Causality between government revenue and expenditure in Malaysia: A seasonal cointegration test

Goh, Soo Khoon and Dawood, Mithani

CenPRIS, USM, Penang, Malaysia

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Causality between Government Expenditure and Revenue in Malaysia

A Seasonal Cointegration Test

D.M. Mithani and Goh Soo Khoon

The objective of this article is to empirically incorporate the effect of seasonality in examining the causal relationship between quarterly government revenue and government expenditure in Malaysia for the period 1970.1–1994.4. The seasonal integration and cointegration tests developed by Hylleberg, Engle, Granger and Yoo (1990) and extended by Engle, Granger, Hylleberg and Lee (1993) are applied prior to determination of causality. Evidence of seasonal cointegration of biannual frequency is found. The seasonal error correction model results indicate a unidirectional causal influence from government expenditure to government revenue, supporting the spend-and-tax hypothesis in the short run, i.e. higher government spending leads to higher taxes. The implication of this result is that the size and growth of the public sector and consequential tax burden as well as fiscal deficit in Malaysia are largely determined by the spending decision.

I. Introduction

In the contemporary era, after World War II, the rapidly increasing size of the public sector in most economies inspired a number of theoretical and empirical studies in explaining the causes of such growth through quantitative analysis. Basically, this type of research was prompted by the need to devise appropriate measures in order to reduce the U.S. federal budget deficit (Manzini and Nejadan 1995). Besides, causality links between government expenditure and revenue have unique significance for developing countries in making budgetary decisions. Governments in developing countries often face greater budgetary constraints. Usually, the government has to make a choice between two possibilities: either to raise taxes or to reduce expenditure in order to contain the fiscal deficit through necessary adjustments in its fiscal operations. This linkage issue is crucial in understanding the causes and consequences of budgetary deficit, which is commonly observed in most countries.

On a theoretical plane, there is no consensus among the economists on the issue of causal relations between government expenditure and tax revenue. Three alternative hypotheses have been advanced in perceiving the intertemporal budgetary links between taxation and government spending: (1) Spend-and-tax hypothesis (2) Tax-and-spend hypothesis, and (3) Fiscal synchronization hypothesis.

According to the spend-and-tax hypothesis, government starts spending first, and then determines how to finance the expenditure through additional taxes at a later date, so that causation runs from expenditure to revenue. Such a view is explored by Peacock and Wiseman (1967) who argue that external shocks (such as the Great Depression, World War II, oil shocks) which require temporary increases in government spending can lead to a permanent increase in taxes. The proposition that taxes adjust to changes in government expenditure is also substantiated in a study by Barro (1974), a supporter of the Ricardian Equivalence proposition, who argues that increases in taxes are the results of higher levels of fiscal expenditure; hence, causality runs from expenditure to revenue with no feedback. The tax-and-spend hypothesis, contrary to the previous hypothesis, posits the reverse relationship. It is argued that changes in taxes lead to changes in government spending. This hypothesis earned its popularity with the supply-side economists. Economists, such as Friedman (1982), argued that increases in taxes only result in increased expenditure, and not in deficit reduction. This view suggests a causal relationship from revenue to expenditure. The fiscal synchronization hypothesis, on the other hand, postulates that governments change expenditure and revenue simultaneously. According to this hypothesis, governments decide on the desired expenditure and taxes, by comparing the marginal benefits and costs of any balanced budget change.

Given this lack of theoretical consensus between expenditure and revenue decisions, the direction of the causal relation, if any, becomes a matter of empirical investigation. Going through the available literature, we observed that most empirical studies on this aspect have been confined to developed countries. Anderson, Wallace and Warner (1986) and Von Furstenberg, Green and Jeong (1986), for instance, found strong evidence of causality running from expenditure to revenue in U.S data. In contrast, Manage and Marlow (1986) and Ram (1988*a*) found causality running from revenue to expenditure in the United States, differing from Anderson et al.'s conclusion. Ram (1988*b*), however, observed that for the developed and less developed countries (22 samples) causality seems to run more frequently from revenue to expenditure. Based on the evidence of cointegration between revenue and expenditure in U.S federal data and using quarterly gross domestic product (GDP) data as a macroeconomic control variable, Baghestani and McNown (1992) observed that neither revenue nor expenditure respond to budgetary disequilibria. In their view, each expansion in revenue and expenditure is determined by long run economic growth reflected in rising gross national product (GNP). Huang and Tang (1992), in the case of Taiwan, on the basis of data for the period 1951–87, found that there is feedback between GNP and government expenditure, as well as government revenue and GNP, but there is only one–way causality running from government revenue to government expenditure. Detecting the seasonal movement in the quarterly data of government expenditure and revenue Barbaro, Craigwell, Leon and Mascoll (1994) applied seasonal unit roots tests prior to the standard cointegrating and Granger Causality tests. Their result indicates a unidirectional causal influence from government revenue to government expenditure.

In this article, we seek to test the alternative hypotheses about the tax-expenditure relationships in the Malaysian context. Our approach relates to seasonal components in the Malaysian data and tests for seasonal unit roots and seasonal cointegration as developed by Hylleberg, Engle, Granger and Yoo (1990, henceforth HEGY) and extended by Engle, Granger, Hylleberg and Lee (1993, EGHY) applied to quarterly government expenditure and revenue data in Malaysia from 1970.1 to 1994.4.[1]

The outline of this study is as follows: in section II we discuss the source of data and the methodology for the seasonal unit root tests and then we report the test results. In section III, we describe the seasonal cointegration methodology and show the test results. In section IV, we present the seasonal error correction model (ECM) estimates of the seasonally cointegrated series and then discuss the inferences that can be drawn on the basis of the empirical evidence. In section V we conclude with a discussion of the findings.

II. Data Base and Research Methodology

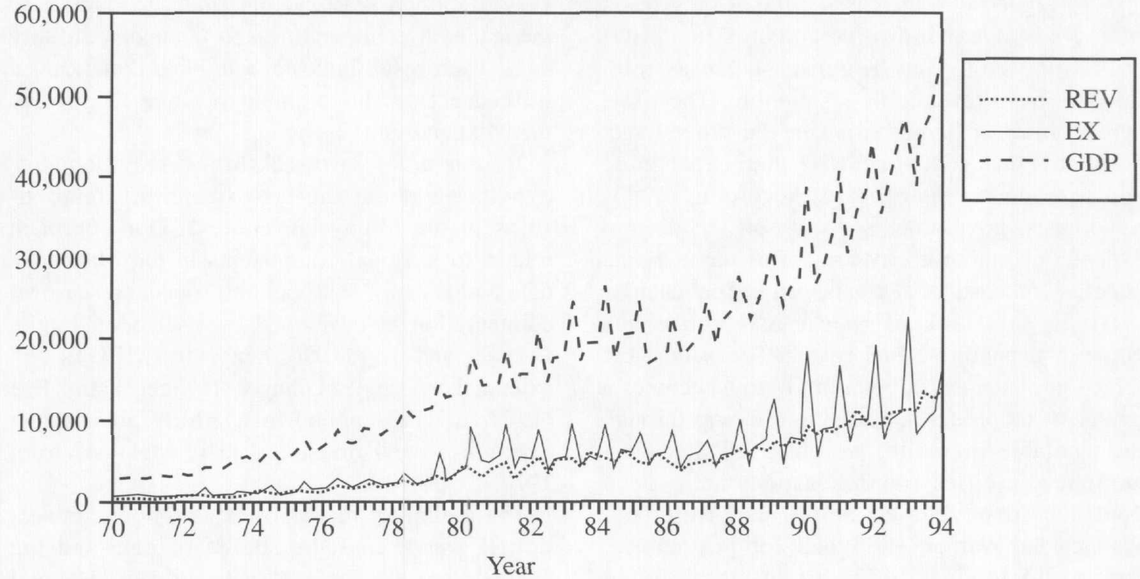
To test the causal relation between government expenditure and revenue in Malaysia, we used a data set spanning the years from 1970.1 to 1994.4, which includes government expenditure (E), revenue (R) and gross domestic product The quarterly data of government revenue and government expenditure were extracted from the data set available on International Financial Statistics CD-ROM, issued by the International Monetary Fund. Since GDP data are not available on a quarterly basis, derivation of this series from the component of GDP (as described in the Appendix) is resorted to. All the variables are expressed in nominal terms. Figure 1 plots the data for the three series. Each series shows a strong seasonality and nonstationarity. It is observed that the fourth quarter seasonal component is largest for the expenditure series. One of the possible reasons for the appearance of peak in the fourth quarter of the expenditure series is that most of the government departments fully utilize their budgetary allocation of funds only towards the end of fiscal year. There is, however, no consistent pattern to the plot of revenue series. Most importantly, Figure 1 suggests that seasonal cointegration among the three series might be present.

When seasonality is not an issue, the non-stationary economic time series are typically integrated of order d, I(d), i.e. differencing d times is required to make the series stationary. With seasonality, however, the order of integration may involve both seasonal and non-seasonal differences.

HEGY propose a set of tests for the presence of unit roots at the seasonal frequencies, as well as at

FIGURE 1

Government Revenue, Expenditure and GDP, 1970.1–1994.4



the zero frequency. One can test for a unit root (i.e. a root at zero frequency), unit root at a biannual frequency (i.e. a root at frequency π) or a unit root at annual frequency (i.e. a root at frequency $\pi/2$). The HEGY test is as follows:

$$\Delta_4 X_t = \pi_1 X_{1,t-1} + \pi_2 X_{\rho 2,t-1} + \pi_3 X_{3,t-2} + \pi_4 X_{3,t-1} + \sum_{j=1} \beta_j \Delta_4 X_{t-j} + \varepsilon_t \qquad (1)$$

where $X_{1t} = (1 + L + L^2 + L^3)X_t$
$X_{2t} = -(1 - L + L^2 - L^3)X_t$
$X_{3t} = -(1 - L_2) X_t$

The X_{1t} transformation removes all seasonal roots and leaves only the unit root at the zero frequency. The X_{2t} transformation leaves a unit root at the biannual frequency for quarterly data so that $(1 + L)X_{2t}$ is stationary and X_{3t} dispenses with all the roots except for a pair of complex roots at an annual frequency so that $(1 + L^2) X_{3t}$ is stationary.

The tests of the null hypothesis of unit roots are based on the t-statistics for π_1 and π_2 for the zero and the biannual frequency, respectively. For unit roots at the annual frequency, one needs to perform either the F-test of the null hypothesis that $\pi_3 = \pi_4 = 0$, or the t-statistic on π_3, when $\pi_4 = 0$. Not rejecting the null hypothesis implies that there is a unit root with an annual frequency. Critical values for the tests were tabulated by HEGY (1990). The corresponding three null hypotheses are not alternatives since a series may have nonseasonal, biannual, and annual unit roots.

The HEGY Test Results

In the HEGY testing, we begin with the most general model, including all the deterministic components (trend, seasonal dummies and intercept), a suitable lag structure for $\Delta_4 X_t$ (for example 4) and all the π_1's (it may be noted that in our model, Xt is referred to all variables, i.e. E,R and GDP). Thereafter, we leave out the non-significant autoregression (AR) parts and the deterministic part one by one, in different combinations. Following the work by Ermini et al. (1996), we use the notation $I_w(1)$ with $W = 0, \pi, \pi/2$ to indicate the presence of a unit root at each of the three frequencies respectively.

The outcomes of the HEGY test are shown in Table 1 and the summary of the results is presented in Table 2. The results indicate that the E and GDP series are a combination of non-seasonal, biannual and annual unit roots, that is they are integrated of order 1 at all frequencies 0, π and $\pi/2$ while the R series has a combination of non-seasonal and biannual unit root, that is this series is integrated of order 1 at the frequencies 0 and π. From Table 1, we notice that the results are not sensitive to the removal of deterministic seasonality. This implies that the seasonal components of R, E and GDP are stochastic rather than deterministic.

III. Seasonal Cointegration and Testing Results

The presence of seasonal unit roots implies that standard cointegration tests are inappropriate. The static cointegration regression does not necessarily give consistent estimates due to the presence of seasonal unit roots as shown in Engle, Granger and Hallman (1989). The problem of the standard cointegration technique in this context motivates an alternative approach to testing for cointegration in the presence of unit roots at other seasonal frequencies. A few attempts have been made in this regard by HEGY (1990) and EGHL (1993) to extend the usual cointegration technique to the case where the data have unit roots at both zero and seasonal frequencies. According to EGHL, a vector of series X_t, each of which is integrated at frequency W, $I_w(1)$, is said to be cointegrated at that frequency if a unique linear combination of them no longer exhibits unit roots at that same frequency. In other words, this relation is stationary. EGHL suggest that given prior information on which unit roots are present, seasonal unit roots can be filtered out before testing for cointegration on the filtered series.

We used OLS to estimate the bivariate and trivariate model at different frequencies using the corresponding unit root variable with or without deterministic components. The residuals from these cointegrating regressions are then tested for

TABLE 1
Tests for Seasonal Unit Roots for the series R, E and GDP, 1970.1–1994.4

Variable	*Auxiliary Regression*	t π_1	t π_2	t π_3	t π_4	F $\pi_3 \cap \pi_4$
R	I, SD, T	0.26	–1.39	1.62	–3.18 ***	6.49 *
	I, SD	1.86	–1.42	1.60	–3.18 ***	7.70 **
	I, T	0.23	–0.11	1.22	–2.59 ***	4.16 **
	I	1.93	–0.14	1.20	–2.59 ***	4.16 **
	—	2.62 **	–0.14	1.21	–2.16 **	4.21 **
E	I, SD, T	–2.32	–2.02	0.37	–1.36	1.01
	I, SD	–0.50	–2.01	0.32	–1.58	1.30
	I, T	–2.37	–1.26	0.27	0.18	0.03
	I	–0.43	–1.15	0.24	0.04	1.35
	—	1.26	–1.18	0.28	0.02	0.04
GDP	I, SD, T	0.80	–0.93	–0.14	–2.12	2.26
	I, SD	2.36	–1.06	–0.21	–2.27	2.58
	I, T	0.78	0.39	–0.33	–0.49	0.17
	I	2.48	0.40	–0.33	–0.49	0.17
	—	3.18 ***	0.40	–0.33	–0.50	0.17

NOTES: Variables R = Tax Revenue, E = Government Expenditure, GDP = Gross Domestic Product. The auxiliary regression were augmented by lagged values of the fourth difference of the regressand. I = intercept, SD = seasonal dummies, T = trend. The statistics are compared with the critical value provided in Hylleberg et al. (1990).
* significant at 10% ** significant at 5% *** significant at 1%

TABLE 2
A Summary of the Seasonal Unit Root Testing Results

Variable	*Zero*	*Biannual*	*Annual*
R	I_0 (1)	I_π (1)	$I_{\pi/2}$ (0)
E	I_0 (1)	I_π (1)	$I_{\pi/2}$ (1)
GDP	I_0 (1)	I_π (1)	$I_{\pi/2}$ (1)

stationarity. In the bivariate model, for example, let us denote the residuals obtained from regressing R_{1t} on E_{1t}, R_{2t} on E_{2t} and R_{3t} on E_{3t} and E_{3t-1} as U_t, V_t and W_t respectively. The test of non-cointegration at zero frequency can then be performed by an auxiliary regression on ΔU_t on U_{t-1} with or without deterministic parts and augmented by the necessary lagged values of ΔU_t to whiten the error term. A similar procedure can be performed to test non-cointegration at the biannual frequency. The auxiliary regression is represented in the following form:

$$\Delta U_t = \pi_1 (U_{t-1}) + \sum_{j=1}^{k} \beta_j \Delta U_{t-j} +$$

$$\text{deterministic components} + e_{1t} \qquad (2)$$

$$V_t + V_{t-1} = \pi_2 (-V_{t-1}) + \sum_{j=1}^{k} \beta_j (V_{t-j} + V_{t-j-1}) +$$

$$\text{deterministic components} + e_{2t} \qquad (3)$$

However, the test of non-cointegration at the annual frequency is more complicated. EGHL (1993) show that the auxiliary regression to be estimated is:

$$W_t + W_{t-2} = \pi_3 (-W_{t-2}) + \pi_4 (-W_{t-1}) + \sum_{j=1}^{k} \beta_j (W_{t-j} + W_{t-j-2}) +$$

$$\text{deterministic components} + e_{3t} \qquad (4)$$

and the F-value of the joint test $\pi_3 \cap \pi_4 = 0$ is computed together with the t values for $\pi_3 = 0$ and $\pi_4 = 0$.

EGHL (1993) suggest that the critical values used for the zero and biannual frequencies cointegration tests can be obtained by following Engel and Yoo (1987, Table 2, p. 157). These critical values are, however, not valid for seasonal cointegration at annual frequency because they are complex conjugates. EGHL provide the asymptotic associate t-statistics (see Table A1-A5 in EGHL, pp. 293–97). We actually do not need these statistics, since from Table 2 we know that cointegration at this frequency is clearly ruled out. From Table 2, we see that the only possibility of cointegration in the bivariate and trivariate models is in the long run (zero) frequency and biannual frequency, for only at these two frequencies are all variables integrated of the same order.

The seasonal cointegration tests of these two frequencies are, thus, performed. Tables 3 and 4 show the cointegration test results for the bi-variate and trivariate model. R and E are used as the left hand side variable for the choice of normalization. As shown in Table 4, we can reject the null hypothesis of no seasonal cointegration at the biannual frequency for the bivariate and trivariate model at the 1 per cent significance level, but not at the zero frequency.[2] Hence, we conclude that the data exhibit one seasonal cointegrating relationship at biannual frequencies only. Surprisingly, we are not able to reject non-cointegration at zero frequency since from Figure 1, we expect a long run relationship between R and E.

IV. Seasonal ECM

To examine the budgetary hypotheses, our main interest lies in the adjustment of revenue and expenditure to budgetary disequilibria. In addition, we shall draw inferences abut causality between time series on the basis of ECM. If two variables are cointegrated, causality must run in at least one direction between them, since one variable can help to forecast the other (Granger 1988). Short-run dynamics in the ECM are captured by the Error Correction Term (ECT) and other right hand side terms. Conventional tests of causality may be based on the significance of these terms. If the ECT is significant, the size of the coefficient will measure the speed at which the fiscal variable adjusts to restore budgetary equilibrium (Baghestani and McNown 1992).

HEGY (1990) gives the general theory of ECM in a seasonal time-series framework. It is shown that the seasonal ECM takes the following form:

$$A^*(B)\, \Delta_4 X_t = \gamma_1 \alpha_1' X_{1,t-1} + \gamma_2 \alpha_2' X_{2,t-1} - (\gamma_3 \alpha_3' - \gamma_4 \alpha_4') X_{3,t-2} + (\gamma_4 \alpha_3' + \gamma_3 \alpha_4') X_{3,t-1} + \varepsilon_t \qquad (5)$$

where $A^*(B)$ is an autoregressive matrix, $\alpha_{i's}$ are cointegration vectors at different frequencies in $n \times r_i$ matrices ($i = 1,2,3$) and the $\gamma_{i's}$ are the error-correction parameters (see HEGY, p. 232, EGHY, p. 281).

In the preceding section, we inferred that the data exhibit one seasonal cointegrating relationship at W= π only (for bivariate and trivariate model), which implies that Model 5 is now reduced to:

TABLE 3

Cointegration Test at Frequency 0: The Long Run

	Cointegration Regression				*Auxiliary Regression*		*Tests for Unit Roots in Residuals*	
Regressand	*Regressor R1t*		*Deterministic Components*	R^2	*Deterministic Part*	*Augmentation*	*Q (20)/Prob*	*DF* $t_{\pi l}$
E_{1t}	1.12		C	0.97	—	2	9.58 (0.97)	–1.51
E_{1t}	1.12		C	0.97	C,T	2	9.97 (0.96)	–1.08
E_{1t}	0.92		C,T	0.97	—	2	9.14 (0.98)	–2.13
E_{1t}	1.18		—	0.97	—	2,6	6.78 (0.99)	–0.23
	R_{1t}	GDP_{1t}						
E_{1t}	1.14	–0.06	C, T	0.97	—	1,2,5	6.54 (0.99)	–3.06
E_{1t}	1.14	–0.06	C,T	0.97	C,T	1,2,5	6.54 (0.99)	–2.86
E_{1t}	0.87	0.06	C	0.97	—	2	9.24 (0.98)	–1.42
E_{1t}	0.87	0.06	C	0.97	C,T	2	9.65 (0.97)	–1.02
	E_{1t}							
R1t	0.86		C	0.97	—	2	9.82 (0.97)	–1.12
R1t	0.86		C	0.97	C,T	2	10.12 (0.96)	–0.94
R1t	0.76		C,T	0.97	—	2	10.72 (0.95)	–0.63
R1t	0.83		—	0.97	—	2	10.16 (0.96)	–0.77
	E_{1t}	GDP_{1t}						
R1t	0.14	0.25	C,T	0.99	—	1,2,3,5	12.07 (0.91)	–5.10 ***
R1t	0.14	0.25	C,T	0.99	C, T	1,2,3,5	12.29 (0.90)	–5.06 ***

NOTES: The tests are based on the ordinary augmented Dickey-Fuller regression $\Delta U_t = \pi_1(U_{t-1}) + \Sigma\beta_j\Delta U_{t-j}$ + deterministic components + e_{1t} where Ut is the residual from the cointegration regression. The Q(20)/Prob in column 7 is the Box-Pierce Q-statistics which detects serial correlation with 20 degree of freedom. Parentheses in this column refer to the level of the probability. The "t" statistic in the last column is distributed as described in Engle and Yoo (1987).

*** significant at 1%.

TABLE 4

Cointegration Tests at Frequency $^1/_2$: Biannual

	Cointegration Regression				Auxiliary Regression		Tests for Unit Roots in Residuals	
Regressand	*Regressor* R_{2t}		*Deterministic Components*	R^2	*Deterministic Part*	*Augmentation*	*Q(20)/Prob*	*DF* $t_{\pi 2}$
E_{2t}	4.10		C	0.76	—	2	15.1 (0.70)	–5.56***
E_{2t}	3.33		C,SD	0.78	—	2	15.6 (0.74)	–4.85***
E_{2t}	3.98		—	0.75	—	2	17.5 (0.62)	–5.98***
	R_{2t}	GDP_{2t}						
E_{2t}	0.68	0.82	C	0.86	—	2	17.5 (0.62)	–5.98***
E_{2t}	0.61	0.78	C,SD	0.86	—	2	15.3 (0.76)	–4.84***
	E_{2t}							
R_{2t}	0.18		C	0.76	—	2	14.7 (0.79)	–6.11***
R_{2t}	0.15		C,SD	0.78	—	2	15.2 (0.76)	–6.39***
R_{2t}	0.18		—	0.74	—	2	17.1 (0.64)	–6.88***
	E_{2t}	GDP_{2t}						
R_{2t}	0.03	0.17	C	0.85	—	2	14.9 (0.78)	–5.20***
R_{2t}	0.02	0.16	C,SD	0.85	—	2	15.7 (0.73)	–5.37***

NOTES: The tests are based on the auxiliary regression $V_t + V_{t-1} = \pi_2(-V_{t-1}) + \Sigma\beta_j (V_{t-j} + V_{t-j-1})$ + deterministic components + e_{2t} where V_t is the residual from the cointegration regression. The Q(20)/Prob in column 7 is the Box-Pierce Q-statistics which detects serial correlation with 20 degree of freedom. Parentheses in this column is the level of the probability. The "t" statistic is distributed as the "Dickey-Fuller" in Engle and Yoo (1987).

*** significant at 1%.

$$A^*(B)\Delta_4 X_t = \gamma_2 \alpha_2' X_{2,t-1} + \varepsilon_t \quad (6)$$

Table 5 reports the OLS regression estimates of both bivariate and trivariate models of the ECM in restricted form — allowing 4 lags on the differences of all variables and specifications omitting insignificant terms.[3] We did not estimate an equation for GDP since GDP may be determined by factors other than fiscal policy variables. All restricted ECMs pass the diagnostic tests to determine that the residuals are white noise (based on the inspection of the Box-Pierce Q-statistics).

Using either residual RER and RER* (from the OLS cointegrating regressions of R_{2t} on E_{2t}, R_{2t} on E_{2t} and GDP_{2t}, both with the deterministic components), the major finding from the ECM estimates in Table 5 is the lack of statistical significance of the budgetary disequilibrium term in expenditure equations. It appears that only revenue adjust to correct budgetary disequilibria with the correct sign. The six-monthly error correction term has the value of –0.31 (in bivariate model) and –0.24 (in trivariate model).

Granger (1988) shows that the causal impact of one variable on another can take place in an ECM in two ways: first, through the impact of lagged changes in the independent variable; second, through the ECT, which itself is a function of lagged levels of the variables. Our next interest lies in the causality results implied by the ECT and by the lagged difference terms. As can be seen from the bivariate model (columns 1 and 2, Table 5), the ECT is significant only in the equations with revenue regressed on expenditure, and significant lagged changes in expenditure appear in the revenue equation, no significant lagged changes in revenue appear in the expenditure equation. With GDP (in the trivariate model), the conclusion remains the same. The causality results indicate unidirectional causality from expenditure to revenue, thus, support the spend-and-tax hypothesis, i.e. changes in government spending lead to changes in taxes.

TABLE 5
Seasonal Error-correction Models for Revenue, Expenditure and GDP, 1970.1–1994.4

	Bivariate Model		*Trivariate Model*	
Dep Variable	$\Delta_4 R$	$\Delta_4 E$	$\Delta_4 R$	$\Delta_4 E$
C	223.79 (3.26)	352.02 (2.94)	129.06 (1.88)	124.05 (0.82)
RER(–1)	–0.31 (–2.67)	0.09 (0.44)		
RER*(–1)			–0.24 (–2.1)	0.27 (1.17)
Δ_4R(–1)	0.41 (3.91)		0.34 (3.19)	
Δ_4E(–2)		0.29 (2.70)	–0.11 (–1.96)	0.27 (2.55)
Δ_4E(–3)	0.09 (1.95)			
Δ_4GDP(–2)			0.14 (3.53)	
Δ_4GDP(–3)				0.16 (2.45)
R^2	0.40	0.08	0.44	0.14
Q(20)/Prob	10.4 (0.96)	15.77 (0.73)	12.63 (0.89)	15.17 (0.76)

NOTES: RER is the residual series from the OLS cointegrating regression of R_{2t} on E_{2t} with the deterministic components. RER* is the residual series from the OLS cointegration regression of R_{2t} on E_{2t}, GDP_{2t} and with the deterministic components. Numbers in the parentheses are the value of t-ratios. The Q(20)/Prob is the Box-Pierce Statistics which detect serial correlation with 20 degree of freedom; the parentheses in this row is the level of probability.

V. Concluding Remarks

In this study, the seasonal unit-root testing and cointegration methodology as developed by HEGY (1990) and extended by EGHL (1993) is used to examine behavioural relationships between tax revenue (R) and government expenditure (E) on the basis of quarterly data in Malaysia from 1970.1 to 1994.4. Unit roots were found at the zero frequency for all the series (Revenue, Expenditure and GDP), but seasonal unit roots at the annual frequency and biannual were found only for E and GDP. Seasonal cointegration was established at zero and biannual frequencies. The cointegration tests reject the existence of a long run relation between R and E in Malaysia at zero frequency, but there is evidence of cointegration at biannual frequency. Based on the error correction model estimates, however, we found only revenue responds to budgetary disequilibria and a uni-directional causality running from expenditure to revenue is confirmed. The result is consistent with the spend-and-tax hypothesis that expenditure leads revenue changes. This means that the public sector in Malaysia is largely determined by spending decision. Any attempt, thus, to reduce the size of government should be accomplished by a reduction in spending, and not in tax collection.

Our results also confirm the procedures adopted by the Malaysian fiscal authorities that the desired fiscal adjustment in historical perspectives (e.g. in the 1980s) has been mainly through public expenditure to bring down the fiscal deficit. Little adjustment has been made to revenue because of its limited scope. This may be due to several constraints, such as Malaysia's revenue dependence on the export of oil and a gradual decline in non-oil revenue (Jaafar 1989).

The policy implication of the validity of the spend-and-tax hypothesis in the Malaysian budgetary process is that to contain a fiscal deficit or lower the tax burden, it is essential that the government expenditure should be reduced. The positive fiscal effect of lowering government spending is at least two-dimensional: (1) it reduces the quantum of fiscal deficit in the first instance, and (2) it eventually paves the way for tax reduction.

A noteworthy feature of the 1997 Malaysia fiscal policy was to assume a "middle path" in the budgetary exercise with a view to avoid fiscal deficit and maintain surplus budget (as since 1994). However, the Malaysian Government is keen to pull the country out of the recession induced by the 1998 currency crisis and economic turmoil. Thus policy makers in framing the 1999 Budget have resorted to a modified Keynesian approach of deficit budgeting — raising government expenditure together with several tax exemptions — as a fiscal device to stimulate the economy, with an emphasis on infrastructure development through public sector initiative and support. Incidentally, in 1998 the government budget envisaged a cut in government spending by 20 per cent; without raising taxes, but ended up with rising expenditure following the counter-cyclical measures. The decline in revenue coupled with higher expenditure have resulted in the government deficit of RM9.59 billion or 3.7 per cent of the GNP in 1998 compared with a surplus of RM6.63 billion in 1997 (Economic Report, 98/99). This kind of fiscal behaviour of the government is apparently indicative of the existence of the spend-and-tax hypothesis on the operational side of the budget in the country. The problem of accumulated fiscal deficits in future, thus, will have to be dealt with both by a substantial cut in government spending and raising tax revenue.

APPENDIX DERIVATION OF QUARTERLY GDP

Since GDP is not available in quarterly series, we derived it by approximation of the quarterly data of major components of the GDP, viz.: private consumption, private investment, government expenditure and net exports from other data bases (the relevant data are taken from *Statistics Quarterly Bulletin* of Bank Negara, various issues). To obtain the quarterly figures for the private consumption (which is only available in the annual form in Malaysia), we multiplied the proportion of the import of consumption goods for each quarter with the annual private consumption expenditure. The process of calculation is illustrated below:

Year	Import on Consumer Goods	(%)	Private Consumption (Annual = 44,856m)
1988.1	2,039.7	(21.96)	9,850.01
1988.2	2,209.8	(23.79)	10,671.45
1988.3	2,531.0	(27.25)	12,222.57
1988.4	2,508.1	(27.00)	12,111.98

The same method is employed for the other components of GDP identity, and their approximation bases are mentioned below:

Component of GDP	Approximation Bases
Private Investment	Quarterly series of Industrial Production Index
Public Consumption	Quarterly series of Federal Government current expenditure
Public Capital Formation	Quarterly series of Federal Government direct expenditure
Export of Goods and Services	Quarterly series of gross exports (f.o.b)
Import of Goods and Services	Quarterly series of gross imports (c.i.f)

The summation of quarterly figures of private consumption, private investment, public consumption and public capital formation plus exports of goods and services and minus imports of goods and services lead to the quarterly GDP in nominal terms.

NOTES

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1. For want of time and data, the analysis is confined up to the year 1994. Though this may be a limitation of this study, however, the conclusions derived are less likely to differ even when the data are updated.
2. In Table 3, we can reject the null hypothesis when using R as the left hand side variable, E_{1t} and GDP_{1t} as the right hand side variable. But we fail to reject the null hypothesis when using E_{1t} as the left hand side variable. This suggests that there is no cointegration at this frequency.
3. The common practice in the selection of the appropriate lag length is to choose a relatively long lag length (some researcher prefer to use 4 or 8 lags, very few would use 12 lags) and pare down the model by the usual t-test or F test. Most important is that once a tentative lag length has been determined, diagnostic checking by the Box-Pierce Q-statistic or some other tests should not appear in any strong evidence of structural change or serial correlation. One can see from our result in Table 5 that all the seasonal error correction models pass the Box-Pierce Q statistic.

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Dawood Mithani, Ph.D., is senior lecturer, School of Economics, Northern University of Malaysia (UUM), Sintok, Malaysia. **Goh Soo Khoon** is lecturer, School of Economics, Northern University of Malaysia (UUM), Sintok, Malaysia.