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Testing for Seasonal Integration and Cointegration: An Expository Note with Empirical Application to KLSE Stock Price Data

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ABSTRAK

Tujuan kertas kerja ini adalah untuk menyiasat sifat-sifat bermusim siri harga saham mengikut sektor di Bursa Saham Kuala Lumpur (KLSE) untuk jangkamasa 1978:1 hingga 1992:3. Keputusan kajian mencadangkan bahawa indeks-indeks harga saham di KLSE mempamerkan punca satu bermusim, bukan sahaja pada frekuensi sifar tetapi dalam kebanyakkan kes pada frekuensi dua kali setahun. Hasil kajian yang mencadangkan harga saham mempamerkan integrasi bermusim memberi implikasi penting terhadap kointegrasi bermusim. Walau bagaimana pun keputusan-keputusan ujian kointegrasi bermusim mencadangkan bahawa indeks-indeks harga saham mengikut sektor di KLSE tidak berkointegrasi bermusim. Keputusan ini menyarankan bahawa hipotesis kecekapan bermaklumat pasaran saham tidak boleh di tolak untuk KLSE.

ABSTRACT

The purpose of this paper is to investigate the seasonal properties of the sectoral stock price series at the Kuala Lumpur Stock Exchange (KLSE) for the period 1978:1 to 1992:3. Our results suggest that the stock price indices at the KLSE exhibit seasonal unit roots, not only at the zero frequency, but in most cases at the biannual frequency. The finding that stock price indices exhibit seasonal integration has important implications for seasonal cointegration. However, our seasonal cointegration test results suggest that sectoral stock price indices at the KLSE are not seasonally cointegrated. These results imply that the informationally efficient stock market hypothesis cannot be rejected for the KLSE.

INTRODUCTION

The concept of cointegration first introduced by Granger (1981) relates to the notion of a long run or equilibrium relationship among two or more variables. Granger points out that the series may be unequal in the short run but they are tied together in the long run, that is, they move parallel to each other over time. According to Granger (1986) and Engle and Granger (1987), a very important consequence of cointegrated variables is that one variable can be used to predict the other. Granger (1986) notes that, "if x_t, y_t are I(1) and co-integrated, there must be Granger causality in at least one direction as one variable can help forecast the other."

The method of cointegration is a very useful tool in economics, particularly in searching for long-run relationships between various economic variables. Some real world examples were given by Granger (1986). He states that, "such variables are interest rates on assets of different maturities, prices of a commodity in different parts of the country, income and expenditure by local government and the value of sales and production costs of an industry. Other possible examples would be prices and wages, imports and exports, market prices of substitute commodities, money supply and prices and spot and future prices of a commodity."

However, before we test for cointegration among variables, we need to know the stationarity status of the series. In empirical work, we only deal with stationary series. A non-stationary variable contains a unit root. It has neither fixed mean nor a constant variance. A non-stationary variable such as a random walk will not fluctuate about a certain mean. On the other hand, a stationary variable has a fixed mean and a constant variance. When plotted over time, a stationary variable is characterised by numerous fluctuations about the mean. For example, white noise is a stationary process, and such does not contain a unit root.

In cointegration analysis, it is important that the series under study have the same order of integration. Series X_t and Y_t are integrated of the same order, denoted by $X_t \sim I(d)$ and $Y_t \sim I(d)$, if the two time series required to be differenced *d* times to achieve stationarity. A series $X_t \sim I(1)$, that is integrated of order one, needs to be differenced only once to achieve stationarity, that is, to become I(0). According to Granger (1986), 'an I(0) series has a mean and there is a tendency for the series to return to the mean, so that it tends to fluctuate around the mean, crossing the value frequently and with rare extensive excursions.'

Inspired by the seminal work of Nelson and Plosser (1982), there is a vast literature that investigates on whether macroeconomic time series contain a unit root. Among others, studies by Schwert (1987), Wasserfallen (1986) and Vujosevic (1992) conclude that many macroeconomic time series contain a unit root. The finding that many economic time series have a unit root (i.e. are nonstationary in their levels) led to the concept of cointegration suggested by Granger (1981). Furthermore, the existence of cointegration among these nonstationary economic time series provides a statistical foundation for the use of error-correction models. The notion of integration, cointegration and error-correction modelling has been extensively tested and investigated in recent years.

More recently, attention has been directed to the testing of integration and cointegration for the presence of seasonality in economic time series. Previous studies that used high frequency (monthly and quarterly) data have either ignored the seasonal components or used seasonal adjusted data in their analysis. Osborn *et al.* (1988) and Hylleberg *et al.* (1990) have pointed out that high frequency economic time series might also have seasonal unit roots besides the unit root at zero frequency. Despite this warning, most researchers avoided the issue of seasonality in economic time series. Kunst (1994) gives three reasons for disregarding the role of seasonality in economic study. Firstly, researchers assumed that the dummy-style determination system can appropriately eliminate seasonal variations in economic time series. Secondly, researchers regard seasonal phenomena as a nuisance and as such seasonal adjustment procedures are used to eliminate them. Thirdly, the usefulness of findings for seasonal integration and cointegration for empirical tests and application has yet to be established.

Nevertheless, the important role of seasonality in economic time series has been given serious attention in recent years. Among the most recent studies include Osborn (1990), Engle *et al.* (1993), McDougall (1994, 1995), Linden (1994), Hurn (1993), Hylleberg *et al.* (1993) and Sarantis and Stewart (1993). The finding of the studies on seasonality in macroeconomic time series by Osborn (1990) for the United Kingdom, Otto and Wirjanto (1990) for Canada, Ghysels *et al.* (1994) for the United States, McDougall (1995) for New Zealand and Hylleberg *et al.* (1993) for several developed countries suggest that many macroeconomic time series exhibit significant seasonality. Osborn (1990) concludes that the finding for seasonal integration in those economic time series have important implications for seasonal cointegration.

McDougall (1994) and Hurn (1993) investigate the long-run relationship between money and income for New Zealand and South Africa respectively. Although McDougall (1994) finds a nonseasonal relationship between money and income for the New Zealand economy, Hurn finds support for the existence of seasonal cointegration in the South African monetary data. Hurn notes that the inclusion of seasonal components improves the overall performance of the final error-correction models. On the other hand, studies by Linden (1994) on labour demand in Finnish manufacturing, and Engle *et al.* (1993) on the Japanese consumption function, provide evidence in favour of seasonal cointegration. Nevertheless, Sarantis and Stewart (1993) found that exchange rates and relative prices of several developed countries under study do not support the existence of seasonal cointegration.

In a more recent study, Moosa (1995) clearly indicates the importance of the order of (seasonal) integration of time series variables. Moosa finds out that previous studies on Australian consumption function are mis-specified and as a result performed poorly as expected. Moosa (1995) concludes that, "the failure of these equations is due to the use of inappropriate filters (by overlooking the time series properties of the variables) and faulty error correction terms (by only allowing for cointegration at the zero frequency or the long run and not at other frequencies). A finding of this study is that Australian non-durable consumption and disposable income are cointegrated at the frequency 1/4 (annual cycle), implying that the consumption-income relationship should be modelled as a seasonal error correction model."

Despite the increasing interest in testing for seasonality in macroeconomic series among researchers, most of the existing studies, except for Hurn (1993), are mainly confined to the developed nation. Therefore, there is an imperative need to conduct a similar study to investigate the seasonal behaviour of macroeconomic time series of the developing economies. Thus, the primary aim of the paper is to complement the existing literature of testing for seasonality in macroeconomic time series of the developing countries.

METHODOLOGY

The importance of seasonality in economic time series has been recognised and has been given proper treatment in economic literature. The work of Box and Jenkins (1970) implicitly assumes that there are seasonal unit roots in the series by using the seasonal differencing filter. Other researchers prefer using seasonally adjusted data in the analysis. However, these approaches have been criticised by Miron (1992) and Ghysels (1988, 1990, 1992). They pointed out that seasonal adjustment might lead to wrong inference about economic relationships between the series under study. The seasonal adjustment biases the outcome toward accepting the null hypothesis that a unit root exists. Olekalns (1994) concludes that, 'tests of the unit root hypothesis should not be carried out with seasonally adjusted data.'

When comparing the performance of a series between seasonally adjusted and seasonally unadjusted data, Ghysels (1990, 1992) found that the nature of unit root between the seasonally adjusted and seasonally unadjusted series gave contradictory results. Ghysels concludes that the seasonal adjustment procedure might alter the outcome of the conventional test and therefore gave substantially different results. On the other hand, Miron (1992) points out that seasonal fluctuations are not a nuisance, instead seasonality has economic importance in economic analysis and acts as a source of information in understanding economic relationships.

Thus, the problems associated with seasonal adjustment have led to the examination of seasonal unit roots and hence tests to determine orders of seasonal integration for economic time series. The essence of seasonality is that not only must each of the series be integrated of the same order but they must be seasonally integrated of the same order, otherwise the estimates of the cointegrating equations will be inconsistent. In other words, the estimates result in a spurious regression problem (Hylleberg *et al.* 1990).

Testing for Seasonal Unit Roots

For a seasonally unadjusted economic time series, the concept of integration will include the possibility of seasonal unit roots. A seasonal economic time series, X_t, is said to be integrated of order (d, D), that is $X_t \sim I(d, D)$ if the series is stationary after first period differencing d times (unit root) and seasonal differencing D times (seasonal unit root) (see Osborn *et al.*, 1988).

According to Hylleberg *et al.* (1990) and Engle *et al.* (1993), for quarterly data, the seasonal difference operator $(1-B^4)$ can be decomposed into four possible roots in the generating process as follows:

$$(1-B^4) = (1-B)(1+B)(1-iB)(1+iB) \qquad (1)$$

In equation (1), the unit roots are 1, -1, i and –i which correspond to zero frequency, one-half (1/2) cycle per quarter or two cycles per year in quarterly data and one fourth (1/4) period cycle corresponding to one-quarter cycle per quarter or one cycle per year in quarterly data. However, the last root , -i, is indistinguishable from the one at i with quarterly data and therefore it is treated as the annual cycle.

The testing procedure for seasonal unit root has been provided by Hasza and Fuller (1982), Dickey *et al.* (1984), Osborn *et al.* (1988), Osborn (1990), Hylleberg *et al.* (1990) and Engle *et al.* (1993). The latter two seasonal unit root testing procedures are the most popular among researchers. Furthermore, Ghysels *et al.* (1994) found out that Hylleberg *et al.* (1990) (thereafter HEGY) procedure compares favourably with other alternative procedures, in particular, with Dickey *et al.* (1984) tests.

The HEGY (1990) approach consists in estimating the following regression:

$$\Delta_4 x_t = \pi_1 y_{1t-1} + \pi_2 y_{2t-1} + \pi_3 y_{3t-2} + \pi_4 y_{3t-1} + \sum_{i=1}^{p} \pi_4 x_{t-i} + \upsilon_t \quad (2)$$

where $y_{1t} = (1+B+B^2+B^3)x_t$, $y_{2t} = -(1-B+B^2-B^3)xt$ and $y^{3t} = -(1-B^2)x_t$. For quarterly time series data, deterministic components are added in equation (2). The test regression now becomes

$$\Delta_4 x_t = \alpha_0 + \alpha_1 SD_{1t} + \alpha_2 SD_{2t} + \alpha_3 SD_{3t} + \theta t + \pi_1 y_{1t-1} + \pi_2 y_{2t-1} + \pi_3 y_{3t-2} + \pi_4 y_{3t-1} + \sum_{i=1}^{p} \Delta_4 x_{t-i} + \varepsilon_t \quad (3)$$

where α_0 is a constant, t is a linear time trend and SD_{it}'s are quarterly seasonal dummy variables. The test for seasonal unit roots is by running ordinary least square (OLS) on equation (3) and the test statistics on π's can be used for inferences. According to a simulation study by Ghysels *et al.* (1994), the inclusion of a constant and seasonal dummies appears to be a prudent decision in testing for seasonal unit roots. Ghysels *et al.* (1994) further conclude that, "it was found that when the data-generating processes have seasonal dummies, the regression without seasonal dummies seriously distorts the test result [i.e. it leads to a large bias in the size or too low power]. Hence, although inclusion of too many lags or irrelevant deterministic terms (i.e. a constant, seasonal dummies, and/or a trend) tends to reduce the power of the tests, the safe strategy in empirical applications is the inclusion of these (possibly irrelevant) terms in the model."

To test for a unit root at zero frequency (i.e. $x_t \sim I_0(1)$ we simply perform a t-test on $\pi_1 = 0$. To test for root -1 (the biannual frequency unit root) that is, $x_t \sim I_{1/2}(1)$, a test on $\pi_2=0$ is performed. For the complex roots (an annual frequency unit root) or $x_t \sim I_{1/4}(1)$, we can perform either a joint *F*-test of $\pi_3=\pi_4=0$, or two sequential t-tests of $\pi_4=0$ and then $\pi_3=0$. For a series to contain no seasonal unit roots, $\pi_2=0$ and the joint *F*-test of $\pi_3=\pi_4=0$ must both be rejected. On the other hand, for a series to be stationary, it must have no unit roots, hence, it must be established that each of the t-test of $\pi_1=\pi_2=0$ and the joint *F*-test of $\pi_3=\pi_4=0$ are rejected. The critical values can be found in Hylleberg *et al.* (1990).

In equation (3), the choice of the truncation lag parameter, p, can be determined according to a variety of lag selection criteria. Engle *et al.* (1993) pointed out that the power and size of the unit root tests depend critically on the 'right' augmentation being used. Too many parameters will decrease the power of the tests while too few will render the size far greater than the level of significance. Engle *et al.* (1993) rely on the augmentation approach by estimating equation (3) for some lag length. After establishing which of the lags are statistically significant, the equation is then re-estimated by including only the statistically significant autoregressive terms. The net result is to leave gaps or 'holes' in the lag distribution of the autoregressive terms in equation (3). Ghysels *et al.* (1994) used Hall's 'data-based model-selection' procedure consisting in estimating the number of autoregressive terms according to the longest lag with a statistically significant coefficient, beginning with a maximum lag length of 7 quarters. A general-to-specific approach used by McDougall (1994), which was based on Perron (1989), consists in starting with a given number of arbitrary maximum lagged regressors, say k, and then successively reduced until the last included lag has significant coefficient based on the usual t-test. On one hand, Lee and Siklos (1991) used the well known Akaike and Schwarz criteria, and on the other, Otto and Wirjanto (1990) and Osborn (1990) based their analysis on the significance of the Lagrange Multiplier test for serial correlation to choose the 'best' model for each series. Yet others have employed the Akaike's (1969) Final Prediction Error (FPE) criterion in selecting the optimal lag length. Hsiao (1981) points out that the FPE criterion is equivalent to using an *F*-test but with a varying level of significance. As Judge *et al.* (1982) argued, the intuition behind this procedure is that as the lag length on the variable under consideration increases, the first term of FPE increases while the second term decreases and as a result, these opposing forces are balanced when their product reaches a minimum. Furthermore, according to Hsiao (1979), 'the criterion tries to balance the risk resulting from the bias when a lower order is selected and the risk resulting from the increase of variance when a higher order is selected by choosing the specification that gives the smallest FPE.'

Seasonal Cointegration and Error Correction Model

According to HEGY (1990), 'a pair of series each of which is integrated at frequency ω are said to be cointegrated at that frequency if a linear combination of the series is not integrated at ω.' For a two-variable case consisting of X and Y, where y_t and $x_t \sim I_\omega(1)$, $\omega = 0, 1/4, 1/2, 3/4$, there may exist one or no cointegrating vector at each frequency. The general form of the error-correcting mechanism which allows for cointegration (at one cointegrating vector) at all frequencies, $\omega = 0, 1/4, 1/2, 3/4$, is shown to be

$$\Delta_4 y_t = \sum_{i=1}^{q} \phi_i \Delta_4 y_{t-i} + \sum_{j=1}^{p} \lambda_j \Delta_4 x_{t-j} + \gamma_1 U_{t-1} + \gamma_2 v_{t-1} + \gamma_3 w_{t-2} + \gamma_4 w_{t-3} + \eta_t \quad (4)$$

where u^{t-1}, v_{t-1}, w_{t-2} and w_{t-3} are lagged residuals of the following respective cointegrating equations (5), (6) and (7), derived by Engle *et al.* (1993),

$$u_{1t} = y_{1t} - \alpha_1 x_{1t} \quad (5)$$

$$v_{1t} = y_{2t} - \alpha_2 x_{2t} \quad (6)$$

$$w_{1t} = y_{3t} - \alpha_3 x_{3t} - \alpha_4 x_{3t-1} \quad (7)$$

where in each case the x^{it} and y_{it} (i=1,2,3) represent the zero, biannual and annual frequencies which are run with or without deterministic components including an intercept (I), seasonal dummies (SD's) and a time trend (T).

The residuals u, v and w are tested for their stationarity characteristics according to the following manner outlined by Engle *et al.* (1993). The test for noncointegration at the zero frequency can be performed by establishing the following equation

$$\Delta u_t = \pi_1 u_{t-1} + \sum_{i=1}^{k} \delta_i \Delta u_{t-1} + \text{deterministic components} + \tau_{1t} \quad (8)$$

When testing for noncointegration at the biannual frequency (i.e. 1/2), we run the following auxiliary regression

$$v_t + v_{t-1} = \pi_2(-v_{t-1}) + \sum_{i=1}^{k} \delta_i (v_{t-i} + v_{t-i-1}) + \text{deterministic components} + \tau_{2t} \quad (9)$$

Similarly to the above tests, the test for seasonal noncointegration at the annual frequency (i.e. 1/4 and 3/4) can be performed by estimating the following equation:

$$w_t + w_{t-2} = \pi_3(-w_{t-2}) + \pi_4(-w_{t-1}) + \sum_{i=1}^{k} \delta_i(-w_{t-1} + -w_{t-2}) + \text{deterministic components} + \tau_{3t} \quad (10)$$

The t-values of the test statistics of π's can be used for inference for noncointegration at zero, biannual and annual frequencies. However, for testing for noncointegration at the annual frequency, the F-value of the joint test $\pi_3 = \pi_4 = 0$ is computed together with the t-values for $\pi_3 = 0$ and $\pi_4 = 0$. The critical values for π_1 and π_2 are tabulated in Engle and Yoo (1987). On the other hand, the critical value for F-statistic for $\pi_3 \cap \pi_4 = 0$ are tabulated in Engle *et al.* (1993).

Description and Sources of Data Used

In this paper, the testing for seasonal integration and cointegration is applied to sectoral stock prices at the KLSE for the period 1978:1 to 1992:3. The stock price indices are the Composite, Industrial, Finance, Property, Agriculture and Tin. The stock price indices were collected from various issues of the Investors Digest published monthly by KLSE. All data used in the analysis are transformed into natural logarithms before estimation.

DISCUSSION ON EMPIRICAL RESULTS

Results of Seasonal Unit Root Tests

The results of applying the HEGY test for seasonal unit roots are presented in Table 1. Based on equation (3), each of the variables in the logarithmic form is then regressed against (i) without the deterministic components, (ii) a constant, (iii) a constant and seasonal dummies, (iv) a constant and trend, and (v) with a constant, seasonal dummies and trend. In Table 1, we report the final specification of equation (3) which was based on the chosen optimal lag length. The optimal lag length, p, was determined using the Perron's (1989) liberal approach which consists in starting with a given number of lagged dependent variables and paring down the model by the usual t-statistics. If the t-statistics on the last lagged term is less than 1.6, the term is dropped from the model. The process is repeated until the t-statistics on the last lagged coefficient is greater

TABLE 1
HEGY tests for seasonal unit roots

Variables	Deterministic components	π_1	π_2	π_3	π_4	$\pi_3 \cap \pi_4$	Lag	LM(4)	Significance of deterministic components *I*	*SD*	*Tr*	Frequencies with a unit root
Composite	–	0.91	-3.43**	-3.02**	-3.59**	11.13**	1	4.12	-	-	-	0
	I	-1.96	-3.34**	-3.21**	-3.24**	10.59**	1	1.11	**	-	-	0
	I, SD	-1.84	-3.51**	-3.57	-3.67**	13.27**	1	2.66	ns	ns	-	0
	I, Tr	-1.92	-2.16**	-2.52**	-0.53	3.39**	8	7.54	ns	-	ns	0
	I, SD, Tr	-2.20	-2.60	-3.77**	-0.81	7.75**	12	7.09	ns	**	**	0, 1/2
Industrial	–	1.00	-3.45**	-3.49**	-3.01**	10.75**	1	5.58	-	-	-	0
	I	-1.47	-3.36**	-3.60**	-2.80**	10.58**	1	3.36	ns	-	-	0
	I, SD	-1.37	-3.59**	-3.91**	-3.23**	13.03**	1	5.76	ns	ns	-	0
	I, Tr	-2.15	-1.89	-2.41**	-0.62	3.18**	8	4.96	**	-	**	0, 1/2
	I, SD, Tr	-2.34	-2.54	-3.32	-0.79	6.09	12	4.94	**	ns	**	0, 1/2, 1/4
Finance	–	1.13	-3.47**	-3.35**	-2.96**	10.12**	1	5.40	-	-	-	0
	I	-2.62	-2.93**	-3.13**	-0.76	10.00**	3	1.02	**	-	-	0
	I, SD	0.43	-2.28	-3.81**	-0.61	7.64**	12	6.01	ns	**	-	0, 1/2
	I, Tr	-2.85	-2.25**	-2.84**	-0.14	4.08**	8	7.63	**	-	**	0
	I, SD, Tr	-2.51	-2.43	-4.33**	-0.42	9.62**	12	6.60	**	**	**	0, 1/2
Property	–	0.61	-2.51**	-3.37**	-2.11**	9.10**	2	5.98	-	-	-	0
	I	-3.25**	-2.14**	-3.08**	0.15	4.76**	5	2.48	**	-	-	stationary
	I, SD	-2.50	-2.24	-3.87**	0.36	7.51**	8	2.47	**	**	-	0, 1/2
	I, Tr	-3.29	-2.07**	-3.10**	0.25	4.83**	5	3.93	**	-	ns	0
	I, SD, Tr	-2.95	-2.30	-3.97**	0.39	7.90**	8	5.77	**	**	ns	0, 1/2
Agriculture	–	0.53	-3.09**	-3.35**	-2.86**	9.74**	1	5.42	-	-	-	0
	I	-3.05**	-3.08**	-3.70**	-2.32**	9.60**	1	3.49	**	-	-	stationary
	I, SD	-1.83	-3.65**	-2.60	-0.87	3.70	9	2.23	ns	**	-	0, 1/4
	I, Tr	-3.32	-2.89**	-3.77**	-1.90**	8.98**	1	4.54	**	-	ns	0
	I, SD, Tr	-3.25	-3.66**	-3.66**	-0.90	7.46**	4	4.08	**	**	**	0
Tin	–	-0.33	-4.05**	-2.27**	-1.72	4.08**	3	3.81	-	-	-	0
	I	-2.40	-2.34**	-2.63**	-0.11	3.49**	8	5.51	**	-	-	0
	I, SD	-2.60	-2.31	-4.20**	-0.13	8.92**	12	5.70	**	**	-	0, 1/2
	I, Tr	-2.37	-2.34**	-2.64**	-0.17	3.54**	8	5.33	**	-	ns	0
	I, SD, Tr	-2.48	-2.29	-4.18**	-0.16	8.86**	12	7.25	**	**	ns	0, 1/2

Notes; The LM Chi-Square statistics for serial correlation with 4 lags is 9.48 with 4 degree of freedom (5%). ns denotes not significant. Asterisk, **, denotes statistically significant at five percent level. Critical values for 48 observations and at 5 percent significance level are as follows (see Hylleberg *et al.* 1990):

Deterministic components	π_1	π_2	π_3	π_4	$\pi_3 \cap \pi_4$
-	-1.95	-1.95	-1.93	-176/1.72	3.26
I	-2.96	-1.95	-1.90	-1.72/1.68	3.04
I, SD	-3.08	-3.04	-3.61	-1.98/1.96	6.60
I, Tr	-3.56	-1.91	-1.92	-1.70/1.64	2.95
I, SD, Tr	-3.71	-3.08	.-3.66	-1.91/1.97	6.55

than 1.6. A series of autoregressions is estimated for equation (3) by varying the lag order p from 1 to 12. For each lag length chosen, the presence of serial correlation in the residuals is checked using the Breusch-Godfrey LM test for fourth-order autoregression[1]. All final equations estimated show that the null hypothesis of no serial correlation can be rejected at 5 percent significance level.

Several observations can be derived from the results presented in Table 1. First, most of the final specifications indicate a shorter lag length. The evidence is stronger for Agriculture and followed by Composite and Industrial sectors. Secondly, in majority of the cases, the HEGY tests are not robust to the inclusion of deterministic components in the equation. For Composite stock price, except one, the results overwhelmingly indicate that seasonal unit roots cannot be rejected at the zero frequency. When a constant, seasonal dummies and trend were included in the equation, the results suggested that seasonal unit roots could not be rejected at frequency zero and 1/2 for the Composite stock price. In the cases of Finance and Tin sectors, the results suggested that seasonal unit roots could not be rejected at frequency zero and biannual when a constant and seasonal dummies or all three deterministic components were included in equation (3). Similar results were also obtained for the Property sector except one, in that, an equation with a constant, the HEGY test suggest that Property stock price is stationary in levels. The HEGY test also suggested that Agriculture stock price is stationary in levels when a constant is included in equation (3). However, when a constant and seasonal dummies were included in the equation, the results suggested that seasonal unit root could not be rejected at frequency zero and 1/4. Last but not least, in case of Industrial stock price, results suggest that seasonal unit roots cannot be rejected at all frequencies (i.e. 0, 1/2 and 1/4) when a constant, seasonal dummies and trend are included; at frequencies zero and 1/2 when a constant and trend are included; and at frequency zero when either a constant or a constant and seasonal dummies or no deterministic components are included in the model.

Since the HEGY test is sensitive to the inclusion of deterministic components, our question is: How do we select the appropriate HEGY regression equation? In this paper, we do this by infering at the significance of the deterministic components. The appropriate HEGY regression equation selected is the one with the most significance deterministic components. Based on this criterion, the appropriate HEGY regression equation for Composite stock price is the one which includes a constant, seasonal dummies and a trend; Industrial with a constant and a trend; Finance with all three deterministic components; Property with a constant and seasonal dummies; Agriculture with a constant, seasonal dummies and trend; and Tin with a constant and seasonal dummies. Based on these results, we conclude that seasonal unit roots at zero and biannual frequencies are suggested for all stock price indices, except for Agriculture stock price where seasonal unit root at zero frequency is suggested.

Results of Seasonal Cointegration Tests

The results of the above unit root tests indicate that stock price series are integrated of order one but at some specific frequencies. Having established that the stock price indices are seasonally integrated, our next attempt is to investigate whether these stock price series are seasonally cointegrated along the lines suggested by Engle *et al.* (1993). In our case, in the quarterly series, cointegration between sectoral stock price integrated at the biannual frequency is said to exist if there is at least one linear combination of the series that is stationary at that frequency. For the annual frequency case, cointegration is said to exist if there is at least one linear combination of the series, all integrated at the annual frequency and the series lagged one quarter which is stationary at that particular frequency.

Following Engle and Granger's (1987) two-step procedure, the test for seasonal noncointegration at a particular frequency is based on a test for a unit root at that frequency

[1] According to Harvey (1985), the LM principle for testing for serial correlation yields more satisfactory test compared to the Box-Ljung test.

in the residuals from a first step regression. The first step regression is a regression of one of the series on the other, but after proper transformations so that no unit roots exist at other frequencies. As mentioned earlier, a test of noncointegration at the long-run frequency is a test for a unit root at the zero frequency in the residuals, μ_t, from a regression y_{1t} on x_{1t} where y_{1t} is $(1+B+B^2+B^3)y_t$, that is the sum of four consecutive values of say, Industrial stock price series while x_{1t} is defined analogously for say, Finance stock price series. Likewise, a test of noncointegration at the biannual frequency is a test of there being a unit root at that frequency in the residuals, v_t, from a regression of $y_{2t} = -(1-B+B^2-B^3)y_t$ on $x_{2t} = -(1-B+B^2-B^3)x_t$. And for the annual frequency the first step regression is $y_{3t} = -(1-B^2)y_t$ on $x_{3t} = -(1-B^2)x_t$ and x_{3t-1} and the test for a unit root at the annual frequency in the residuals w_t is based on the *F*-value for $\pi_3=\pi_4=0$ in the regression $(w_t+w_{t-2}) = \pi_3(-w_{t-2}) + \pi_4(-w_{t-1})$. In testing for cointegration, we allow for augmentation of the lagged dependent variable so as to induce white noise following Perron's (1989) liberal approach mentioned earlier.

Seasonal cointegration is usually conducted between those series which appeared to be seasonally integrated at common frequencies. In our case, we conducted a pairwise seasonal cointegration at the long-run between Industrial, Finance, Property, Agriculture and Tin. On the other hand, seasonal cointegration at the biannual frequency is conducted between Industrial, Finance, Property and Tin. The results of testing for seasonal cointegration at the long-run and biannual frequencies are presented in Tables 2 and 3 respectively. Looking through Table 2, the results suggest that sectoral stock prices at the KLSE are not cointegrated at the zero frequency. The t-statistics for π_1 in all cases are smaller (in absolute term) than the critical value tabulated in Engle and Yoo (1987). On the other hand, results in Table 3 also suggest that cointegration at the biannual frequency can also be rejected between the sectoral stock prices at the KLSE. In all cases the t-statistics for π_2 are smaller (in absolute term) than the critical value tabulated in Engle and Yoo (1987).

CONCLUSION

More recently, the work of Hylleberg *et al.* (1990) and Engle *et al.* (1993) has enabled researchers to investigate the time series properties of an economic series when they contain seasonal components not only at the zero frequency but

TABLE 2
Tests for cointegration at frequency zero: the long-run

Regressand	Regressor	Cointegrating regression			Tests for unit roots in residuals		
		Coefficient on regressor	R^2	D.W.	Augmented Dickey-Fuller test π_1	Lag	LM(4)
Industrial	Finance	0.507	0.829	0.07	-1.86	5	2.21
	Property	0.383	0.894	0.09	-1.55	5	1.79
	Agriculture	0.704	0.817	0.08	-1.31	5	5.67
	Tin	0.769	0.961	0.19	-3.19	10	2.59
Finance	Property	0.548	0.966	0.09	-2.11	5	3.46
	Agriculture	1.226	0.913	0.12	-1.72	9	2.94
	Tin	0.422	0.775	0.05	-2.24	5	3.14
Property	Agriculture	1.981	0.737	0.11	-1.71	9	6.92
	Tin	1.089	0.605	0.05	-0.87	11	3.16
Agriculture	Tin	0.360	0.691	0.09	-1.00	8	4.28

Notes: The t-statistics for π_1 isdistributed as described in Engle and Granger (1987) and Engle and Yoo (1987). The critical value at five percent significance level is 3.29 for T equals 50 observations. The LM Chi-Square statistics for serial correlation with four lags is 9.48 with four degree of freedom (5%). All cointegrating regressions and the auxillary regressions are estimated with a constant and seasonal dummies.

TABLE 3
Tests for cointegration at frequency 1/2: biannual cycle

Regressand	Regressor	Cointegrating regression			Tests for unit roots in residuals		
		Coefficient on regressor	R^2	D.W.	Augmented Dickey-Fuller test π_2	Lag	LM(4)
Industrial	Finance	0.886	0.725	2.67	-1.85	11	5.76
	Property	0.624	0.671	2.84	-2.87	7	4.12
	Tin	0.729	0.648	2.88	-2.44	4	1.85
Finance	Property	0.662	0.813	2.52	-3.06	7	4.94
	Tin	0.696	0.636	1.88	-3.07	3	6.48
Property	Tin	0.938	0.622	1.50	-2.21	11	4.99

Notes: The t-statistics for π_2 is distributed as described in Engle and Granger (1987) and Engle and Yoo (1987). The critical value at five percent significance level is 3.29 for T equals 50 observations. The LM Chi-Square statistics for serial correlation with four lags is 9.48 with four degree of freedom (5%). All cointegrating regressions and the auxillary regressions are estimated with a constant and seasonal dummies.

also possibly at the biannual and annual frequencies. The fact that a time series is integrated at seasonal frequencies implies that it possesses long memory properties so that shocks tend to last permanently and moreover they tend to alter the seasonal pattern permanently. The finding that time series exhibit seasonal unit roots at different frequencies suggest that some series may be cointegrated at the seasonal frequencies. Cointegration established at different frequencies will lead to an interesting seasonal error correction model. Moosa (1995) points out that an error correction model will be misspecified if cointegration at the seasonal frequencies is present but is not accounted for.

In this paper we have endeavoured to investigate the seasonal properties of sectoral stock price indices at the KLSE, by applying the recent technique of seasonal unit root test proposed by Hylleberg *et al.* (1990). In our analysis, we also conducted the seasonal cointegration test proposed recently by Engle *et al.* (1993) to the sectoral stock price data. Generally, we found that stock price indices at the KLSE exhibit seasonal unit roots, not only at the zero frequency, but, in most cases at the biannual frequency. However, when tested for seasonal cointegration, our results suggest that sectoral stock price series are not seasonally cointegrated either at zero frequency (in the case of Industrial, Finance, Property, Agriculture and Tin) or the biannual frequency (in the case of Industrial, Finance, Property and Tin). An important implication of this study is that the informationally efficient stock market hypothesis cannot be rejected for the Kuala Lumpur Stcok Exchange for the period under study. This implies that investors cannot earn abnormal profit consistently using the returns of sectoral stock price to predict the returns of other sectoral stock price at the KLSE.

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