Cointegration, causality and domestic portfolio diversification in the Cyprus Stock Exchange

by

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Abstract

In this paper we provide an investigation on the potential benefits that may exist for portfolio managers, private and institutional investors from domestic portfolio diversification. We employ daily data for the period 1996-2002 from the Cyprus Stock Exchange, recently established emerging market. Cointegration as well as linear and nonlinear causality analysis is used in order to reveal whether there are benefits from domestic portfolio diversification. The cointegration analysis leads to the conclusion that we are unable to reject the null hypothesis of no cointegration in most bivariate cases of the 56 pairs of sectoral indices and this finding is taken to imply that the are benefits from portfolio diversification, when domestic investors construct portfolios which include stocks from the sectors which are not cointegrated. Furthermore, the application of linear and nonlinear Granger causality leads to a pattern of causality between these pairs of sectoral indices which is almost identical and therefore the linearity hypothesis is rejected. Furthermore, based on our causality analysis we provide evidence that traders and investors in the CSE set up short-run investment strategies. Moreover, this implies that the Cypriot investors do not adopt contrarian and momentum investment strategies. Therefore, we argue that the investors in the Cyprus stock market exhibit myopic investment behaviour.

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1. Introduction

Over the last fifteen years there has been a growing interest among portfolio managers for the emerging capital markets as they provide opportunities for higher asset returns compared to those of the developed markets. This was caused by the substantial increase of capital flows from the mature markets to the emerging markets of the South East Asia and the economies of transition of Central and Eastern European countries. The purpose was to invest in portfolios consisting to a great extent with securities from these new financial markets. Indeed, the study by Singh and Weisse (1998) reports that, during the period 1989-1995 the inflow of funds in emerging markets amounted to 107.6 billion US dollars as opposed to a mere 15.1 billion US dollars in the previous period 1983-1988. However, in the aftermath of the financial crisis in Southeast Asia, Latin America and Russia in 1997-1998, we have experienced a substantial increase in financial uncertainty as a result of the increased volatility that stock returns of the mature markets but mainly of those of the emerging markets exhibited.

Given these stylized facts academics and practitioners became aware of the importance that a thorough study of the new markets had for portfolio managers and institutional investors. Research has been focused on two independent issues which at the same time belong to the class of issues that time series analysis encompasses, namely existence of long-run relationships between stock markets (especially between emerging markets and emerging and mature markets) and modeling the volatility of stock returns in the emerging markets.

The main aim of the present paper is to investigate whether there are long-run benefits from domestic portfolio diversification for the Cypriot investors who invest in the Cyprus Stock Exchange (CSE) and what is the causality direction among the sectors of the Cypriot economy whose firm equity are traded in this capital market. Specifically, we examine the case of the Cypriot investor who invests on CSE and he/she is interested for the stocks whose prices are expressed in Cyprus pounds.[1] To achieve our target we adopt the framework of cointegration theory and we statistically examine whether cointegration exists for the following cases. First, between the general price index and the volume of transactions. Second, among the twelve sectoral price indices and finally between the sectoral indices in a bivariate framework giving rise to 56 pairs to be examined.

Recent studies that deal with the issue of the multivariate analysis of the relationship among the stock of different stock markets have applied cointegration theory with the purpose of studying the long-run properties of stock prices. Most of these studies have reached the conclusion that there is cointegration between two stock prices or between two stock price indices and they have interpreted this finding as evidence that there is a long-run linkage and therefore a long-run relationship.[2] Equivalently, we could argue that the existence of cointegration between two or more stock prices this could be seen as evidence of long-run relationships between these series. With respect to the issue of portfolio diversification, existence of cointegration between two or more stock prices implies that in the long run these prices are moving together and therefore, the benefits from diversification with the construction of a portfolio that consists of these stocks are limited. In contrast, lack of cointegration implies that there are significant long-run benefits from the reduction of risk without loss in the expected returns.

[1] Therefore, in the present study we abstract from the issue of exchange rate volatility and the problems which are tied with the international portfolio diversification.

[2] Taylor and Tonks (1989), Arshanapalli and Doukas (1993), Byers and Peel (1993), Kasa (1992), Richards (1995), Kanas (1998, 1999) και Georgoutsos and Kouretas (2003).

Since the early 1990s economists have paid attention to the analysis of mature and emerging markets by applying cointegration theory in order to confirm whether benefits from international portfolio diversification exist. Arshanapalli and Doukas (1993) focused their analysis on the capital markets of the U.K., France, Germany and U.S.A. and they performed statistical tests for the existence of bivariate cointegration between the U.S. stock market with each of the other major markets. For the U.S. investor this study finds out that there exists cointegration for all potential pairs and therefore there are small benefits from the international portfolio diversification of the American investors. Contrary to these findings, Taylor and Tonks (1989) found no evidence of cointegration between the U.S. and the U.K. stock markets. Byers και Peel (1993) as well as Kasa (1992) studied the case of three European of three European stock markets and the those of Japan and Canada and they reached the conclusion that there is partial evidence of cointegration. Kanas (1998) examined the case of the six largest stock markets vis-à-vis the NYSE and found no evidence in favour of cointegration leading to the conclusion that there may be substantial benefits from the international portfolio diversification. Other studies like Gallagher (1995) and DeFusco *et al*. (1996) which analyse the case of some of the major European stock markets also confirmed that there is no evidence of cointegration between the stock markets of Ireland, Germany and the U.K. 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 examined the long-run interrelationships of the European capital markets and they concluded that although cointegration exists the gains from diversification are short-lived since the adjustment to the common trend is very slow. Finally, Georgoutsos and Kouretas (2003) analyzing the major stock markets argued that the most significant cause for the lack of evidence in favour of cointegration among them is the rejection of the long-run Purchasing Power Parity.

We also provide a Granger causality analysis between the 12 sectors of the Cyprus economy which are included in the stock market in a bivariate context as we do for the cointegration analysis. Cointegration analysis examines whether a long-run relationship between two or more variables exists or not. Granger causality analysis is adopted in order to investigate the causal dynamic relationships between the same set of variables. We first conduct a linear Granger causality between the first differences for every pair of the sectoral price indices by estimating a VAR model in each case. We then employ the corrected statistical criterion due to Baek and Brock (1992) in order to conduct a non-linear Granger causality analysis. We do that in order to examine whether the results of the linear causality analysis depend of the linearity hypothesis or not (robustness test).

The rest of the paper is organized as follows. Section 2 presents the institutional and functional framework of the Cyprus Stock Exchange. Section 3 presents and discusses the data and preliminary empirical results. Section 4 discusses cointegration analysis and the obtained results. Finally, section 5 provides the summary and the concluding remarks.

2. The Cyprus Stock Exchange

The Cyprus Stock Exchange is the primary stock market in Cyprus. It is considered to be a small emerging capital market with a very short history since it was established in April 1993 when the inaugural Stock Exchange Law passed through the Cypriot House of Representatives. In July 1995 the Cypriot House of Representatives passed the laws for the stock exchange function and supervision, while additional laws led to the establishment of the Central Securities Depository. On 29 March 1996 the first day of transactions took place. The Cyprus Stock Exchange S.A. is supervised by the Ministry of Finance and the Minister of Finance is responsible for choosing the seven member executive committee that runs CSE. Furthermore, the Securities and Exchange Committee is mostly responsible for the well functioning of the capital market of Cyprus. Trading takes place electronically through the Automated Trade System. The main index is the CSE General Price Index that reflects, approximately, 93% of the trading activity and 96% of the overall capitalization. In November 2000 the FTSE/CySE 20 was constructed with the cooperation of CSE, the Financial Times and the London Stock Exchange in order to monitor closer the market. To highlight the increasing need for regional capital market integration the FTSE Med 100 was created in June 2003 with the cooperation of CSE, ASE and the Tel-Aviv Stock Exchange. Figure 1 shows the evolution of the CSE general price index and its returns, (Chisostomidou *et al*. 2006 provide a comprehensive analysis of the institutional framework of CSE).

The Cyprus accession in the European Union on 1st May 2004 it also determines the starting period within which all required changes for the adjustment of the financial system and the operation of the financial markets of Cyprus in order to become an integral part of the European financial system and the European financial markets. The capital market of Cyprus seeking to achieve its primary goal which is the efficient allocation of sources in their alternative uses in the production and investment process it has to adjust its institutional framework of operation by fully incorporated all procedures that regulate the European capital markets.

The forthcoming European financial market integration implies that the benefits from international portfolio diversification within Europe may be substantially reduced and this provides a further motivation of the present study since we are seeking for exploring intra-firm benefits from domestic portfolio diversification. This argument provides a further motivation for the present study.

3. Data and preliminary empirical results

Our data consists of daily observations for the period 29 March 1996 (first day of operation of the CSE) to 19 April 2002, excluding all weekends, holidays and days during which the CSE was closed. The final sample consists of 1444 observations. The data has been taken from DATASTREAM and all series have been transformed to natural logarithms. The variables used for the present analysis are the following: The general stock price index, Banks, Construction companies, Fisheries and Fishtrading companies, Investment companies, Manufacturing, Insurance, Hotels, Tourist services, Real Estate, Informatics, Financial Services, Other companies. We also use the volume of transactions.

We begin by providing a discussion of the financial developments in the operation of the CSE. We divide the time period of the operation of CSE in three periods. The first period (29/03/1996-30/06/1999) is characterized by low interest by domestic and foreign investors, thin trading as well as low volatility and persistence of the general price index around the 100 units. The second period (01/07/1999-31/10/2000) is characterized by the presence of a rational bubble. The stock market bubble is an expected event in new and emerging markets, like the Cyprus capital market, although this is a phenomenon that often appears in mature markets as well. For the Cyprus case the appearance of the rational bubble was the result of the sudden increased interest of a great number of domestic investors who, as it is always the case in the new and emerging markets were not well informed about the workings of this new market and the trading process in the CSE. We could partially attributed the presence of the rational bubble to the rational bubble that existed in the Athens Stock Exchange which is a market whose developments influence the investors' behaviour in the Cyprus capital market. This event occurred in a year before the one in Cyprus and during its upward trend made domestic investors to believe that there will be a continuous increase in stock prices. However, it is well known that there comes a time period that the bubble burst leading to panic and continuous fall in the stock prices. The leading view about the creation of a bubble in a financial market is that it occurs when the price of financial instrument (stock price, exchange rate, etc.) deviates substantially and systematically from the fundamentals (either of the corresponding firm or the economy) for long time periods. The bubble in the CSE lasted for one and a half years. Looking at the third and final period (01/11/2000 – 19/04/2002) we clearly see that the general price index of CSE has gradually returned to its initial level and for a long period its value was around the baseline of 100 while at the present its value is a bit higher, steadily leading to reduced interest by investors and low volume of transactions. Figure 1 presents the evolution of the general price index as well as of the stock returns for the period under investigation. We clearly see the bubble during the period 01/07/1999 – 31/10/2000.

Table 1 reports the results from unit root and stationarity tests for the CSE general stock price index and its first difference in order to obtain a clear picture of the stochastic properties of the series. It also report these test statistics for the 12 sectoral indices and the volume of transactions. Specifically, in order to test for the presence of a unit root in the level of the series we apply a set of unit root tests developed by Elliott *et al*. (1996) and Elliott (1999) as well as by Ng and Perron (2001). These tests modify conventional ADF and Philips-Perron unit root tests in order to derive tests that have better size and power. The use of these recently developed tests lead to firmer conclusions with respect to the integration properties of the stock price series since rejections of the null hypothesis of nonstationarity will not be attributed to size distortions, whereas nonrejections is not the outcome of a low probability of rejecting a false null hypothesis. The null hypothesis for the Elliot *et al*. (1996) GLS augmented Dickey-Fuller test (DF-GLS$_u$) and Ng and Perron (2001) GLS version of the modified Phillips-Perron (1988) tests (MZ_a^{GLS} and MZ_t^{GLS}) is that of a unit root against the alternative that the initial observation is drawn from its unconditional distribution. These tests use the GLS-detrending technique 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 methodology resulted to a substantial increase in power for the DF-GLS$_u$ test deriving 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 and they have also derived a modified version of the AIC criterion (MIC) that give rise to substantial size improvements over alternative selection rules such as BIC. Finally, we apply the Kwiatkowski *et al*. (1992) KPSS test for the null hypothesis of level or trend stationarity against the alternative of non-stationarity and these additional results will provide robust inference. The overall evidence for this set of tests is that all price indices as well as the trading volume are nonstationary while their first difference is a stationary process.

Provided that the stock price index is a nonstationary variable we only consider the first differences of the general price index:

$$\Delta p_t = 100 * (p_t - p_{t-1}) \text{ and } \Delta q_t = 100 * (q_t - q_{t-1}) \qquad (1)$$

which corresponds to the approximate percentage nominal return on the stock price series obtained from time t to *t-1*. Specifically, the daily returns have been calculated by taking the first difference of the logarithms of two consecutive days. By the same token the percentage daily changes in the volume of transactions is 100 times the first difference of the logarithms of two consecutive trading values

Table 2 reports several descriptive statistics for the returns of the general index as well as of the sectoral indices of the CSE. The descriptive statistics include the mean, the variance, the asymmetry and kurtosis of the distribution of stock returns. According to Table 2 almost all series exhibit asymmetry and kurtosis since the respective statistics are statistically significant leading to the conclusion that we observe statistically significant deviations from normality. Further analysis of these descriptive statistics show that the stock market of Cyprus is not efficient which is expected given that this market is an emerging one. These results are of importance when we use the VAR models in order to examine the long-run properties of the stock prices as well as when we conduct the linear and nonlinear Granger causality analysis for the stock returns. Finally the value of the Q^2 statistic is statistically significant which implies that there is 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 different stock prices and we should take that into consideration when we employ the non-linear models for Granger causality.

4. Cointegration analysis

To conduct the cointegration analysis between the sectoral indices we apply the well known multivariate cointegration methodology developed by Johansen (1988, 1991) and Johansen and Juselius (1990, 1992). Given the numerous application of this methodology we only provide a short description of it.[3]

Johansen's methodology (1988, 1991) is based on the estimation of an autoregressive (VAR) system of n x 1 vectors of nonstationary variables X_t.

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

where z_t is a column vector of stochastic variables, $\varepsilon_t \sim Nid_p(0,\Sigma)$. The parameters $(\Gamma_1,.........,\Gamma_{k-1},\gamma)$ 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:

$$\Pi = \alpha\beta' \qquad (2)$$

where Π has reduced rank, r, α and β are $(p\mathrm{x}r)$ matrices, and

[3] For a comprehensive and rigorous presentation of this methodology see Hamilton (1994).

$$\Psi = \alpha_{\perp}(-I + \Gamma_1)\beta_{\perp} = \varphi\eta' \quad (3)$$

where Ψ is a $(p-r)$x$(p-r)$ matrix of full rank, φ and η are $(p-r)$x$(p-r)$ matrices, and $\alpha_{\perp}$ and $\beta_{\perp}$ are px$(p-r)$ matrices orthogonal to α and β, respectively. The parameterization in (2) and (3) 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.

Using the maximum likelihood estimation method, Johansen developed two statistical criteria to test the null hypothesis of no cointegration. The first test is the maximum eigenvalue test and the second test is the Trace test.[4] We apply the multivariate cointegration methodology due to Johansen (1988, 1991) on the system given by (1), while the lag structure of the system is determined with the use of a Likelihood Ratio (LR) test developed by Sims (1980). In case that there are more than two variables z_t (n > 2), then there exists the likelihood that exist more than one linear stationary cointegration vectors of the non-stationary time series. In principle, the larger is the number of cointegration vectors the greater is the probability that a long run relationship exists between the series.

We begin the cointegration analysis by examining the pair of the general price index and the volume of transactions (GEN-VOL). Both the maximum eigenvalue and

[4]The algebraic expressions of these two statistical criteria are given in Hamilton (1994). The critical values of these two statistical criteria have been calculated by Osterwald-Lenum (1992) and extended by MacKinnon *et al*. (1999).

the trace test statistics could not reject the null hypothesis of no cointegration and therefore, we argue that there does not exist a long-run relationship between the general price index and the volume of transactions in the CSE. The lack of cointegration between these two variables does not necessarily imply the lack of Granger causality, an issue which is examined in the next section. The results are given in Table 3.

We next move to the cointegration analysis of the 12 sectoral indices within a multivariate framework. Table 4 reports the results. Both statistics reject the null hypothesis of no cointegration among the 12 sectors and thus, we argue that there is at least one long-run relationship between the 12 sectoral indices. Specifically, with the trace test we find nine cointegration vectors, while on the basis of the maximum eigenvalue we confirm the existence of at most five cointegration vectors. Therefore, given our discussion above, the fact that there exists a large number of cointegration vectors leads to the conclusion that we can have at least one statistically significant long-run relationship between the 12 sectoral indices. This finding is consistent with our intuition that since all the sectors of the economy of Cyprus are subject to a greater or lesser extent to common disturbances like macroeconomic disturbances (inflation level, interest rate level, tax rates, monetary policy) as well as to various political events. This statistically significant long-run relationship between the 12 sectoral indices implies that there are no benefits from portfolio diversification in terms of reduction in risk for that portfolio which also includes stocks from these 12 sectors. Based on this evidence we can investigate the case of potential combinations of portfolios which include stocks from some of the sectors and which could give to the domestic investors benefits from the diversification.

Table 5 reports the results of the bivariate cointegration analysis for all 56 pairs of sectoral indices of the CSE. According to the cointegration analysis we observe that there are indeed possible cases in which portfolio diversification can result to long-run benefits for the Cypriot investor. Specifically, the banking index BANK appears to have a cointegrating vector with all other sectors besides BUILD, INSUR, MANUF and REALESTATE. Based on these findings we could achieve a risk reduction (without reduction of the expected returns), in case we construct a portfolio which includes stocks of the banking sector with stocks either of the building sector, the insurance sector, the manufacturing sector or the realestate sector. Furthermore, we conclude that the building sector, BUILD, has no long-run relationship with any of the other 11 sectoral indices and this implies that there will be long-run benefits from the portfolio diversification. Respectively, the index of the financial services companies, FINSERV as well as that of the fisheries companies, FISH, appear to have a statistically significant cointegration vector only with the banking sector index each one of them. Hence, any portfolio which includes stocks from either the financial services sector or the fisheries sector and any of the other sectors besides that of the banking sector will provide the opportunity to the investors opportunities to accrue benefits from the reduction of risk.

We also observe that index FIN is cointegrated with the banking index, BANK and the index of hotels, HOTELS and therefore any portfolio of two assets which include stocks from the financial sector FIN, should not also include stocks from the banking sector and from the hotels sector on the basis of the benefits criterion of portfolio diversification. Furthermore, we find that there exists a long-run relationship between the index of the tourist services sector, HOTELCOMP, and the index of the manufacturing and banking sectors. Therefore, a combination of the index

HOTELCAMP, and either of these two indices will not offer long-run benefits to the domestic investors from diversification of their portfolios. The sectoral index of hotels, HOTELS, has a statistically significant cointegration vector with the index of all other companies, MISC, as well as with the index of the banking sector, whereas the index of the informatics, INFORM, has a long-run comovement with the banking sector and the realestate sector. Finally, the index of the insurance sector, INSUR, is found to have a statistically significant cointegrating vector with the manufacturing sector and with the realestate vector.

To summarise, we argue that there exists a rich variety of results stemming from our cointegration analysis with respect to the behaviour and predictability of the general price index and all 12 sectoral indices. Our bivariate systems lead to the conclusion that the CSE offers the opportunity for making long-run profits from the portfolio diversification.[5]

5. Linear and nonlinear Granger causality analysis

Cointegration analysis examines whether or not a long-run relationship between two or more variables. Granger causality analysis is adopted in order to investigate the existence of causal dynamic relationships between the same variables.

The linear Granger causality analysis is conducted by regressing the first differences between two sectoral indices at a time through the estimation of a VAR model. The complete set of results is presented in Table 6. Given the large number of cases and therefore the extensive length of results we provide a summary of them. According to the estimated pairwise cases we observe that in most cases we are

[5] To save space we only report the quantitative results given that there is an extensive set of tables for all 56 bivariate cases. These tables are available upon request.

unable to reject the null hypothesis of no-Granger causality between any two pairs of sectoral indices.

With the purpose to examine whether the results obtained from the linear Granger causality analysis is independent of the linearity hypothesis we also apply the non-linear Granger causality analysis (robustness test). The application of the non-linear Granger causality is based on the corrected statistical criterion developed by Baek and Brock (1992).

Consider two strictly stationary and weakly dependent scalar time series $\{W_t\}$ and $\{Z_t\}$. We denote the m-length lead vector of W_t with W_t^m, and the L_w-length and L_z-length lag vectors of W_t and Z_t, respectively, by W_{t-Lw}^{Lw} and Z_{t-Lz}^{Lz} :

$$W_t^m = (W_t, W_{t+1}, ..., W_{t+m-1}),\ m = 1, 2, ..., t=1, 2, ...$$

$$W_{t-Lw}^{Lw} = (W_{t-Lw}, W_{t-Lw+1}, ..., W_{t-1}),\ L_w = 1,2,..., t = L_w+1, L_w+2, \tag{4}$$

$$Z_{t-Lz}^{Lz} = (Z_{t-Lz}, Z_{t-Lz+1}, ..., Z_{t-1}),\ L_z = 1,2,..., t = L_z+1, L_z+2, ...$$

For given values of m, L_w, and $L_z \geq 1$ and for $e > 0$, Z does not strictly Granger cause W if

$$\Pr\{\| W_t^m - W_s^m \| < e \mid \| W_{t-Lw}^{Lw} - W_{s-Lw}^{Lw} \| < e, \| Z_{t-Lz}^{Lz} - Z_{s-Lz}^{Lz} \| < e \} =$$

$$= \Pr\{\| W_t^m - W_s^m \| < \mathbf{e} \mid \| W_{t-Lw}^{Lw} - W_{s-Lw}^{Lw} \| < e \} \tag{5}$$

where Pr{.} denotes probability and || . || denotes the maximum norm. The probability on the left hand side of equation (5) is the conditional probability that two arbitrary m-length lead vectors of $\{W_t\}$ are within a distance e of each other, given that the corresponding L_w-length lag vectors of $\{W_t\}$ and L_z-length lag vectors of $\{Z_t\}$ are within a distance e of each other. The probability on the right hand side of equation

(5) is the conditional probability that two arbitrary m-length lead vectors of $\{W_t\}$ are within a distance e of each other, given that their corresponding L_w-length lag vectors are within a distance e of each other. It can be shown[6] that, given values for m, L_w, L_z and $e > 0$, under the null hypothesis that $\{Z_t\}$ does not strictly nonlinearly Granger cause $\{W_t\}$, the statistic

$$\sqrt{n}\left\{\frac{C1(m+L_w, L_z, e, n)}{C2(m+L_w, e, n)} - \frac{C3(m+L_w, e, n)}{C4(L_w, e, n)}\right\} \sim AN\left(0, \sigma^2(m, L_w, L_z, e)\right) \qquad (6)$$

where $C1(m+L_w, L_z, e, n)$, $C2(m + L_w, e, n)$, $C3(m+L_w, e, n)$, and $C4(L_w, e, n)$ are correlation-integral estimators of the point probabilities corresponding to the left hand side and right hand side of equation (5). This test has remarkably good power properties against a variety of nonlinear Granger causal and noncausal relations, and its asymptotic distribution is the same if the test is applied to the estimated residuals from a vector autoregressive (VAR) model (Hiemstra and Jones, 1994).

In carrying out the modified Baek and Brock tests, values for the lead length m, the lag lengths L_w and L_z, and the scale parameter e must be chosen. On the basis of the Monte Carlo results of Hiemstra and Jones (1994), we set for all cases, m=1, $L_w = L_z$ using common lag lengths of 1 to 5 lags. Moreover, for all cases, we set $e = 1.0\sigma$, where $\sigma = 1$ denotes the standard deviation of each series.

The results of the non-linear Granger causality analysis are presented in Table 7 for all bivariate cases of the sectoral indices. The overall findings lead to the conclusion that the causality direction is almost identical to the one found in the linear

[6] For more details on the derivations, see Hiemstra and Jones (1994).

Granger causality analysis and therefore we argue that the causality direction is given and it does not depend on the linearity hypothesis.

More specifically, in the case of the nonlinear Granger causality we observe we were able to reject the null hypothesis of no causality between the indices in few cases. The fact that in only few cases we found existence of linear or nonlinear Granger causality between the sectoral indices can provide some interpretations with respect to the behaviour of the investors who are active in the stock market of Cyprus. Specifically, it is clear that there are no short run dynamic interrelationships between the indices. This finding implies that overtime the sectoral indices are independent. Furthermore, this evidence leads to the conclusion that traders and investors in the CSE set up short-run investment strategies. Moreover, this implies that the Cypriot investors do not adopt contrarian and momentum investment strategies. Therefore, we argue that the investors in the Cyprus stock market exhibit myopic investment behaviour. [7]

6. Summary and concluding remarks

In this paper we provide a comprehensive analysis of the potential benefits that may be realized from domestic portfolio diversification. Specifically, we use daily data for the period 1996-2002 for the Cyprus Stock Exchange a recently established emerging market.

We employ two well known econometric methodologies to accomplish our aim. First, we use the Johansen (1988, 1991) and Johansen and Juselius (1990, 1992) multivariate cointegration methodology to examine whether there are long-run relationships among the 12 sectors of the Cyprus economy. Looking into the

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[7] To save space we only report the quantitative results given that there is an extensive set of tables for all 56 bivariate cases. These tables are available upon request.

relationships between sectors can also be justified on the grounds that the recent enlargement of the European Union may eventually result to minimum benefits from international portfolio diversification whereas substantial benefits from risk reduction due to exercise of domestic portfolio diversification may still existed. We also provide a linear and nonlinear Granger causality to reveal any short-run dynamics between the sectoral indices.

Our cointegration analysis provided us with a rich variety of results with respect to the behaviour and predictability of the general price index and all 12 sectoral indices. First, we found no cointegration between the general price index and the volume of transactions. Second, within a multivariate context we showed that there is at least one statistically significant long-run relationship between the 12 sectoral indices. Based on this finding, we finally examine all bivariate systems of sectoral indices and we can conclude that the CSE offers the opportunity for making long-run profits from the portfolio diversification

The linear and nonlinear Granger causality analysis has led to very similar pattern of causality with only few cases of causality between the bivariate cases of all sectoral indices. Therefore, the linearity hypothesis was rejected while it is clear that there are no short run dynamic interrelationships between the indices. This finding implies that overtime the sectoral indices are independent. Furthermore, this evidence leads to the conclusion that traders and investors in the CSE set up short-run investment strategies. Moreover, this implies that the Cypriot investors do not adopt contrarian and momentum investment strategies. Therefore, we argue that the investors in the Cyprus stock market exhibit myopic investment behaviour.

The results of the present paper are particularly useful to private and institutional investors as well to the financial institutions, for the evaluation and management of their portfolios which include stocks of companies which are listed in the CSE. These results are also useful to the pension funds (when they will be allowed to invest part of their reserves in stocks traded in the CSE), to the insurance companies and to the mutual funds (whose establishment and introduction to the CSE is expected).

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Table 1. Unit root and stationarity tests

A. Levels

Variables	DF-GLS$_u$				KPSS	
	t_μ	t_τ	MZ_a^{GLS}	MZ_t^{GLS}	η_μ	η_τ
GENERAL INDEX	-0.60 [4]	-0.34 [4]	-0.14 [1]	-0.15 [1]	2.251*	0.619*
VOLUME	-0.41 [5]	-0.51 [5]	-0.37 [3]	-0.57 [3]	0.889*	0.736*
BANK	-0.22 [1]	-0.31 [1]	-0.59 [2]	-0.63 [2]	2.333*	2.261*
BUILD	-0.41 [0]	-0.38 [0]	-0.98 [2]	-1.03 [2]	1.167*	1.127*
FIN	-0.21 [2]	-0.77 [2]	-0.20 [4]	-0.19 [4]	1.098*	1.111*
FINSERV	-1.41 [1]	-1.11 [1]	-1.19 [3]	-1.03 [3]	1.209*	1.034*
FISH	-0.98 [2]	-1.12 [2]	-0.34 [3]	-0.29 [3]	2.145*	0.786*
HOTEL COMP	-0.43 [2]	-0.56 [2]	-0.78 [4]	-0.54 [4]	0.989*	0.546*
HOTELS	-0.91 [0]	-0.88 [0]	-1.16 [1]	-1.02 [1]	1.333*	0.338*
INFORM	-1.46 [4]	-1.55 [4]	-0.88 [3]	-0.89 [3]	1.654*	0.790*
INSUR	-0.29 [1]	-0.56 [1]	-0.78 [2]	-0.68 [2]	1.335*	1.003*
MANUF	-1.02 [3]	-1.12 [3]	-0.34 [0]	-0.45 [0]	1.991*	0.675*
MISC	-0.78 [4]	-0.63 [4]	-0.23 [5]	-0.37 [5]	1.565*	0.454*
REAL ESTATE	-0.33 [6]	-0.67 [6]	-1.11 [3]	-1.15 [3]	2.160*	1.656*

B. First Differences

Variables	DF-GLS$_u$ t_μ	t_τ	MZ_a^{GLS}	MZ_t^{GLS}	KPSS η_μ	η_τ
GENERAL INDEX	-16.75* [3]	-16.63* [3]	-424.52* [3]	-14.56* [3]	0.221	0.136
VOLUME	-19.57* [2]	-18.90* [2]	-20.13* [4]	-9.77* [4]	0.198	0.078
BANK	-26.44* [5]	-29.88* [5]	-11.33* [5]	-8.98* [5]	0.233	0.067
BUILD	-33.92* [0]	-34.55* [0]	-49.22* [2]	-15.99* [2]	0.335	0.105
FIN	-44.23* [2]	-40.19* [2]	-31.12* [3]	-9.01* [3]	0.178	0.077
FINSERV	-15.66* [7]	-12.29* [7]	-25.11* [8]	-7.44* [8]	0.201	0.035
FISH	-13.22* [2]	-14.25* [2]	-18.92* [4]	-10.03* [4]	0.099	0.111
HOTEL COMP	-20.11* [3]	-19.18* [3]	-33.90* [5]	-12.03* [5]	0.103	0.099
HOTELS	-30.11* [4]	-25.12* [4]	-18.24* [2]	-9.01* [2]	0.201	0.105
INFORM	-12.55* [6]	-14.79* [6]	-18.03* [5]	-6.22* [5]	0.198	0.034
INSUR	-17.33* [3]	-19.25* [3]	-21.05* [4]	-10.11* [4]	0.095	0.077
MANUF	-52.17* [1]	-45.33* [1]	-33.05* [1]	-8.88* [1]	0.122	0.056
MISC	-38.11* [2]	-40.12* [2]	-19.77* [3]	-7.56* [3]	0.088	0.044
REAL ESTATE	-23.11* [5]	-26.12* [5]	-14.93* [11]	-2.70* [11]	0.244	0.067

Notes:

- 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_a and MZ_t 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 theses 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(l) 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 2. Desciptive statistics – Daily returns

	mean (x 10^2)	variance	m_3	m_4	JB	Q(8)	Q^2(8)
General Index	0.001	0.003	1.88 * [0.00]	25.91 * [0.00]	45120.0 * [0.00]	89.25* [0.00]	103.85 * [0.00]
Transactions Volume	0.20	0.068	-0.38 * [0.00]	28.37 * [0.00]	52823.0 * [0.00]	176.5 * [0.00]	303.8 * [0.00]
Banks	0.011	0.0004	3.03 * [0.00]	47.94 * [0.00]	153685.0 * [0.00]	49.92 * [0.00]	149.0 * [0.00]
Building Materials	-0.17	0.0004	-1.81 * [0.00]	2.92 * [0.00]	146.83 * [0.00]	24.17 * [0.00]	77.60 * [0.00]
Fisheries	-0.40	0.0007	0.17 * [0.15]	3.05 * [0.00]	158.9 * [0.00]	5.14 * [0.74]	49.98 * [0.00]
Investment	-0.007	0.0005	0.91 * [0.00]	10.75 * [0.00]	7836.0 * [0.00]	230.8 * [0.00]	439.47 * [0.00]
Financial Services	-0.4	0.0007	-0.14 [0.25]	5.09 * [0.00]	438.4 * [0.00]	34.30 * [0.00]	4.45 [0.81]
Insurance	-0.069	0.0006	0.02 [0.70]	8.20 * [0.00]	4429.9 * [0.00]	80.40* [0.00]	192.05 * [0.00]
Manufacturing	-0.03	0.0005	1.13 * [0.00]	11.62 * [0.00]	9227.3 * [0.00]	77.81* [0.00]	373.7 * [0.00]
Other	0.000	0.004	-0.13 * [0.03]	80.35 * [0.00]	425071.0 * [0.00]	174.4 * [0.00]	328.08 * [0.00]
Tourist Services	-0.02	0.0006	-0.08 [0.17]	9.13 * [0.00]	5491.1 * [0.00]	159.0 * [0.00]	339.8 * [0.00]
Hotels	-0.12	0.0008	0.19 [0.12]	16.13 * [0.00]	4385.0 * [0.00]	17.50 * [0.02]	37.99 * [0.00]
Realestate	-0.09	0.0008	0.06 [0.35]	23.15 * [0.00]	35280.0 * [0.00]	43.87 * [0.00]	278.1 * [0.00]
Informatics	-0.7	0.0014	0.32 * [0.01]	14.48 * [0.00]	3540.0 * [0.00]	8.22 [0.41]	56.52 * [0.00]

Notes: The average return is expressed in terms of $x10^2$; m_3 and m_4 are the coefficients of skewness and kurtosis of the standardized residuals respectively; JB is the statistic for the null of normality; $Q(8)$ and $Q^2(8)$ are the Ljung-Box test statistics for up to 8th-order serial correlation in the Δp_t and Δp_t^2 series, respectively. (*) denotes statistical significance at the 5 percent critical level.

Table 3. Johansen-Juselius cointegration analysis-(general index-volume)

			5% Critical Values	
R	Trace	λmax	Trace	λmax
r=0	5.75	5.08	18.11	15.02
r=1	0.66	0.66	8.19	8.19

Notes: r denotes the number of eigenvectors. Trace and λmax denote, respectively, the trace and maximum eigevalue likelihood ratio statistics. The 5% critical values are taken from MacKinnon *et al*. (1999; Table III). A structure of nine 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 constant in the VAR equation was estimated following the Johansen (1992a,b; 1994) testing strategy.
(*) denotes statistical significance at the five percent critical level.

Table 4. Johansen-Juselius cointegration analysis-(All sectoral indices)

			5% Critical Values	
R	Trace	λmax	Trace	λmax
r=0	552.9938*	103.7403*	336.22	76.61
r=1	449.2535*	85.80932*	286.39	70.59
r=2	363.4442*	81.59040*	240.58	64.56
r=3	281.8538*	60.80931*	198.72	58.51
r=4	221.0444*	59.72142*	160.87	52.41
r=5	161.3230*	42.03399	127.05	46.31
r=6	119.2890*	35.69364	97.26	40.19
r=7	83.59539*	30.56815	71.44	34.03
r=8	53.02724*	22.93630	49.64	27.80
r=9	30.09094	20.29929	31.88	21.49
r=10	9.791651	9.733150	18.11	15.02
r=11	0.058501	0.058501	8.19	8.19

Notes: r denotes the number of eigenvectors. Trace and λmax denote, respectively, the trace and maximum eigevalue likelihood ratio statistics. The 5% critical values are taken from MacKinnon *et al*. (1999; Table III). A structure of ten 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 constant in the VAR equation was estimated following the Johansen (1992a,b; 1994) testing strategy.
(*) denotes statistical significance at the five percent critical level.

Table 5. Johansen-Juselius bivariate cointegration analysis

	BANK	BUILD	FIN	FINSERV	FISH	HOTELCO MP	HOTELS	INFORM	INSUR	MANUF	MISC	REALE STATE
BANK	-											
BUILD	YES	-										
FIN	NO	NO	-									
FINSERV	YES	NO	NO	-								
FISH	YES	NO	NO	NO	-							
HOTELCOMP	YES	NO	NO	NO	NO	-						
HOTELS	YES	NO	NO	YES	NO	NO	-					
INFORM	YES	NO	NO	NO	NO	NO	NO	-				
INSUR	NO	NO	NO	NO	NO	NO	NO	NO	-			
MANUF	NO	NO	NO	NO	NO	YES	NO	NO	YES	-		
MISC	YES	NO	NO	NO	NO	NO	YES	NO	NO	NO	-	
REALESTATE	NO	NO	NO	NO	YES	NO	NO	YES	YES	NO	NO	-

Notes: Yes=the null hypothesis of nocointegration is rejected; No=the null hypothesis of nocointegration could not be rejected. The level of significance is five percent.

Table 6. Linear Granger causality

	BANK	BUILD	FIN	FINSERV	FISH	HOTELCOMP	HOTELS	INFORM	INSUR	MANUF	MISC	REALESTATE
BANK	-											
BUILD	NO	-										
	YES											
FIN	NO	YES	-									
	NO	YES										
FINSERV	YES	YES	NO	-								
	YES	NO	NO									
FISH	NO	YES	NO	NO	-							
	YES	YES	YES	YES								
HOTELCOMP	NO	YES	NO	NO	YES	-						
	NO	NO	NO	YES	NO							
HOTELS	YES	YES	NO	YES	YES	YES	-					
	YES	NO	NO	YES	NO	NO						
INFORM	NO	NO	NO	NO	NO	NO	YES	-				
	YES	NO	YES	NO	NO	NO	NO					
INSUR	YES	YES	NO	YES	YES	NO	NO	YES	-			
	YES	NO	NO	YES	NO	NO	NO	NO				
MANUF	YES	YES	NO	YES	YES	NO	NO	NO	YES	-		
	NO	NO	NO	YES	NO	YES	NO	NO	YES			
MISC	NO	YES	NO	NO	YES	YES	NO	NO	YES	YES	-	
	YES	YES	NO	NO	NO	NO	NO	NO	YES	YES		
REALESTATE	YES	YES	NO	NO	YES	NO	NO	NO	YES	YES	YES	-
	YES	NO	NO	YES	NO	NO	YES	YES	YES	YES	YES	

Notes: NO = F is not statistically significant at the 5% level of significance. YES = F is statistically significant at the 5% level of significance.

Table 7. Nonlinear Granger causality

	BANK	BUILD	FIN	FINSERV	FISH	HOTELCOMP	HOTELS	INFORM	INSUR	MANUF	MISC	REALESTATE
BANK	-											
BUILD	YES NO	-										
FIN	NO NO	YES NO	-									
FINSERV	NO NO	NO NO	NO NO	-								
FISH	NO NO	NO NO	NO NO	NO NO	-							
HOTELCOMP	NO YES	NO NO	NO NO	NO YES	NO NO	-						
HOTELS	NO NO	YES NO	NO YES	NO NO	NO NO	YES YES	-					
INFORM	YES NO	NO NO	NO NO	NO NO	NO NO	NO NO	NO NO	-				
INSUR	YES YES	YES YES	YES YES	NO NO	NO NO	NO YES	NO NO	NO NO	-			
MANUF	NO NO	YES NO	YES NO	YES NO	NO NO	NO YES	NO NO	NO NO	NO NO	-		
MISC	NO YES	YES YES	NO NO	NO NO	NO NO	NO NO	NO NO	NO NO	YES YES	NO NO	-	
REALESTATE	YES NO	YES NO	YES NO	NO NO	NO NO	NO YES	YES NO	NO NO	NO NO	NO NO	YES NO	-

Notes: NO = F is not statistically significant at the 5% level of significance. YES = F is statistically significant at the 5% level of significance.

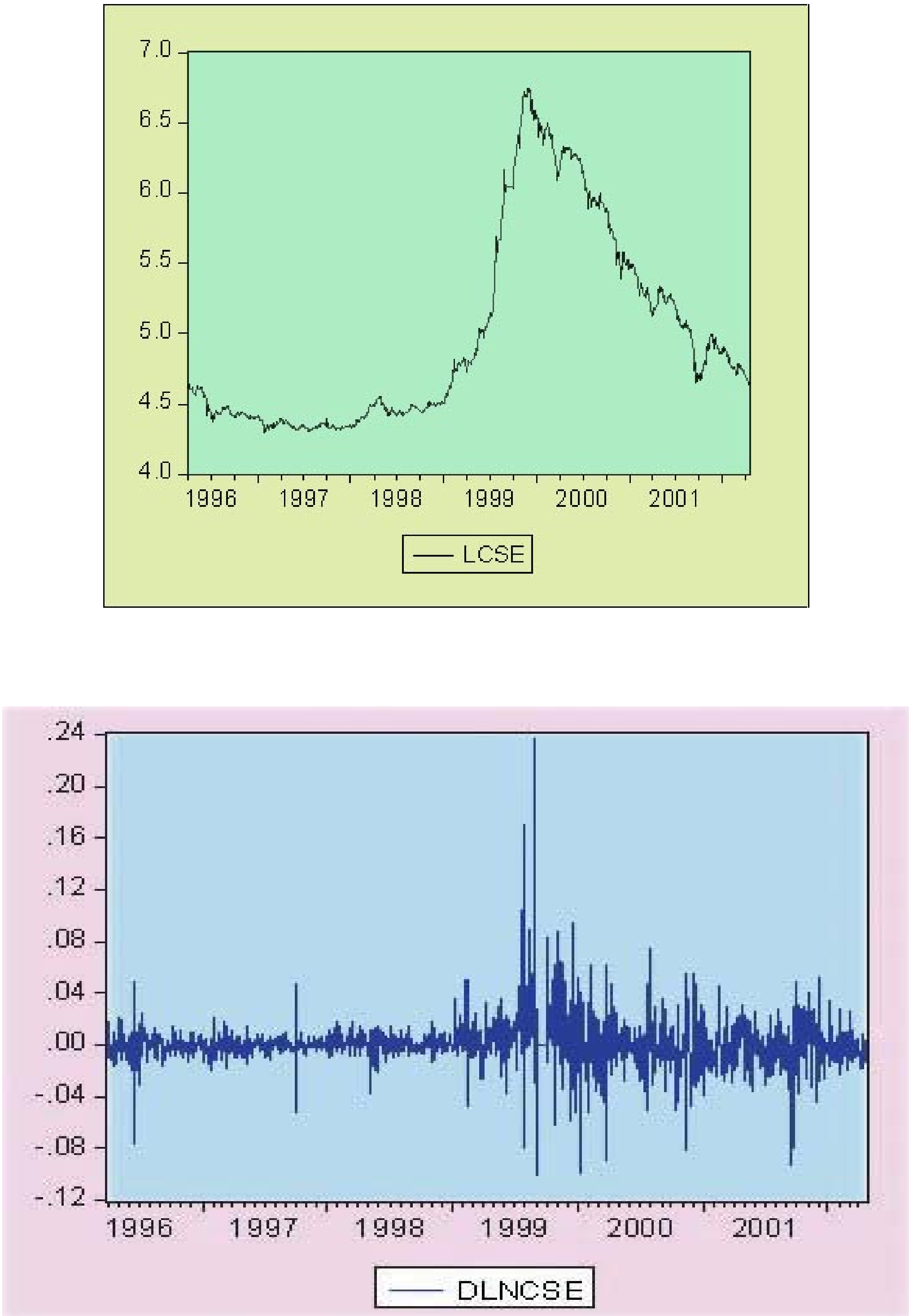

Figure 1. CSE General Price Index and Returns