

Common Stochastic Trends among the Cyprus Stock Exchange and the ASE, LSE and NYSE

by

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Abstract

Common stochastic trends among major international stock price indices has been a very intensively analyzed issue mainly as a result of the 1987 stock market crash and the need for policy coordination in financial markets. This paper investigates the existence of common stochastic trends among an emerging equity market, the Cyprus Stock Exchange and three mature equity markets namely ASE, LSE and NYSE. The main finding of our analysis is that there is evidence of one long-run relationship among the four equity markets and therefore three common stochastic trends. We use the Gonzalo and Granger (1995) methodology to identify, estimate and test for the number of common trends that leads to permanent changes among the four stock markets. Furthermore, we identify as driving forces of the system the ASE, LSE and NYSE equity markets while the emerging stock market of Cyprus does not enter significantly the common trends. Finally, we show that although cointegration exists there are small long-run benefits from international portfolio diversification since the stock prices adjust very slowly to these common trends.

Keywords: cointegration, common trends, identification, international stock markets, international portfolio diversification.

JEL Classification: G15, G21, G32, G34

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1. Introduction

International portfolio theory argues that investors would benefit through international diversification when there is low degree of comovements among the national stock markets. Thus, during the last two decades an important and challenging issue in empirical finance is whether national stock markets show similar price patterns and are converging over time or capital market integration is a rather elusive concept. The vast literature on the many aspects of stock market comovements gives an indication of this interest among researchers. The stock market crash of 1987 has resulted to an increase in the interest of researchers on the matter and as a result a great number of papers have attempted to document the existence of linkages between returns of stock indices of the major equities markets and more recently of several emerging capital markets.

The issues concerning capital markets linkages are significant for a number of reasons. First, from a long-run perspective if there exist a comovement among the prices of national stock markets and therefore they share a single common trend with only transitory deviations from this trend, then there will be limited benefits from long-run diversification. From a European point of view this result would imply the need for an increase financial markets integration, which is considered by some economists as an important step for the increasing international role of Euro and the successful implementation of EMU. Second, evidence of strong linkages among world capital markets may lead to the rejection of the efficient markets hypothesis in the framework defined by Granger (1986) who considers the incompatibility between cointegration of two or more asset prices and the fact that these prices are derived

from efficient markets.¹ However, today it is argued that evidence of cointegration among stock prices may not necessarily imply violation of market efficiency since if fundamentals are cointegrated, this will also lead to cointegration among stock prices. Finally, it is important to distinguish between the issue of long-run comovements among stock indices and capital market integration. The latter describes a state where increased interdependence and policy coordination among countries exist. As Taylor and Tonks (1989) underline that an implication of capital market integration is that a specific model of required returns is employed to all assets irrespectively to the market they are traded.²

Kasa (1992) provided the initial evidence on international stock price linkages and comovements of fundamentals over long period in a multivariate cointegration framework. His main finding was that there is strong evidence in favour of a single common stochastic trend that drives the system of the stock indices for the U.S.A. Canada, Germany, Japan and the United Kingdom. Following Kasa's work a substantial number of papers have further analyzed the existence of long-run linkages between stock markets and by now it has been well documented that stock markets move together (Cohray *et al.*; 1993, Kasa, 1995; Richards; 1995, Serletis and King, 1997, Francis and Leachman, 1998; Bracker *et al.*; 1999, Georgoutsos and Kouretas, 2003; Engsted and Tangaard, 2004 and Fraser and Oyefeso, 2005 are some of these studies).

¹ Granger's (1986, p.213) argument relies on the error correction representation theorem and the one-to-one correspondence between cointegration and the error correction. The existence of an error correction model implies predictability of one asset price from the other. However, Richards (1995) argues that these results apply strictly on the total returns (i.e. cum dividends) or otherwise or non-interest / dividend paying assets. Therefore, the efficient markets hypothesis may still be compatible with the existence of cointegration if we consider the error correction as a proxy for the risk premium.

² Lucas (1982) and Kasa (1995) argue that capital market integration implies the equalization among countries of marginal rates of substitution in consumption both inter-temporally and across states of nature.

In this paper we examine the issue of finding common trends between an emerging capital, the Cyprus Stock Exchange (CSE), and three mature markets that of the Athens Stock Exchange (ASE), the London Stock Exchange (LSE) and the New York Stock Exchange (NYSE). The reason for paying attention on an emerging market is due to the fact that during the 1990s there has been an increasing flow of funds from developing countries towards these markets (Hawanini, 1994). Indeed, during the 1990s the opening of the economies of several developing countries and in particular those of Central and Eastern European countries, Latin America and South East Asia along with their consequent economic growth has attracted the interest of international portfolio managers of American and European funds. Thus, while the net assets invested (in US dollars) in the Southeast Asia region was 4000 in 1990 they have grown to 65,300 in 2001.³ Moreover, the total capitalization of these markets during this period has increased enormously ranging from 362 per cent in Singapore to 5105 per cent in Thailand, which has attracted a large number of assets and funds.

The starting point in our analysis is the argument set forth by Kasa (1992) that it is not only significant to provide evidence for the existence of cointegration among the four stock indices under question but also the degree to which these stock prices respond to the international common factor(s). Kasa (1992, p.97) notes

“Of course, the economic relevance of any long-term comovement hinges on the speed of adjustment towards the common trend. If transitory deviations from the common trend have a persistence measured in decades, then to an investor with a finite horizon the existence of a common trend is of little significance.”

Our testing approach follows four steps. First we use some recently developed unit root tests due to Elliot *et al.* (1996), Elliott (1999) and Ng and Perron (2001) which correct some deficiencies of the standard ADF and Phillips and Perron tests.

³ See the Micropal Directory of Emerging Markets.

Second, we test for cointegration with the use of the single equation Philips and Hansen (1990) fully modified OLS methodology and Johansen (1988, 1991) and Johansen and Juselius (1990, 1992) multivariate cointegration methodology. Third, by considering the Moving Average representation of the vector error correction model we look into the issue of decomposing the stock market price series into their permanent and temporary components and we then analyze how long does it take stock prices to return to the common trend or permanent component. Finally, given Kasa's (1992) aforementioned argument this is a crucial step since even in the case that stock markets exhibit common stochastic trend we may still face the situation that the returns do not react to the trend. From the point of view of portfolio diversification this implies that in the case that the returns react very slowly (or not at all) to the trend then the benefits from international diversification still exist even in the presence of cointegration (see also Garrett and Spyrou, 1999).⁴

The rest of the paper is organized as follows. In section 2 we present a brief discussion of the literature on common stochastic trend in international stock markets. Section 3 presents the econometric methodology. Section 4 discusses the data and presents the empirical results with our conclusions given in section 5.

2. Common stochastic trends in international stock markets

An important issue of portfolio theory deals with the advantages derived from international diversification. The rapid technological innovations in the last two decades have led to the elimination of the barriers in trading in global financial markets and that resulted to a dramatic increase in capital flows from one country to the other. Trading now takes place on a 24 hour basis whereas in most developed

⁴ Crowder and Wohar (1998), Rangvid (2001), Georgoutsos and Kouretas (2003) and Pascual (2003) also consider the stability issue when analyzing stock markets comovements.

countries foreign exchange controls have been abolished. Therefore, these developments have led to an increased degree of capital market integration and research has focused on the issue that stock markets drift upwards over time and whether there is a common trend among national stock markets.

Over the years the econometric techniques employed range from simple regression, to autocorrelation analysis, to spectral analysis, to vector autoregressions, to variance decomposition, to cointegration methods. Cointegration theory provides the analytical framework within which the main strand of econometric analysis is conducted recently. The initial works on the issue Taylor and Tonks (1989), Chan *et al.* (1992) and Arshanapalli and Doukas (1993) employed the two-step Engle-Granger (1987) cointegration technique provide conflicting evidence in favour of cointegration between stock prices of major equity markets in a bivariate context.⁵ As we have already discussed Kasa (1992) identifies a single common factor for the case of five major equity markets. However, Richards (1995) has criticized these findings on the grounds that they have been obtained with an arbitrary choice of the lag length of the estimated VAR and no adjustment of the critical values of the Johansen's tests. The evidence from subsequent works is not clear-cut. Thus, Francis and Leachman (1988) found one cointegration vector among the stock indices of U.S.A., U.K., Germany and Japan but this implies that there are three stochastic trends that can not be identified. Serletis and King (1997) examined the issue within the European Union context and they failed to find on common stochastic trend. They explained this evidence on the low integration of the Athens Stock Exchange with the other European capital markets. Fraser and Oyefeso (2005) also examine the long-run

⁵ Other works have studied the issue by focusing on the short-run interdependencies of price levels and/or price volatility across national equity markets (Eun and Shin, 1989; King and Wadhwani; 1990, Longhin and Solnik, 1995, Masih and Masih; 2001 and Ratanapakorn and Sharma; 2002). Moreover, Engsted and Tanggaard (2004) provide evidence in favour of comovements between the US and UK stock markets within a simple present value model.

interrelationships of the European capital markets and they conclude that although cointegration exists the gains from diversification are short-lived since the adjustment to the common trend is very slow. Crowder and Wohar (1998) employed the same dataset as Kasa (1992) and they found that the system is better described by four common trends a results which also consistent with the recursive stability tests they employ. Recently Georgoutsos and Kouretas (2003) have shown that the failure to identify one common trend may be explained by the validity of PPP. Thus, if PPP holds then the use of stock indices denominated in local currency will be equivalent to the use of a common currency in real terms. However, if PPP does not hold then the conversion of local prices to a real common currency may affect our search for a common stochastic trend. Finally, with respect to emerging markets Garrett and Spyrou (1999) find evidence of three common stochastic trends for the stock markets of Latin America and four common trends for South East Asia with very slow adjustment of returns to these trends while Scheicher (2001) provide evidence of cointegration for the stock markets of Poland, Hungary and the Czech Republic.

3. Econometric methodology

In order to evaluate whether the potential benefits from international portfolio diversification that are accumulated in the short run carry over to the long run, we are required to analyze whether a long run relationship exists among the four equities markets. We conduct the cointegration analysis first in a bivariate using the single equation Philips and Hansen (1990) fully modified OLS.

Consider the following pairwise cointegrating regression:

$$y_t = \alpha + bx_t + z_t \quad (1)$$

where z_t is the residual series of the cointegrating regression, and y_t and x_t are the two variables to be tested for cointegration. The \hat{P}_z statistic tests the null hypothesis of no cointegration, and is calculated as

$$\hat{P}_z = T \text{ trace}[\Omega_p T^{-1} \Sigma_1^T X_t X_t'] \quad (2)$$

where $\Omega_p = T^{-1} \Sigma_1^T z_t z_t' + T^{-1} \Sigma_1^k w_{sk} \Sigma_{s+1}^T (z_t z_{t-s}' + z_{t-s} z_t')$ for some choice of lag window such as $w_{sk} = (1 - s/(k+1))$, T is the sample size, $X_t' = (y_t, x_t)$, and z_t are the residuals from estimating the above bivariate model with orthogonal regression. Using orthogonal regression, which is invariant to the formulation of the above regression, renders the \hat{P}_z statistic invariant to the normalisation of the cointegrating regression (Phillips and Ouliaris, 1990, p. 172). Moreover, the power of the \hat{P}_z test has been found to be relatively high. Critical values for \hat{P}_z test are reported in Phillips and Ouliaris (1990, Table IV). Its asymptotic distribution is free of nuisance parameters, and is dependent only on the number of explanatory variables in the cointegrating regression.

We next turn to the multivariate cointegration analysis. We apply the Johansen (1988, 1991) and Johansen and Juselius (1990, 1992), multivariate cointegration methodology with well behaved (Gaussian) errors need to be estimated in the form of unrestricted VARs.

The Johansen (1988, 1991) framework involves estimating the following vector-error correction model (ECM):

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \dots + \Gamma_{k-1} z_{t-k+1} + \Pi z_{t-k} + \gamma D_t + \mu + \varepsilon_t \quad (3)$$

where z_t is a column vector of stochastic variables, $\varepsilon_t \sim N(0, \Sigma)$. The parameters $(\Gamma_1, \dots, \Gamma_{k-1}, \gamma)$ define the short-run adjustment to the changes of the process, whereas $\Pi = \alpha\beta'$ defines the short-run adjustment, α , to the cointegrating relationships, β . If the short-run effects are basically different from the long-run effects, due for instance, to costly arbitrage and/or imperfect information, the explicit specification of the short-run effects is probably crucial for a successful estimation of the steady-state relations of interest. D_t is a vector of nonstochastic variables, such as centered seasonal dummies which sum to zero over a full year by construction and are necessary to account for short-run effects which could otherwise violate the Gaussian assumption, and/or intervention dummies; μ is a drift and T is the sample size.

Johansen (1991) shows that if $Z_t \sim I(1)$, the following restrictions on model (3) have to be satisfied:

$$\Pi = \alpha\beta' \quad (4)$$

where Π has reduced rank, r , α and β are $(p \times r)$ matrices, and

$$\Psi = \alpha_{\perp}(-I + \Gamma_1)\beta_{\perp} = \varphi\eta' \quad (5)$$

where Ψ is a $(p-r) \times (p-r)$ matrix of full rank, φ and η are $(p-r) \times (p-r)$ matrices, and α_{\perp} and β_{\perp} are $p \times (p-r)$ matrices orthogonal to α and β , respectively. The parameterization in (3) and (4) facilitates the investigation of, on the one hand, the r linearly-independent stationary relations between the levels of the variables and, on the other hand, the $p-r$ linearly-independent non-stationary relations. This duality between the stationary relations and the non-stationary common trends is very useful for a full understanding of the generating mechanisms behind the chosen data. While the AR representation of the model is useful for the analysis of the long-run relations in the data, the MA representation is useful for the

analysis of the common stochastic and deterministic trends that have generated the data.

The interesting problem to solve is to determine the restrictions that model (3) should satisfy in order to derive, if possible, this long-run specification that is common in the literature. Furthermore, in anticipation of the empirical results, those restrictions should comply with interesting economic scenarios that are better understood through the moving average representation of z_t . This representation of model (3) for the case of $I(1)$ variables is

$$z_t = C \sum_{i=1}^t \varepsilon_i + C \mu t + C \Phi \sum_{i=1}^t D_i + C^*(L)(\varepsilon_t + \mu + \Phi D_t) + Z_0, \quad (6)$$

where $C = \beta_{\perp}(\alpha'_{\perp} \Gamma \beta_{\perp})^{-1} \alpha'_{\perp}$, $C^*(L)$ is a polynomial in the lag operator L , Z_0 is a function of the initial values and $\alpha_{\perp}, \beta_{\perp}$ are both $(4 \times (4-r))$ matrices orthogonal to α and β respectively. In the MA representation $(\alpha'_{\perp} \sum \varepsilon)$ determine the $(4-r)$ stochastic trends while β_{\perp} the variables that are being influenced by them. The realization at time t of the variables in z_t are determined by a stochastic trend component, described by the first term of eq. (6), a deterministic trend component, the cumulated value of the non-stationary variables D_t , a stationary component and the initial values.

It is obvious from equation (6) that the common trends in z_t are contained in the first term of that expression. Johansen (1995, p.41), given the definition of C , defines the common trends by the cumulated disturbances $\alpha'_{\perp} \sum_{i=1}^t \varepsilon_i$. Based on the assumption that the common trends are considered as a linear transformation of z_t , in

the form $f_t = \alpha'_\perp z_t$, Gonzalo and Granger (1995) proposed the following decomposition of any cointegrating system into its permanent and transitory components

$$z_t = A_1 f_t + A_2 y_t \quad (7)$$

where, $y_t = \beta' z_t$, $A_1 = \beta'_\perp (\alpha'_\perp \beta_\perp)^{-1}$ and $A_2 = \alpha (\beta' \alpha)^{-1}$. In addition Granger and Gonzalo (1995) provide the maximum likelihood estimation of α_\perp as well as a likelihood ratio test in order to test whether or not certain linear combinations of z_t can be considered as common trends. The null hypothesis on α_\perp takes the following form:

$$H_0 : \alpha_\perp = K\theta \quad (8)$$

where K is a $p \times m$ known matrix of constants, θ is an $m \times (p-r)$ matrix of unknown coefficients and $p-r \leq m \leq p$. Solving the corresponding restricted eigenvalue problem we construct the following likelihood ratio test statistic for testing the null hypothesis given in (8). The LR test statistic is given by:

$$L = -T \sum_{i=r+1}^p \ln \left[(1 - \hat{\lambda}_{i+(m-p)}^*) / (1 - \hat{\lambda}_i) \right] \quad (9)$$

Under the null hypothesis $H_0 : \alpha_\perp = K\theta$, the L -statistic in equation (9) is distributed asymptotically as $\chi^2_{(p-r) \times (p-m)}$.

4. Empirical results

4.1. Data and preliminary results

In this paper we study four equity markets, those of Cyprus, Athens, London and New York. We use daily data for the period March 29, 1996 (the date of the official opening of the Cyprus capital market) to April 19, 2002. The price data for the Cyprus Stock Exchange (CSE) is the general index, for the Athens Stock Exchange (ASE) we also use the general index while for the London Stock Exchange (LSE) and New York Stock Exchange (NYSE) we use the FTSE100 and the Dow Jones Industrial Average (DJIA), respectively and they have been drawn from DATASTREAM. All indices are expressed in terms of US dollars.⁶

Table 1 reports contemporaneous correlation coefficients of the stock market daily returns of the CSE, ASE, LSE and NYSE. It is shown that there is a very low correlation between the returns on equity of CSE with those of ASE, LSE and NYSE, while in the case of ASE and NYSE we observe even negative correlation and finally a moderate correlation exists between the two major markets of the US and the UK. This evidence may suggest that at least in the short run there exist benefits from portfolio diversification. Furthermore, Figures 1-4 show the evolution of stock general price index along with that of the returns series for each equities market. A visual examination of Figures 1(a) and 2(a) reveal the presence of a rational bubble that it was first experience in the ASE (1997-1999) transmitted to a great extend to the CSE (1999-2000).

To examine, whether the series under consideration are stationary, we apply the Elliot *et al.* (1996) GLS augmented Dickey-Fuller test (DF-GLS_u) and Ng and Perron (2001) GLS versions of the modified Phillips-Perron (1988) tests

⁶ Chrisostomidou et al. (2006) provide a thorough analysis of the institutional and performance of the Cyprus Stock Exchange.

(MZ_a^{GLS} and MZ_t^{GLS}). The null hypothesis is that of a unit root against the alternative that the initial observation is drawn from its unconditional distribution and uses GLS-detrending as proposed by Elliott *et al.* (1996) and extended by Elliott (1999), to maximize power, and a modified selection criterion to select the lag truncation parameter in order to minimize size distortion. In the GLS procedure of Elliot *et al.* (1996), the standard unit root tests (without trend) are applied after the series are first detrended under the local alternative $\rho = 1 + \alpha/T$. This was found to provide substantial power gains for the DF-GLS_u test resulting to power functions that lie just under the asymptotic power envelope. Ng and Perron (2001) find similar gains for the MZ_a^{GLS} and MZ_t^{GLS} tests. They also found that a modification of the AIC criterion (MIC), give rise to substantial size improvements over alternative selection rules such as BIC. For robustness, we then apply the Kwiatkowski *et al.* (1992) KPSS test for the null hypothesis of level or trend stationarity against the alternative of non-stationarity. The results of the unit root and stationarity tests are presented in Table 1. The results are reported in Table 2 and they show that we are unable to reject the null hypothesis of non-stationarity with the DF-GLS_u and MZ_a^{GLS} and MZ_t^{GLS} tests and we reject the null hypothesis of stationarity with the KPSS test for the levels of both series. The results are reversed when we take the first difference of each exchange rate series which leads us to the conclusion that all variables are realizations of $I(1)$ processes.

We also calculate several descriptive statistics for monthly percentage changes in the stock prices. These descriptive statistics are reported in Table 3. All four countries offer positive returns on equity although the potential risk (measured by the standard deviation) associated with these returns is higher in the emerging market of Cyprus and decrease as we move to more developed markets. The skewness and

kurtosis measures indicate that all series are positively skewed and highly leptokurtic relative to the normal distribution. This result is further reinforced from the Jacque-Bera statistic which implies that we reject the null hypothesis of normality. These results are in line with the well established evidence of all previous econometric studies in the literature for the stock markets (mature and emerging), i.e. that the distribution of daily stock returns is not the normal one. Rejection of normality can be partially attributed to intertemporal dependencies in the moments of the series. We also calculate the Ljung-Box (1978) portmanteau test statistics Q and Q^2 (for the squared data) to test for first- and second-moment dependencies in the distribution of the stock price changes. The Q statistic indicates that percentage monthly changes of each rates are serial correlated. This outcome can be interpreted as evidence against the market efficiency hypothesis for the CSE, which was expected given that this market is an emerging one. The Q^2 statistics for all returns series are statistically significant, providing evidence of strong second-moment dependencies (conditional heteroskedasticity) in the distribution of the stock price changes. This finding implies that there is strong evidence for the presence of non-linear dependence between the stock indices. It is also evident that the size of the statistics improves as we move from an emerging market (CSE) towards the mature markets (LSE and NYSE).

4.2. Cointegration analysis

We begin our analysis with the estimation of the cointegration equation (1) to the \hat{P}_z test. Table 4 contains results of estimating the bivariate cointegrating regression and of testing for a unit root in the cointegrating residuals. Test results are reported in Table 4, the \hat{P}_z test results reveal that the null of no cointegration is

rejected for all three pairs of indices considered. Thus, the \hat{P}_z test results lead to the conclusion that there is statistically significant evidence in favour of cointegration, i.e. there is a long-run comovement between the CSE and each of the indices of the ASE, LSE and the NYSE. From a portfolio diversification perspective, this finding is interpreted as evidence that for a Cypriot investor diversifying his/her portfolio internationally in Athens, London or New York, the long-run diversification benefits, in terms of risk reduction, may be reduced. This finding might be the consequence of the economic policy initiated in Cyprus aiming at the convergence of the Cypriot economy to the EU average, in order to facilitate the Cyprus entry to the EU.

Given this evidence in favour of bivariate cointegration we move now to the multivariate cointegration analysis in order to obtain richer insights to possible long-run relationships among the four equities markets. The first stage of our analysis is the determination of the cointegration rank index, r . As a first check for the statistical adequacy of model (1) we report some univariate misspecification tests in Table 5, in order to investigate that the estimated residuals do not deviate from being Gaussian white noise errors.

A structure of nine lags for each case was chosen based on these misspecification tests. We note that our conditional VAR model is well specified, except for the presence of non-normality. Normality can be rejected as a result of skewness (third moment) or excess kurtosis (fourth moment). Since the properties of the cointegration estimators are more sensitive to deviations from normality due to skewness than to excess kurtosis we report the univariate Jarque-Bera test statistics together with the third and fourth moment around the mean. It turns out that the rejection of normality is essentially due to excess kurtosis, and hence not so serious for the estimation results. The ARCH(9) tests for ninth order autoregressive

heteroscedasticity and is rejected for all equations. Again cointegration estimates are not very sensitive to ARCH effects.⁷ The R^2 measures the improvement in explanatory power relative to the random walk (with drift) hypothesis, i.e. $\Delta x_t = \mu + \varepsilon_t$. They show that with this information set we can explain quite a large proportion of the variation in the inflation rates, but to a much lesser extent the variation in the bilateral exchange rates and the stock price indices. Finally, based on Johansen's (1992a,b) testing methodology we choose to estimate a model with a constant restricted in the cointegrating vector.

The results of both the trace and maximum eigenvalue tests are presented in Table 6. Our results indicate that there exists one cointegration vector between the for equity markets since we reject the null hypothesis of no cointegration but we are unable to reject the hypothesis that there exist more than one long-run relationship. This finding suggests that there are three common stochastic trends driving these four stock markets. These results are further reinforced by the estimated roots of the companion matrix and the values of the corresponding eigenvalues as suggested by Juselius (1995). Thus, in Table 6 we also report the estimated eigenvalues and the estimated roots of the unrestricted and restricted to one cointegrating vector model. It is clear that the first eigenvalue is large enough while the other three approach zero, while when we restrict our model to have one cointegrating vector then we observe that the fourth root is substantially far from one.

⁷ It is well known how important is to select the correct lag specification in the VAR. Kasa (1992) has estimated VAR models with very high order lag structures in an attempt to eliminate non-normality and increase the likelihood to capture the presence of any long –run horizon mean reversion. However, Gonzalo (1994) shows (a) that the performance of the maximum likelihood estimator of the cointegrating vectors is little affected by non-normal errors and (b) nonnormality is not an appropriate criteria upon which to select a lag length. Thus, choosing a higher order lag structure to eliminate non-normality may result to obtaining misleading results regarding the number of cointegrating vectors due to the decrease of power and the loss of inference. The crucial criterion for selecting lag length in the VAR specification is that lag length that eliminates serial correlation in the residuals. Lee and Tse (1996) have shown similar results when conditional heteroskedasticity is present.

Table 7 provides three test statistics that are important for assessing the statistical properties of the chosen model. The first test statistic tests for long-run exclusion of each variable that is included in the cointegrating vector and it provides a check of the adequacy of the chosen measurements and shows that none of the variables can be excluded from the cointegration space. These results imply that all four indices enter the cointegrating relation and we do not have the case that some of the indices are cointegrated between them and carry over this property to all other stock indices. This test is equivalent to testing for statistical significance of the estimated coefficients of the cointegrating vector. Indeed, Table 8 reports the estimated coefficients of the cointegrating relationship (normalized with respect to the CSE index) and we observe that they are all statistically significant. These results have implications concerning the benefits from international portfolio diversification between the four stock markets. The second statistic is a stationarity test and indicates that none of the indices can be considered stationary for any reasonable choice of r . Finally, we test for weak exogeneity to the long-run parameters. The results show that the ASE, LSE and NYSE price indices are weakly exogenous for the long-run parameters β , while for the CSE price index the null hypothesis of weak exogeneity is strongly rejected. These results are further confirmed by the second part of Table 2 which report the estimated speed of adjustments (α) of returns to the long-run equilibrium and it is clear that the estimated α 's for the ASE, LSE and NYSE stock indices are not significant while that for the CSE is statistically significant. These last findings take us to the discussion of the common trends.

4.3. Common trends analysis

Given this robust evidence for one cointegrating vector we now move on to the analysis of the three common trends which drive the system. We test whether the group of variables that are weakly exogenous enter the combination of the three common trends significantly. Table 9 reports the calculated estimated L -statistic proposed by Gonzalo and Granger (1995) based on the null hypothesis given by (8).

In this case the matrix K is given as follows:

$$K = \begin{bmatrix} 0 & 0 & 0 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \quad (10)$$

The results show that the null hypothesis $H_0 : \alpha_{\perp}^{ASE} = \alpha_{\perp}^{LSE} = \alpha_{\perp}^{NYSE} = 1$ could not be rejected at the five percent level of significance since the value of the L -statistic is equal to 3.51 with a p -value of 0.319. This implies that the three indices of ASE, LSE and NYSE enter the three common trends significantly. Therefore, they jointly contribute to these common trends and it is clear that these three capital markets are the driving forces for the system. These results are of course in line with the finding in the previous section that the estimated α 's for the ASE, LSE and NYSE stock indices are not significant while that for the CSE is statistically significant.⁸

What are the implications for international portfolio diversification that investors can achieve by including stocks in their portfolio from these four markets? Apparently, the benefits are of small magnitude, however, they do exist since the size of estimated α 's for the ASE, LSE and NYSE are of small size indicating that the

⁸In a relevant study Constantinou et al. (2006) show that when we examine volatility spillovers between the same four markets then we observe that ASE, LSE and NYSE are exporters of volatility to the CSE while CSE is importer of volatility with no effect to the other three markets.

returns may not react to the common trend(s) significantly and they take a long horizon to achieving this, which is partially in line with the statement by Kasa (1992) mentioned in section 2.

5. Concluding remarks

In this paper we used cointegration and common trends analysis to study the co-movements of four stock markets indices namely those of the Cyprus Stock Exchange (an emerging market) and the Athens Stock Exchange, London Stock Exchange and New York Stock Exchange (mature markets). The main issue under investigation is whether these four markets follow common trends and if so what are the implications in terms of benefits from international portfolio diversification. Evidence of cointegration and the existence of common trends implies that in the long-run any benefits from portfolio diversification are diminished. However, as Kasa (1992) points out even in the presence of cointegration there may still be benefits from portfolio diversification if the speed of adjustment (reaction) of the returns do not actually react to the common trend(s).

Our econometric analysis shows that there exists one cointegrating vector among the four equity markets and therefore three common trends. Then by analysing the Moving Average representation of the VAR system among the four stock indices we were able to draw statistical inference with respect to these three common trends. Thus, with the use of the Gonzalo and Granger (1995) methodology we test whether each index enters the common trend(s) significantly or not. We find that the ASE, LSE and NYSE stock indices are identified with each of the common trend or a combination of the common trends and they are the driving forces of the system, while the CSE stock index does not enter significantly the common trend.

Our overall findings suggest that the initial short-run evidence that there maybe potential benefits from portfolio diversification based on the correlations between the four markets does not carry over to the long-run. Therefore, it appears that there are small benefits from diversification in the long-run, possibly also due to the fact that the value of the speed of adjustments to the common trends is very small as Kasa (1992) argues.

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Table 1. Contemporaneous correlations between daily returns

	CSE	ASE	LSE	NYSE
CSE	1.0			
ASE	0.009	1.0		
LSE	0.037	0.047	1.0	
NYSE	0.028	-0.020	0.253	1.0

Notes: The stock returns are in nominal terms in domestic currency.

Table 2. Descriptive statistics

	CSE		ASE		FTSE		DJ	
	Levels	Returns	Levels	Returns	Levels	Returns	Levels	Returns
Mean	4.97	0.003	7.80	0.05	8.60	0.23	9.1	0.4
Standard Deviation	0.69	0.10	0.54	0.02	0.20	0.01	0.22	0.01
Skewness	1.05 *	7.60 *	-0.31*	-0.10	-0.76 *	-0.14 *	-0.84 *	-0.52 *
Kurtosis	0.20	354.1 *	-0.90 *	2.52 *	-0.55 *	0.95 *	-0.44 *	4.0 *
JB	284.5 *	7.9 X10 ⁶	76.5 *	400.9 *	164.8 *	62.2 *	195.0 *	1069.8 *
$Q(24)$	1560.7 *	2570.1 *	182.1 *	145.5*	192.9 *	100.0 *	199.1 *	141.0*
$Q^2(24)$	1670.7 *	1990.0 *	243.1 *	187.1*	199.1 *	143.9 *	122.0 *	191.1*

Notes:

1. The mean returns are expressed in terms of $\times 10^3$
2. LB(4) denotes the Ljung-Box statistic with 4 lags for the returns. This is distributed as χ^2_4 .
3. LBS(4) denotes the Ljung-Box statistic with 4 lags for the squared returns. This is distributed as χ^2_4 .
4. JB denotes the Jarque-Bera normality test. This is distributed as χ^2_2 .
5. * denotes statistical significance at the 5% level.

Table 3. Unit root and stationarity tests

Variables	DF-GLS _u		MZ_a^{GLS} MZ_t^{GLS}		KPSS	
	t_μ	t_τ			η_μ	η_τ
$\ln cse$	-0.60 [4]	-0.34 [4]	-0.14 [1]	-0.15 [1]	2.251*	0.619*
$\Delta \ln cse$	-16.75* [3]	-16.63* [3]	-424.52* [3]	-14.56* [3]	0.221	0.136
$\ln ase$	-0.15 [1]	-0.56 [1]	-0.74 [4]	-0.65 [4]	2.883*	1.061*
$\Delta \ln ase$	-31.93* [0]	-32.49* [0]	-753.13* [0]	-19.40* [0]	0.172	0.117
$\ln ftse$	-0.19 [2]	-0.85 [2]	-0.20 [2]	-0.19 [2]	2.584*	1.121*
$\Delta \ln ftse$	-5.85* [11]	-7.19* [11]	-23.37* [11]	-3.39* [11]	0.306	0.036
$\ln dj$	0.28 [0]	-1.33 [0]	0.26 [0]	0.28 [2]	3.771*	1.054*
$\Delta \ln dj$	-3.96* [11]	-28.92* [11]	-14.93* [11]	-2.70* [11]	0.160	0.024

Notes: e_o and e_n are, respectively, the official and parallel exchange rate. Δe_o and Δe_n are the first differences.

- The DF-GLS_u is due to Elliot et al. (1996) and Elliott (1999) is a test with an unconditional alternative hypothesis. The standard Dickey-Fuller tests are detrended (with constant or constant and trend). The critical values for the DF-GLS_u test at the 5% significance level are: -2.73 (with constant) and -3.17 (with constant and trend), respectively (Elliott, 1999).
- MZ_o and MZ_n are the Ng and Perron (2001) GLS versions of the Phillips-Perron tests. The critical values at 5% significance level are: -8.10 and -1.98 (with constant), respectively (Ng and Perron, 2001, Table 1).
- η_μ and η_τ are the KPSS test statistics for level and trend stationarity respectively (Kwiatkowski *et al.* 1992). For the computation of these statistics a Newey and West (1994) robust kernel estimate of the "long-run" variance is used. The kernel estimator is constructed using a quadratic spectral kernel with VAR(1) pre-whitening and automatic data-dependent bandwidth selection [see, Newey and West, 1994 for details]. The 5% critical values for level and trend stationarity are 0.461 and 0.148 respectively, and they are taken from Sephton (1995, Table 2).

(*) indicates significance at the 95% confidence level.

Table 4: Philips-Perron tests for cointegration: P_z -test

$$S_{CSE} = \alpha + \beta S_{index} + u$$

Index	$\hat{\alpha}$	$\hat{\beta}$	R^2	P_z -test
ASE	-2.74 (-15.96)	0.99 (45.11)	0.59	106.80*
LSE	-18.04 (-25.73)	2.68 (32.85)	0.40	112.30*
NYSE	15.15 (-26.37)	2.21 (35.08)	0.43	100.30*

Notes:

1. The P_z -test is invariant to the normalisation of the cointegrating regression (Phillips and Ouliaris, 1990, p. 172).
2. The asymptotic distribution of P_z -test is free of nuisance parameters and is dependent only on the number of explanatory variables in the cointegrating regression (Phillips and Ouliaris, 1990, p. 172). As the asymptotic distribution is free of nuisance parameters, the critical values reported in Phillips and Ouliaris (1990) are valid.
3. The null hypothesis is no cointegration. If the computed value of the P_z -statistic is greater than the critical value then we reject the null of no cointegration. The 5% critical value is 55.2202 (Phillips and Ouliaris, 1990, Table V).
4. The reported test statistics are based on a lag window of 2. Alternative lag windows yield qualitatively similar results.
5. Numbers in parenthesis are t-statistics calculated from the estimated corrected West standard errors.

Table 5. Residual misspecification tests of the model with $k = 9$

Eq.	σ_ε	LB(36)	ARCH(9)	η_3	η_4	NORM(6)	R^2
Δp	0.002	22.11	50.58*	0.36	1.16	12.40*	0.696
Δi	0.045	25.72	11.06	-0.83	3.81	45.07*	0.161
Δe	0.028	33.65	12.84*	-0.01	0.62	5.18	0.189
Δp^f	0.001	29.00	10.04	-0.21	1.70	23.43*	0.574
Δi^f	0.040	27.67	3.14	-0.65	2.94	37.16*	0.131

Notes: σ_ε is the standard error of the residuals, η_3 and η_4 are the skewness and kurtosis statistics. LB is the Ljung-Box test statistic for residual autocorrelation, ARCH is the test for heteroscedastic residuals, and NORM the Jarque-Bera test for normality. The ARCH and NORM statistics are distributed as χ^2 with 9 and 6 degrees of freedom, respectively and the LB statistic is distributed as χ^2 with 36 degrees of freedom. *(**) denotes significance at the 5% (1%) level.

Table 6. Johansen - Juselius likelihood cointegration tests

r	5% Critical Values			
	Trace	λ_{\max}	Trace	λ_{\max}
R=0	49.37*	30.00*	47.85	27.58
R=1	19.16	10.79	29.78	21.13
R=2	8.36	6.35	15.49	14.26
R=3	2.00	2.00	3.84	3.84

Notes: r denotes the number of eigenvectors. Trace and λ_{\max} denote, respectively, the trace and maximum eigenvalue likelihood ratio statistics. The 5% critical values are taken from MacKinnon *et al.* (1999; Table IV). A structure of four lags was chosen according to a likelihood ratio test, corrected for the degrees of freedom (Sims, 1980) and the Ljung-Box Q statistic for detecting serial correlation in the residuals of the equations of the VAR. A model with an unrestricted linear trend in the VAR equation and a constant restricted in the cointegrating vector is chosen according the Johansen (1992 a, b) testing strategy .

(*) denotes statistical significance at the five percent critical level.

The roots of the companion matrix

Modulus of 5 largest roots

Eigenvalues	0.018	0.007	0.004	0.001
Unrestricted model	0.99	0.95	0.95	0.93
r = 1	1.00	1.00	1.00	0.73

Notes: The table shows the modulus of the estimated $p \times k$ roots of the companion matrix from the VAR system, p is the number of variables and k is the number of lags of the VAR. We report the first five roots which are of our interest.

Table 7. Statistical Properties and Misspecification Tests of the Model*(a) Tests for long-run exclusion, stationarity, and weak exogeneity*

	Long-run exclusion	Stationarity	Weak exogeneity
i_{cse}	15.64*	13.42*	16.98*
i_{ase}	7.34*	13.94*	0.55
i_{lse}	10.45*	11.43*	2.06
i_{nyse}	22.24*	13.50*	2.19

Notes: The long-run exclusion restriction tests for the null hypothesis ($\beta_i = 0$) and the weak exogeneity tests for the null hypothesis ($a_i = 0$). Both tests are χ^2 distributed with one degree of freedom and the 5% critical level is 3.84. The stationarity test is a χ^2 distributed with four degrees of freedom and the 5% critical level is 9.49. An (*) denotes statistical significance at the 5 percent critical level.

(b) Multivariate Residuals Diagnostics

L-B(60)	LM(1)	LM(4)	χ^2 (10)
1517.67(0.01)	17.46(0.86)	30.09(0.22)	81.03(0.00)

Notes: L-B is the multivariate version of the Ljung-Box test for autocorrelation based on the estimated auto- and cross - correlations of the first $[T/4=60]$ lags distributed as a χ^2 with 1400 degrees of freedom. LM(1) and LM(4) are the tests for first and fourth-order autocorrelation distributed as χ^2 with 25 degrees of freedom and χ^2 is a normality test which is a multivariate version of the Shenton-Bowman test (1977) test modified in Doornik and Hansen (1994). Numbers in parentheses refer to marginal significance levels.

Table 8. Estimated Coefficients and Hypothesis Testing

$$i_{cse} = \beta_0 + \beta_1 i_{ase} + \beta_2 i_{lse} + \beta_3 i_{nyse}$$

Eigenvectors				Adjustment Speeds			
β_0	β_1	β_2	β_3	a_{cse}	a_{ase}	a_{lse}	a_{nyse}
-11.23 (1.36)	-4.477 (0.05)	5.592 (0.34)	3.362 (0.42)	0.0157 (0.0031)	-0.0029 (0.0034)	0.0030 (0.0018)	0.0031 (0.0019)

Notes: Numbers in parentheses report standard errors. The eigenvector is normalized with respect to the CSE price index.

Table 9. Testing for linear combinations on the common trends

$H_0 : a_{\perp} K$	$LR - stat.$	p -value	$(p - r)x(p - m)$
$\alpha_{\perp}^{ASE} = \alpha_{\perp}^{LSE} = \alpha_{\perp}^{NYSE} = 1$	3.51	0.319	3

Notes: H_0 : The group of indices contributes significantly to the common trend(s). This is a L -test statistic distributed as $\chi^2_{(p-r)x(p-m)}$, where p is the number of variables, r is the number of cointegrating vectors and m is the number of common trends. (*) denotes significance at the 5% critical level.







