Exports, export composition and growth: cointegration and causality evidence for Malaysia

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This paper comprehensively tests the export-led growth (ELG) hypothesis for Malaysia for the period 1955–90, using cointegration and causality testing based on Hsiao’s synthesis of the Granger test and Akaike’s minimum final prediction error criterion. The results provide support for the ELG hypothesis; aggregate exports Granger-cause real GDP and non-export GDP. This relationship is found to be driven by manufactured exports rather than by traditional exports.

I. INTRODUCTION

Empirical testing for developing countries of the ‘export-led growth’ (ELG hereafter) have a long pedigree. The work originally concentrated on simple correlations between exports and income (Emery, 1967; Kravis, 1970). It then progressed to a consideration of the appropriate variables to be correlated (Michaely, 1977; Heller and Porter, 1978). Tyler (1981) and Feder (1982) subsequently estimated aggregate production functions that included exports as an explanatory variable for crossections of developing countries. With the creation of time series data bases and the development of new time series techniques (such as the cointegration technique), work followed the pioneering study of Jung and Marshall (1985) and Chow (1987). They have sought to investigate causality issues. Indeed, the last few years have witnessed an explosion of studies on different countries or aspects of the export-growth relationship using various time series techniques (for relevant references, see Section II).

What contribution can a further paper on the subject make? At the crudest level, it should be noted that none of the studies is specific to Malaysia. More importantly, this paper seeks to provide a comprehensiveness in terms of the application of ‘good practice’. The time period used here is 1955–90; a considerably larger period than that used by some earlier studies (see, e.g., Dodaro, 1993; Greenaway and Sapsford, 1994a). The present work distinguishes between the ‘economic’ and ‘accounting’ effects of exports on growth; a distinction missing in many recent contributions. It also investigates the effects of decomposing aggregate exports into traditional and non-traditional categories; a consideration omitted by most recent papers. Finally, it seeks to use a comprehensive econometric methodology which investigates nonstationarity, structural break, cointegration, causality and error-correction characteristics of the data.

The aim of this paper, therefore, is to synthesize the existing empirical work and to comprehensively examine the relationship between exports and output growth, using time series data for Malaysia. The remainder of the paper is organized as follows. Section II reviews the issues and recent empirical evidence on the ELG hypothesis. Section III provides a brief review of the development of the Malaysian economy in the post-War period and of the role of the export sector in that development. Section IV sets out the econometric methodology used. The data and empirical results are described in Section V. The final section provides a discussion of the implications of the results, and some summary conclusions.

II. REVIEW OF THE EMPIRICAL LITERATURE

The idea that export growth is one of the major determinants of output growth (i.e. ‘export-led growth’ hypothesis, ELG hereafter) is a recurrent idea in economics. Export
growth might affect output growth through a number of channels. The exports sector may generate positive externalities on nonexports sector, through more efficient management styles and improved production techniques. Export growth may increase the scope for economies of scale in exporting firms and encourage the allocative efficiency and dynamic competitiveness. If there are incentives to increase investment and improve technology, this would imply a productivity differential in favour of the export sector (i.e. marginal factor productivities are expected to be higher in the exports sector than in the rest of the economy). Thus, an expansion of exports, even at the cost of other sectors, will have a positive net effect on aggregate output. Export expansion may also affect aggregate growth by relaxing the foreign exchange constraint. By allowing an increase in imports of intermediate inputs, export expansion relaxes a crucial bottleneck and facilitates the export of inputs embodying recent techniques. Export growth may expand output via the foreign trade multiplier. This argument is based on a short-run Keynesian macromodel which is not strictly relevant to long-run economic growth. Nonetheless, inadequate domestic demand may constrain industrialization in the context of developing countries. Finally, the above arguments have recently been supplemented by the literature on ‘endogenous’ growth theory which emphasizes the role of increasing returns to scale and the dynamic spill-over effects of the export sector’s growth. Exports may increase long-run growth by allowing the economy to specialize in those sectors with scale economies that arise from R&D, human capital accumulation, or learning-by-doing. In this framework, increasing returns to scale are associated with the use of new technology and with the complementarities between human and physical capital.

Despite the popularity of the ELG hypothesis, the empirical evidence is rather mixed. The recent time-series evidence fails to provide uniform support for the ELG hypothesis whereas a substantial literature, applying a range of cross-section type methodologies, rigorously supports an association between exports and growth. In other words, cross-section results appear to find a close and fairly robust relationship, while time-series results are less conclusive (for reviews of exports and growth literature, see, for example, Jung and Marshall, 1985; Ahmad and Kwan, 1991; Edwards, 1993; GharTEy, 1993; Greenaway and Sapsford, 1994a, 1994b; Love, 1994). As a result, the validity of the ELG hypothesis has been brought into question, contrary to the strong earlier empirical support. Recent papers, notably those by Jung and Marshall (1985), Chow (1987), Durrant (1987), Kwan and Cotomitis (1990), Afxentiou and Serletis (1991a), Ahmad and Kwan (1991), Bahmani-Oskooee et al. (1991), Kugler (1991), Dodaro (1993), Oxley (1993), Greenaway and Sapsford (1994a), Love (1994), cast some doubt on the validity of the ELG hypothesis. Others, notably those by Afxentiou and Serletis (1991b), Bahmani-Oskooee and Alse (1993), GharTEy (1993), Khan and Saqib (1993), Kugler and Dridi (1993), Dutt and Ghosh (1994), Sengupta and Espana (1994), Ghatak et al. (1995), provide fairly robust evidence which support the ELG hypothesis.

Most of the time-series studies just mentioned employ the Granger or Sims methods in order to investigate the causal relationship between real export growth and real economic growth. Only a few of them (see Bahmani-Oskooee et al., 1991; GharTEy, 1993; Giles et al., 1993; Oxley, 1993; Love, 1994) follow the approach of C. Hsiao (1979, 1981) which combines the Granger test with the Akaike’s Final Prediction Error (FPE) criterion to determine the optimal lag length in the Granger causality test. Hsiao’s synthesis allows determination of the optimal lag length for each of the variables employed in the Granger test on the criterion of minimum FPE, and thus avoids ambiguity in the arbitrary choice of the lags.

Furthermore, with the few exceptions (Afxentiou and Serletis, 1991b; Kugler, 1991; Bahmani-Oskooee and Alse, 1993; Giles et al., 1993; Kugler and Dridi, 1993; Oxley, 1993; Dutt and Ghosh, 1994; Sengupta and Espana, 1994), recent contributions do not consider whether exports and income are cointegrated. Thus, previous results do not necessarily imply that there exists a ‘genuine’ relationship between the long-run development of exports and output (either GDP or GNP), as they may arise from a purely short-run relationship. A necessary precondition to causality testing is to check the cointegrating properties of the variables under consideration since standard tests for causality are not valid if there exists cointegration (see, especially, Granger, 1988; Bahmani-Oskooee and Alse, 1993). Accordingly, most previous studies have not dealt with the problem of non-stationarity of the exports and income data properly. This resulted in some ‘spurious’ regressions in the literature. Differencing has been proved not to be an appropriate remedy due to possible loss of valuable long-run information in the data. Cointegration analysis and error-correction modelling are recommended as an effective remedy. Error-correction models establish causality between two variables after reintroducing long-run information (through the error-correction term) into the analysis. Note that if two variables are cointegrated, then the Granger causality would run in, at least, one direction. In other words, cointegration implies causal effects (Engle and Granger, 1987).

One important problem arises from the fact that exports, via the national accounting identity, are themselves a component

1 Kugler (1991) and Kugler and Dridi (1993) apply the multivariate cointegration methodology with GDP, consumption and investment on the one hand, and exports on the other, whereas others use standard bivariate cointegration methodology between output and exports. In addition, Ghatak et al. (1995), by employing the multivariate cointegration methodology, provide evidence for Turkey that suggests a stable, joint, long-run relationship among real GDP per capita, export volume index, human and physical capital proxies in accordance with the ‘endogenous’ growth theory.
of output. To remedy this problem, it is necessary to separate the ‘economic influence’ of exports on output, from that incorporated in the ‘growth accounting’ relationship mentioned above (see, e.g. Greenaway and Sapsford, 1994a). It is surprising that this rather important issue is generally ignored in the existing literature.2 Recently, this issue has been raised by Hwa (1988), Sheehey (1990, 1992), Afxentiou and Serletis (1991b), Greenaway and Sapsford (1994a), Love (1994), who indicated that the problem of autocorrelation arises in models which employ any of the major components of output such as exports or government expenditures (or revenues) as the determinant of output growth. There are several methods of dealing with this issue. Heller and Porter (1978) suggest that this problem may be remedied by defining the income variable net of the variable alleged to be the source of growth. Sheehey (1990, 1992) tested the Heller and Porter connection in a Feder/Ram type bivariate model and found that export growth and then government expenditure growth were positively and statistically significant related to income growth but that the coefficients on both explanatory variables lost statistical significance when the dependent variable was measured in net terms. Similarly, using the models of Falvey/Gemmell and Heller and Porter, Greenaway and Sapsford (1994a) find little support for the ELG hypothesis for 13 developing countries.

Another shortcoming of the recent time-series literature, with the exception of Giles et al. (1993) and Ukpolo (1994), is that they focus on ‘aggregate’ real exports only. As stressed by Giles et al. (1993), this may mask important underlying differences between different export categories. Even if there is evidence of ELG relating to certain groups of export goods, this may not be reflected at the aggregate level, and spurious conclusions may be drawn when disaggregated data are not examined. In the case of the New Zealand economy, Giles et al. (1993) provides mixed evidence in support of the ELG hypothesis. While they reject the hypothesis at the aggregate level, there is some support for it in the case of certain export groups, especially if attention focuses on exports of minerals, chemicals and plastic materials; exports of metal and metal products; and (to a lesser degree) exports of live animals and meat. Ukpolo (1994), on the other hand, examines the hypothesis that economic growth is linked to export composition with the use of time-series data for some low-income African countries. His findings support the hypothesis of a positive linkage between the growth of non-fuel primary exports and growth. However, the results cast some doubt on the significance of the positive contribution of the manufactured exports sector to the growth of the low-income African countries. The results also suggest that the existence of a critical level of development (i.e. ‘threshold effect’) is crucial for the attainment of economic growth through the adoption of manufactured-product ELG policies. It may be difficult for LDCs to promote manufactured-product exports to the extent necessary to have a significant impact on economic growth.4

### III. EXPORT PERFORMANCE AND DEVELOPMENT IN MALAYSIA

Since independence in 1957, Malaysia has followed a relatively open economy, generally driven by the market forces. During the colonial period, the Malaysian economy was mainly supported by tin and rubber industry. Between 1964–1984, along with increasing diversification, real GDP grew by an impressive 6–8 percent p.a. Malaysia went into recession during 1985–1986 as commodity prices collapsed. Since 1987, manufactured exports registered strong growth. In 1991, real GDP grew by 8.6 percent (see World Bank, 1993).

Although the Malaysian economy mainly relies on the market forces, government played an important role in the economy, both as a producer and a regulator. Malaysia also encourages foreign investment in the export sector. Indeed, US and Japanese multinationals dominate a substantial section of the manufacturing sector. To create a favourable macroeconomic environment, the government has so far followed a prudent fiscal policy, with a surplus in the operating account and a small deficit in the capital budget, funded mainly by bond sales. Recently, the government is also pre-paying foreign debt. National savings as a percentage of GDP reached 30 percent in 1991. Similarly, the use of monetary policy has so far been targeted to achieve price

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2 Heller and Porter (1978), together with Michaely (1977, 1979), suggested that those results obtained from studies of export growth and output growth may be largely a statistical artefact since exports are frequently a substantial (this is also the case for Malaysia!) component of domestic output.

3 The two-sector model, developed by Feder (1982) and extended by Ram (1987), subdivides the economy into two sectors: export oriented sector and non-export oriented or traditional sector. The model allows exports to influence output growth through two channels: externality effect and productivity differential effect. Thus, this model, in principle, allows researchers to derive estimates of the separate externality and productivity differential effects. A modified version of the Feder/Ram model is presented by Falvey and Gemmell (1989). Particularly, they simplify somewhat the task of isolating the separate productivity differential and externality effects from the estimated parameters of the model. The Appendix of Greenaway and Sapsford (1994a) provides a formal statement of the Feder model, together with the Falvey/Gemmell modification.

4 According to the World Bank Development Reports, most LDCs that experienced growth through expansion of manufactured exports are primarily middle-income countries. Manufactured export volume rose annually by only 2.4 percent from 1965 to 1973, 8.7 percent from 1973 to 1985 and 15 percent in 1989 in low-income countries and by 14.9 percent, 12.9 percent and 43 percent respectively in middle-income countries.
### Table 1. ADF test for unit roots

<table>
<thead>
<tr>
<th>Variables</th>
<th>Test statistic</th>
<th>Levels</th>
<th>1st differences</th>
</tr>
</thead>
<tbody>
<tr>
<td>LY</td>
<td>0.43</td>
<td></td>
<td>−3.88</td>
</tr>
<tr>
<td>LNY</td>
<td>−2.20</td>
<td></td>
<td>−3.25</td>
</tr>
<tr>
<td>LX</td>
<td>−0.02</td>
<td></td>
<td>−6.33</td>
</tr>
<tr>
<td>LXm</td>
<td>−0.31</td>
<td></td>
<td>−4.91</td>
</tr>
<tr>
<td>LXf</td>
<td>−0.53</td>
<td></td>
<td>−3.14</td>
</tr>
<tr>
<td>LXp</td>
<td>1.06</td>
<td></td>
<td>−2.45</td>
</tr>
<tr>
<td>LK</td>
<td>−1.50</td>
<td></td>
<td>−3.99</td>
</tr>
<tr>
<td>LH</td>
<td>−2.31</td>
<td></td>
<td>−3.29</td>
</tr>
</tbody>
</table>

The reported critical values in our tables are obtained from Charemza and Deadman (1992), and correspond to 30 number of observations. The augmentation of one or two, generally, appear to be sufficient to secure lack of autocorrelation of the error terms.

Critical values: at 5% − 2.26
At lower level: at 10% − 1.80
At upper level: at 5% − 2.05
At 10% − 1.64

stability by controlling monetary aggregates, the rate of interest, reserve requirements and sometimes, open market operations. A tight control of public sector deficit as proportion to GDP helped the creation of a macroeconomic environment within which commerce and industry flourished considerably (see Table 1).

As regards exchange rate policy, Malaysia has so far followed an open foreign exchange regime as reflected in low rates of distortions between the black market and the official exchange rate. Occasionally, the Central Bank intervenes in the market to smooth the fluctuations of the Malaysian Ringgit against, say US dollar. The long run stability of Ringgit in the international currency market reflects its real underlying value rather than direct interventions to stimulate exports. Foreign payments (remittances) are freely permitted. Most prices (excepting fuel, rice, sugar, flour, public utility goods and tobacco) are market determined. In comparison with most LDCs, Malaysia experienced a fairly stable rate of inflation. Between 1980–1991, manufacturing exports registered strong growth among all the export categories. Both short and long-term debt/GDP ratio declined substantially between 1987 and 1991 (see World Bank Reports, 1992). Average tariffs account for only 15 percent of prices on a tradeweighted basis. However, in the agricultural sector, there are important tariff and non-tariff barriers to encourage domestic production.

### IV. ON ECONOMETRIC METHODOLOGY

In the light of our previous discussion about export growth and GDP growth, we now set out a model to test rigorously the long-run relationship and causality issues. Standard tests for causality (i.e. Granger/Sims tests) are only valid if the original time series, from which growth rates are obtained, are not cointegrated. In case of cointegration, however, any causal inferences will not be valid (Granger, 1988; Bahmani-Oskooee and Alse, 1993). To make the process valid, one should include the relevant error-correction term, obtained from the cointegrating regression, in the standard causality tests. Therefore, in this section, the export-promotion (i.e. ELG) hypothesis is empirically tested for Granger causality with error-correction models (ECMs).

Accordingly, our methodology consists of two steps. Since it is essential to check for the cointegrating properties of the original real exports and output series before applying the standard Granger test, *in the first step*, considering the cointegration analysis in the Engle–Granger sense (see Engle and Granger, 1987), we estimate the following cointegration regressions by ordinary least squares (OLS):

\[
LY_t = \alpha_0 + \alpha_1 LX_t + u_t \quad (1a)
\]

\[
LNY_t = \lambda_0 + \lambda_1 LX_t + u'_t \quad (1b)
\]

where \( LX \), \( LY \) and \( LNY \) denote the natural logarithms of real exports, real GDP and non-export real GDP respectively. Let us now apply integration and cointegration analyses in the Engle–Granger sense. Accordingly, a time series, say, \( X_t \) is said to be integrated of order \( d \) if, after differencing \( d \) times, it becomes stationary, denoted as \( X_t \sim I(d) \). Moreover, two time series, \( X_t \) and \( Y_t \), are said to be cointegrated of order \( d \), \( b \) where \( d > b > 0 \), denoted as

\[ X_t, Y_t \sim CI(d,b), \text{ if:} \]

a) both are \( I(d) \), and
b) their linear combination \( a_1 X_t + a_2 Y_t \) is \( I(d-b) \); that is, the residuals of the long-run regression should be stationary (i.e. integrated of order zero). The vector \( [a_1, a_2] \) is

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5 Exchange rate distortion index (ERDI) of Malaysia constructed on the basis of the formula \( \text{ERDI} = (\text{BMS}/\text{OFS})/\text{BMS} \) where BMS and OFS represent annual average black market and official exchange rates both expressed in domestic currency per US dollar. The data for BMS and OFS are taken from [Cowitt, P. P., *World Currency Yearbook* (previously known as *Pick's Currency Yearbook*), various issues, International Currency Analysis, Brooklyn] and [IMF, *International Financial Statistics, Yearbooks* (various issues) and 1972 supplement] respectively. The ERDI, here is like the black market premium used by others as it measures the difference between the BMS and OFS as a proportion of BMS [see Edwards (1992), Ghatak et al. (1995)]. Note that, in LDCs, BMS-OFS is generally observed to be positive.

6 We prefer using the natural logarithms of the variables in consideration, since their first differences reflect the rate of change of each variable. In addition, we assume that the explanatory variables are weakly exogenous in Equations 1a and 1b. This is because it is believed that the value of each explanatory variable is determined outside the system and thus, independent of the error term.
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referred to as the 'cointegrating vector' (see Engle and Granger, 1987). Thus, we employ the Augmented Dickey–Fuller (ADF hereafter) test and the residual-based ADF test to determine the integration level and the possible cointegration between the variables respectively.\(^7\) To see if any (exogenous) outlier changes the integration level of the series under consideration, we apply Perron’s additive outlier integration level test with structural break (see Perron, 1990; Perron and Vogelsang, 1992).\(^8\)

The second step involves testing whether export growth Granger causes output growth. As noted earlier, the relevant error-correction term is included in the standard procedure if the variables are found to be cointegrated. Otherwise, the standard Granger test is carried out without the error-correction term.\(^9\) The ECM is formulated as follows:

\[
DLY_t = \beta_0 - \beta_1 u_{t-1} + \sum_{i=1}^{m} c_i DLY_{t-i} + \sum_{j=1}^{n} d_i DLX_{t-j} + e_t (2a)
\]

\[
DLNY_t = \theta_0 - \theta_1 u'_{t-1} + \sum_{i=1}^{q} e_i DLNY_{t-i} + \sum_{j=1}^{p} f_j DLX_{t-j} + e'_t (2b)
\]

where \(u_{t-1}\) and \(u'_{t-1}\) are the lagged residuals (i.e. error-correction terms) obtained from the cointegrating regressions Equation 1a and 1b respectively. The latter \(D\) represents the first differences. Statistically significant \(\beta_1\) and \(\theta_1\) suggest that exports Granger cause output and non-export output respectively. The ECMs introduce an additional channel through which Granger causality could be detected. This is because if two variables are cointegrated, there should be a causal relationship between them. This causal relationship between the two variables provides the short-run dynamics necessary to obtain long-run equilibrium (Granger, 1988). For instance, focussing on Equation 2, exports are said to Granger cause output not only if the \(d_i\)'s are jointly significant, but also if \(\beta_1\) is significant. Thus, unlike the standard Granger test for causality, the ECMs allow for the finding that exports Granger cause output as long as the error-correction term, \(u_{t-1}\), carries a significant coefficient with negative sign.\(^10\) Jones and Joulaian (1991) suggest the interpretation that the changes of the lagged independent variable describe the short-run causal impact, while the error-correction term captures the long-run effect.

To determine the optimal lag lengths in the Granger causality tests (see Equation 2a and 2b), we use the approach of C. Hsiao (1979, 1981) which combines the Granger test with the Akaike’s minimum Final Prediction Error (FPE) criterion to avoid the ambiguity in the arbitrary choice of lags.

In addition to the above ‘aggregate’ export analysis, we also focus on the possible effects of ‘disaggregated’ exports (i.e. separate export categories, such as manufacturing, fuel and non-fuel primary products) on the real GDP and non-export real GDP. Accordingly, we adopt the following aggregate production functions which are also compatible with the ‘new’ growth theory (see especially Fosu, 1990; Uko, 1994, and also see Balassa, 1978; Tyler, 1981; Feder, 1982; Kavoussi, 1984; Ram, 1987; Sheehy, 1990, 1992; Ghatak et al., 1995):

\[
Y = f_1 (K, H, Xm, Xf, Xp) (3a)
\]

\[
NY = f_2 (K, H, Xm, Xf, Xp) (3b)
\]

where \(Y\) represents real GDP; \(K\) and \(H\) are measures of physical capital (proxied by the real gross domestic investment as percentage of real GDP) and human capital (proxied by the enrolment ratio in primary + secondary school); \(Xm, Xf\) and \(Xp\) represent real exports of manufactured products, fuel and non-fuel primary products respectively. Let us now express the functions Equation 3a and 3b as linear logarithmic regression form and apply multivariate cointegration technique:

\[
LY_t = \alpha_0 + \alpha_1 LK_t + \alpha_2 LH_t + \alpha_3 LX_m + \alpha_4 LX_f + \alpha_5 LX_p + u_t (4a)
\]

\[
LN_{Y_t} = \lambda_0 + \lambda_1 LK_t + \lambda_2 LH_t + \lambda_3 LX_m + \lambda_4 LX_f + \lambda_5 LX_p + u'_t (4b)
\]

where we again use the natural logarithms of the variables (prefixed with the letter \(L\)), since their first-differences reflect the rate of change of each variable.\(^11\) In the first step, by

\(^7\) Haug (1993) compares seven different residual-based tests for cointegration with the Monte Carlo method. Among the tests considered, Engle-Granger’s residual-based ADF test shows the least size distortion.

\(^8\) Perron (1990) suggests two types of model to test for unit roots with structural break, namely, the Additive Outlier Model (AOM) and the Innovational Outlier Model (IOM). The AOM is recommended for ‘sudden’ structural changes while the IOM is said to be more appropriate for ‘gradual’ structural changes. We believe that structural changes in the economy are generally ‘sudden’ in nature. This is also thought to be the case for Malaysian economy. Thus, the AOM is preferred in this study.

\(^9\) Note that if two time series are cointegrated, then there would exist an error-correction representation (i.e. error-correction mechanism is well determined) and vice versa. This is known as the Granger Representation Theorem (GRT). Empirically, in small samples, statistically significant error-correction terms provide further evidence in favour of the presence of a ‘genuine’ long-run relationship. It is also true that, by definition, cointegration implies causal effects.

\(^10\) Note that, to avoid an explosive process, the coefficient should take a value between \(-1\) and 0.

\(^11\) We assume, again, that the explanatory variables in long-run regressions Equations 4a and 4b are weakly exogenous. This is because it is believed that the value of each of the variables is determined outside the system and thus, independent of the error term.
In this case, there is no longer a unique long-run relationship among the variables. One drawback of the Engle–Granger cointegration analysis is that it assumes uniqueness of the cointegrating vector. However, in a multivariate context, such as regression Equations 4a and 4b, the number of cointegrating vectors could be more than one.\(^{12}\) The Johansen Full Information Maximum Likelihood (ML) method can be used to check for the number of cointegrating vectors (if there is any!) (see Johansen, 1988; Johansen and Juselius, 1990).\(^{13}\) As a check for the robustness of the cointegrating estimates we also employ Saikkonen’s method which provides asymptotically efficient estimates (see Saikkonen, 1991). Once a joint cointegration among variables is confirmed, the next step would be to model the short-run with the use of ECM. The existence of joint cointegration among the variables in Equation 4a and 4b and statistically significant error-correction terms with negative signs in ECM models, would be regarded as robust evidence in favour of export-led growth hypothesis.

Another shortcoming of the Engle–Granger method is that, due to non-normality of the distribution of the estimators of the cointegrating vector, no sensible judgement can be made about the significance of the parameters. As a remedy, Engle and Yoo (1991) propose a ‘three-step’ estimation technique to overcome this shortcoming of the classical two-step Engle–Granger method. The third step corrects the parameter estimates of the first step so that standard tests, such as the \(t\)-test, can be applied (for further details, see Engle and Yoo, 1991). In this paper, we follow the same approach to make more sensible judgements on the significance of the explanatory variables in our model. This type of search will also give us a chance to be more specific on what categories of export (i.e. manufactured products, fuel or non-fuel primary products) are the driving forces (if there exists any!) of the export-led growth for the Malaysian case.

### V. DATA AND EMPIRICAL RESULTS

**Aggregate data analysis**

In the light of the econometric methodology presented in the previous section, we now check for the cointegrating properties of the variables involved, i.e. natural logarithms of real GDP and real exports in Equation 1a, and non-export real GDP and real exports in Equation 1b. Accordingly, we apply integration and cointegration analyses in the EG sense. The Malaysian data used in this study are annual for the period 1955–1990 and are taken from various yearbooks of IMF International Financial Statistics.

The visual inspection of the variables in hand (i.e. \(LY, LX\) and \(LNY\)) suggests that they are all nonstationary in levels, but stationary in first differences. The relevant sample autocorrelation functions (i.e. correlograms) are also supportive in the same direction (graphs are available on request). We then apply the Augmented Dickey–Fuller test for unit roots as a formal test. The ADF test results for unit roots (see Table 1) confirm that all variables are integrated of order one in levels but integrated of order zero (i.e. stationary) in first differences [denoted as \(LY \sim I(1), LX \sim I(1),\) and \(LNY \sim I(1)\)].\(^{14}\)

As regards non-export real GDP (\(LNY\)) for the period 1955–1990, we observe a decline after 1985. Since this could imply a structural break in the non-export real GDP of Malaysia, we apply the Additive Outlier Perron test for unit roots with structural break (for details of the test, see Perron, 1990; Perron and Vogelsang, 1992). The small sample critical values of the Additive Outlier Perron test was recently tabulated by Rybinski (1994). As a result, our ADF unit root test results for \(LNY\) are validated despite some structural changes that have occurred after 1985 (see Table 2). That is, lack of ‘spurious’ unit root resulting from structural change for \(LNY\) is ensured.

Let us now estimate the EG cointegrating regressions Equation 1a and 1b by OLS, and test for the stationarity of the residuals by applying the residual-based ADF test. Indeed, i.e. if \(LY \sim I(1)\) and \(LX \sim I(1)\), in order for \(LY\) and \(LX\) to be cointegrated, \(u_t\) should be \(I(0)\). Table 3 reports the results of the residual-based ADF test for cointegration for both Equation 1a and 1b. As regards Equation 1a, we are able to reject the null of \(I(1)\) at 5% significance level, i.e. the variables are said to be cointegrated. However, results for Equation 1b are more complicated to evaluate and appear to be inconclusive [low Cointegrating Durbin–Watson (CRDW) statistic is especially noticeable]. This could be due to a sudden decline in \(LNY\) after the 1985 period. It is

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\(^{12}\) In this case, there is no longer a unique long-run relationship towards which the error-correction model (ECM) is adjusting. Then, a single equation cointegration regression will estimate the linear combination of existent vectors. Many researchers in applied econometrics, overcome this problem by choosing the cointegrating vector which make ‘economic sense’. This implies choosing the cointegrating vector where the estimated long-run elasticities correspond closely (in both magnitude and sign) to those predicted by economic theory.

\(^{13}\) It should be noted that the Engle-Granger and Johansen methods are grounded with different econometric methodologies and thus, cannot be directly compared. The Engle-Granger method is a single equation-based modelling while the Johansen method is a system approach based on VAR modelling. Charemza and Deadman (1992, p. 201) suggest that the Johansen method can be used for single equation modelling as a supplementary tool. Assume, for instance, that the Johansen method indicates the existence of a unique cointegrating vector. In this case, as pointed out by Charemza and Deadman, if the estimated cointegrated coefficients have, after normalization, economically sensible signs and approximately similar in size to those estimated by the Engle-Granger method, this could be regarded as some confirmation of the single equation model to which the Engle-Granger method was employed.

\(^{14}\) Econometric computations in this study have been carried out by Microfit 3.0 version.
Note that, in our case, the error-correction term, \( u_{t-1} \), should be taken from the cointegrating regression with dummy variable.
The intercept term is not included in the residual-based ADF equations. The reported slopes and the t-statistics are the corrected and thus valid test statistics obtained by the use of Engle–Yoo third-step correction procedure (Engle and Yoo, 1991). Our original Engle–Granger estimates show only small bias compared with the reported Engle–Yoo slopes. All Engle–Yoo t-statistics reported belong to the explanatory variable, \( L^2X \). \( DU \) is the dummy variable created in line with Perron (1990). CRDW represents the Cointegrating Durbin–Watson statistic.

Table 3. The residual-based ADF test for cointegration: aggregate data

<table>
<thead>
<tr>
<th>Cointegrating regression</th>
<th>Slope (Engle–Yoo)</th>
<th>t-statistic (Engle–Yoo)</th>
<th>Adjusted ( R^2 )</th>
<th>Calculated ADF</th>
<th>Critical value</th>
<th>Lower-value 5%</th>
<th>Upper-value 5%</th>
</tr>
</thead>
<tbody>
<tr>
<td>( L^2 = f(L^2X) )</td>
<td>1.11</td>
<td>84.6</td>
<td>1.02</td>
<td>-3.39</td>
<td>-2.87</td>
<td>-2.85</td>
<td>-2.65</td>
</tr>
<tr>
<td>( L^2N = f(L^2X) )</td>
<td>1.02</td>
<td>16.9</td>
<td>0.45</td>
<td>-1.96</td>
<td>-2.87</td>
<td>-2.85</td>
<td>-2.65</td>
</tr>
<tr>
<td>( L^2N = f(L^2X, DU) )</td>
<td>1.28</td>
<td>21.4</td>
<td>0.88</td>
<td>-3.38</td>
<td>-3.32</td>
<td>-2.97</td>
<td>-3.16</td>
</tr>
</tbody>
</table>

Table 4. Johansen maximum likelihood (ML) procedure: cointegration likelihood ratio (LR) test to determine the number of cointegrating vectors (\( r \)), based on maximal eigenvalues of the stochastic matrix

<table>
<thead>
<tr>
<th>Cointegrating regression</th>
<th>Null hypothesis</th>
<th>Alternative hypothesis</th>
<th>Test statistic</th>
<th>Critical value at 5%</th>
<th>Critical value at 10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>( L^2N = f(L^2X, DU) )</td>
<td>( r = 0 )</td>
<td>( r = 1 )</td>
<td>24.01</td>
<td>22.00</td>
<td>19.77</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 1 )</td>
<td>( r = 2 )</td>
<td>11.91</td>
<td>15.67</td>
<td>13.75</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 2 )</td>
<td>( r = 3 )</td>
<td>3.39</td>
<td>9.24</td>
<td>7.73</td>
</tr>
<tr>
<td>( L = f(L^2m, LXf, LXp, L^2K, L^2H) )</td>
<td>( r = 0 )</td>
<td>( r = 1 )</td>
<td>66.61</td>
<td>39.37</td>
<td>36.76</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 1 )</td>
<td>( r = 2 )</td>
<td>24.98</td>
<td>33.46</td>
<td>30.90</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 2 )</td>
<td>( r = 3 )</td>
<td>17.83</td>
<td>27.07</td>
<td>24.73</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 3 )</td>
<td>( r = 4 )</td>
<td>15.27</td>
<td>20.97</td>
<td>18.60</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 4 )</td>
<td>( r = 5 )</td>
<td>11.03</td>
<td>14.07</td>
<td>12.07</td>
</tr>
<tr>
<td>( L^2N = f(L^2m, LXf, LXp, L^2K, L^2H) )</td>
<td>( r = 0 )</td>
<td>( r = 1 )</td>
<td>74.68</td>
<td>39.37</td>
<td>36.76</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 1 )</td>
<td>( r = 2 )</td>
<td>49.61</td>
<td>33.46</td>
<td>30.90</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 2 )</td>
<td>( r = 3 )</td>
<td>16.87</td>
<td>27.07</td>
<td>24.73</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 3 )</td>
<td>( r = 4 )</td>
<td>14.99</td>
<td>20.97</td>
<td>18.60</td>
</tr>
<tr>
<td></td>
<td>( r &lt; 4 )</td>
<td>( r = 5 )</td>
<td>10.01</td>
<td>14.07</td>
<td>12.07</td>
</tr>
</tbody>
</table>

Table 5. Granger causality test from error-correction model (ECM): aggregate data

<table>
<thead>
<tr>
<th>Dependent variable</th>
<th>Coefficient for ( EC(-1) )</th>
<th>t-statistic for ( EC(-1) )</th>
<th>F-statistic for ( \Sigma DLX )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( DL^2Y )</td>
<td>-0.50</td>
<td>-4.03&lt;sup&gt;a&lt;/sup&gt;</td>
<td>5.01(4)&lt;sup&gt;b&lt;/sup&gt;</td>
</tr>
<tr>
<td>( DL^2N )</td>
<td>-0.34</td>
<td>-3.61&lt;sup&gt;b&lt;/sup&gt;</td>
<td>NS</td>
</tr>
</tbody>
</table>

\( EC(-1) \) denotes the error-correction term. Number in the bracket indicates the maximum optimal lags determined by the use Akaike’s FPE criterion. \( D \) represents first-differences. We only report \( F \)-statistic for explanatory lagged variables in first-differences. NS means ‘no significant lag found’. The error-correction term for the short-run equation with \( DL^2N \) as the dependent variable comes from the cointegrating regression with dummy variable, \( DU \).

<sup>a</sup> means ‘significant at 1% level’,
<sup>b</sup> means ‘significant at 5% level’.

In a multivariate context, maximum number of cointegrating vectors can be \( r = N - 1 \) where \( N \) represents the number of variables in cointegrating regression. For the critical values reported by Microfit 3.0 version, see Osterwald–Lenum (1992). Both the \( \lambda \)-max and the Trace statistics have been used for testing the number of co-integrating vectors.

involved. However, as far as the EG cointegrating regression results are concerned, we still face the following problems: a) The number of cointegrating vectors may be more than one, since we have more than two variables involved in each cointegrating regression. Because the EG approach assumes the uniqueness of the cointegrating vector, we need to employ a system-based Johansen method to check for the number of cointegrating vectors. b) OLS long-run estimates may be remarkably biased. c) Resulting \( t \)-statistics may not be valid due to nonnormality of the distribution.

To solve the first problem, we applied the Johansen method. For cointegrating regression \( L^2X = f(LXm, LXf, LXp, L^2K, L^2H) \) one can reject the null \( r = 0 \) against the alternative \( r = 1 \), but cannot reject the null \( r \leq 1 \) against the alternative \( r = 2 \), that is, there exists a ‘unique’ cointegrating vector (see Table 4). As regards \( L^2N = f(LXm, LXf, LXp, L^2K, L^2H) \), however, evidence, at first glance, suggest two cointegrating
Table 6. The residual-based ADF test for cointegration: disaggregated data for exports

<table>
<thead>
<tr>
<th>Cointegrating regression</th>
<th>Adjusted $R^2$</th>
<th>CRDW</th>
<th>Calculated ADF residuals</th>
<th>Critical value</th>
<th>Lower-value 5%</th>
<th>Lower-value 10%</th>
<th>Upper-value 5%</th>
<th>Upper-value 10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>$LY = f(LX_m, LX_f, LX_p, LK, LH)$</td>
<td>0.99</td>
<td>1.64</td>
<td>-4.10</td>
<td>-4.22</td>
<td>-3.83</td>
<td>-4.05</td>
<td>-3.65</td>
<td></td>
</tr>
<tr>
<td>$LNY = f(LX_m, LX_f, LX_p, LK, LH)$</td>
<td>0.96</td>
<td>1.84</td>
<td>-5.08</td>
<td>-4.22</td>
<td>-3.83</td>
<td>-4.05</td>
<td>-3.65</td>
<td></td>
</tr>
</tbody>
</table>

The intercept term is not included in the residual-based ADF equations. An augmentation of one appeared to be sufficient to secure lack of autocorrelation of the error terms.

vectors. Although the existence of multiple cointegrating vectors is regarded as an identification problem for single equation cointegrating estimation, this problem, in practice, may be solved by choosing the particular cointegrating vector where the long-run estimates correspond closely (in both magnitude and sign) to those predicted by economic theory and also to those obtained by some other alternative long-run estimation techniques. Accordingly, in the case of $LNY = f(LX_m, LX_f, LX_p, LK, LH)$, the number of cointegrating vectors is reduced to one (those results of the Johansen procedure and of some alternative long-run estimation procedures are available on request). As regards $LNY = f(LX_m, LX_f, LX_p, LK, LH)$, we can also assume that there is only one economically sensible cointegrating vector among the variables considered. To solve the second and third problems, we use the Engle–Yoo third step corrections. Table 7 reports the long-run estimates and $t$-statistics for the Engle–Yoo method for cointegrating regression Equation 4a and 4b. As mentioned earlier, the Engle–Yoo correction produces unbiased long-run estimates and statistically valid $t$-statistics. As a whole, it is important to note that, in both Equation 4a and 4b, manufacturing exports turn out to be the most significant explanatory variable together with physical capital measure. Another interesting finding is the negative effect of the exports of non-fuel primary products on the real GDP and non-export real GDP. This is, we believe, due to the well-known pattern of structural changes in the economy undergoing rapid industrialization.

Since the existence of joint cointegration among the variables in long-run regressions Equation 4a and 4b is confirmed, the next step would be to model the short-run with the use of ECM. In order to save some degrees of freedom due to small sample size, this time, we decided to use the first lags (only) as the maximum lag length instead of applying Akaike’s FPE criterion to determine them. Those lags with insignificant parameter estimates are, naturally, eliminated in the process. It is important to note that both short-run models with the use of ECM produce statistically significant coefficient estimates with negative signs for the corresponding error-correction terms (see Table 8). Since we take the contemporaneous effects into consideration in both short-run models, we report the relevant $t$-statistics and the coefficient estimates of the error-correction terms for both the OLS and the IV estimation methods. We also found that most of the explanatory variables (in first-differences) in short-run models are statistically significant either in the first-lag or in the contemporaneous level (or both) (those results are not reported, but available on request). Overall, the findings imply the righthand side variables in both ECM models jointly Granger cause real GDP growth and non-export real GDP growth respectively. Following Jones and Joulfaian (1991) we can suggest an interpretation that the changes of the lagged independent variables describe the short-run causal impact, while the statistically significant error-correction term introduces the long-run effect.

To sum up, those results with disaggregated real exports data appear to suggest that we can use (especially) manufactured exports as the major engine of growth. Since Malaysian economy has already reached a certain critical level of economic development, one may expect to see manufactured-product export promotion policies working...
Table 8. Granger causality test from error-correction model (ECM): disaggregated data for exports

<table>
<thead>
<tr>
<th>Dependent variable</th>
<th>Coefficient for EC(-1)</th>
<th>t-statistic for EC(-1)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>OLS</td>
<td>IV</td>
</tr>
<tr>
<td>DLY</td>
<td>-0.36</td>
<td>-0.31</td>
</tr>
<tr>
<td>DLNY</td>
<td>-0.71</td>
<td>-0.71</td>
</tr>
</tbody>
</table>

\(^{EC(-1)}\) denotes the error-correction term. \(D\) represents first-differences. OLS and IV represent the Ordinary Least Squares and the Instrumental Variable estimations respectively. \(^a\) means ‘significant at 5% level’, \(^b\) means ‘significant at 1% level’.

EC(-1) denotes the error-correction term. \(D\) represents first-differences. OLS and IV represent the Ordinary Least Squares and the Instrumental Variable estimations respectively. \(^a\) means ‘significant at 5% level’, \(^b\) means ‘significant at 1% level’.

efficiently. Thus, it seems reasonable for Malaysia to continue to promote manufactured-product exports to the extent necessary to have a significant impact on real economic growth. This, in a sense, marks a significant ‘product innovation’ effect for Malaysian exports on its real GDP. Results are also validated for non-export real GDP.

VI. CONCLUSIONS AND IMPLICATIONS

Time series studies of ELG have proliferated in recent years, and the results for this work have not been as consistently supportive of the ELG hypothesis as cross-sectional analysis. This pattern of results may not be very surprising, if some countries have experienced growth with export growth, some low growth with low export growth, and still others growth associated with non-export factors. Cross-sections, especially for particular samples of countries, may be supportive of ELG, whereas the outcome of time series testing may be country-dependent. Of course another reason for apparently inconsistent time series evidence may be variability in the time series methodology. Here we seek to be both comprehensive and apply ‘good practice’ to a relatively long time series. Thus when we find support for the ELG hypothesis for aggregate exports, this is not only due to the ‘accounting’ effect. The results indicate that real (aggregate) export growth also Granger-causes non-export real GDP growth for Malaysia and for this time period. We also find this to be a ‘genuine’ long run relationship, rather than possibly a spurious short term relationship, since there is robust support for cointegration between exports and GDP.

It turns out in the case of Malaysia that this aggregate support for the ELG hypothesis is driven by the relative importance of non-traditional exports in total exports. Herein may be a reason for the failure of some studies to find support for the hypothesis for other countries when using aggregate data, i.e. where non-traditional exports are relatively unimportant. Certainly one would anticipate that non-traditional export growth might have a greater effect on output and operate through more channels than traditional exports would. Indeed for Malaysia we find a significant negative, causal effect for some traditional exports (non-fuel primary exports) on GDP and non-export GDP. One could envisage traditional export growth which appreciates the real exchange rate and squeezes out non-traditional exports. If there is a positive causal link between non-traditional exports and output growth, then traditional export growth will in turn ‘crowd in’ output growth. This possibility is consistent with the evidence for Malaysia for non-traditional, manufactured exports, which are shown to be an engine of growth and to have an influence which is reflected in aggregate exports because they achieve sufficient importance within the sample period used. The evidence for other countries needs to be re-investigated therefore, in the light of this result.

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REFERENCES

Exports, export composition and growth


