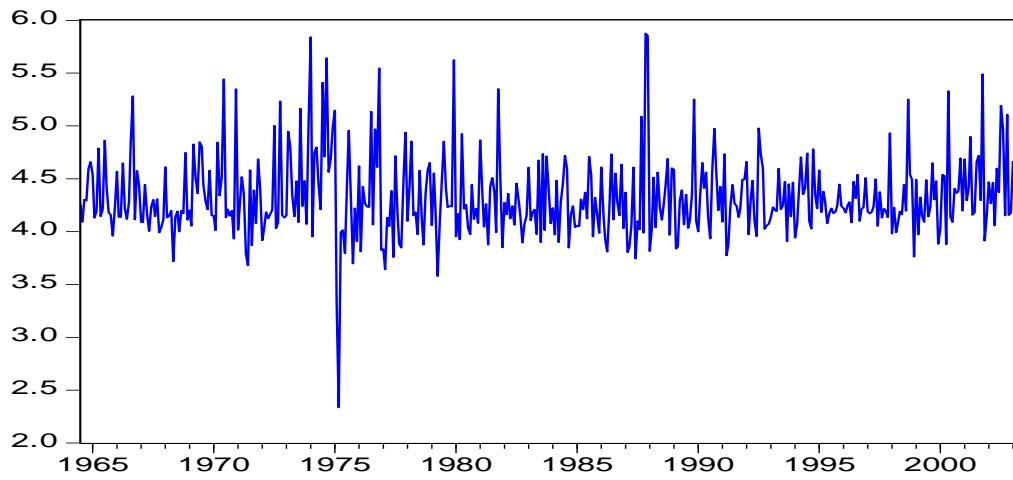

29

Figure 6: Conditional standard deviation, UK

Figure 7: Conditional standard deviation, USA
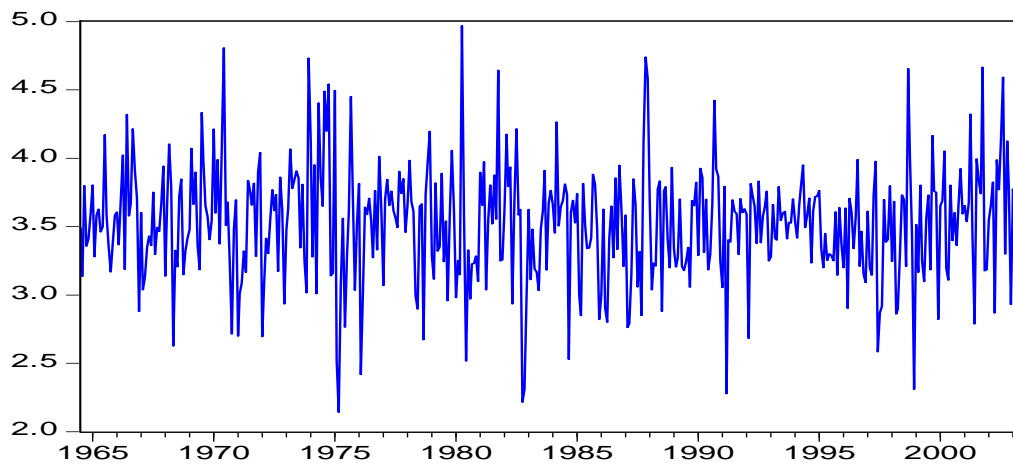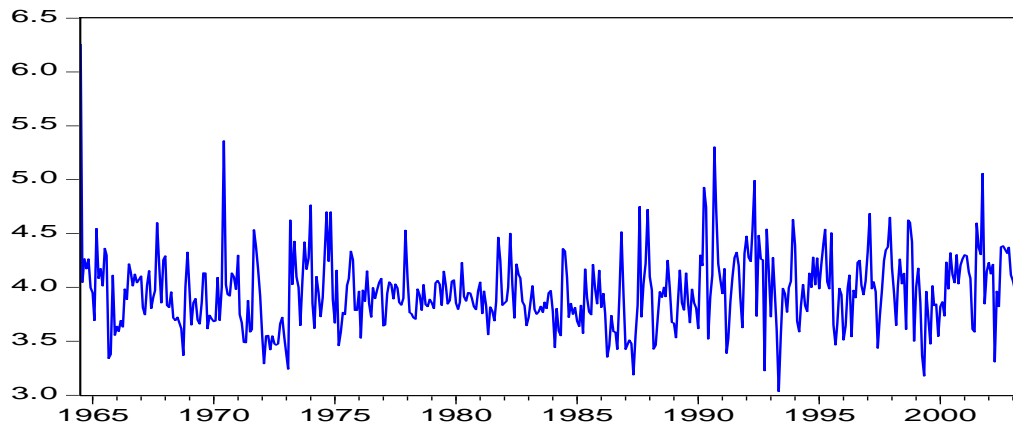

Figure 8: Conditional standard deviation, Japan

ENDNOTES

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[1] We also conducted the Lee and Strazicich (2003) test for unit root which allows one to endogenously search and account for two structural breaks. We find similar results; thus, we do not report the results here. However, the results are available from the author upon request.