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EXCHANGE RATE VARIABILITY AND THE EXPORT DEMAND FOR MALAYSIA'S SEMICONDUCTORS: AN EMPIRICAL STUDY

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ABSTRACT

This paper examines the effects of exchange rate variability on export demand for semiconductors, which is the largest sub-sector of electronics industry in Malaysia as reported by MIDA (Malaysian Industrial Development Authority, 2004). The empirical results, which are estimated based on the Johansen's multivariate co-integration tests and error correction model, suggest that there is a unique long-run relationship among quantities of export, relative price, real foreign income, and real exchange rate variability. The major finding of this paper is that the variability of real exchange rate has some effect on semiconductor exports in both the long-run and the short-run. In the light of rapid advances in technology in the global markets for electronics products, the findings are useful to policy makers for the design and target of appropriate exchange rate and industrial policies to enhance the export competitiveness of semiconductor industry.

Keywords: Exchange rate volatility; semiconductor exports; Malaysia;

JEL classification codes: C32; F14

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I. INTRODUCTION

It is an essential fact that exports sector plays a substantial role as 'engine of growth' in a country's industrialization i.e. exports led growth strategy. For the case of Malaysia, Ghatak *et al.* (1997) found aggregate exports Granger-cause real GDP and non-export GDP. Also, it is interesting to highlight that this relationship is driven by manufactured exports rather than by traditional exports. In this relation, Malaysia's semiconductor industry plays a significant role in the country's economy in terms of export trade and industrialization process, and also has the prospect of moving up the industrialization scale for Malaysia. In 2003, the semiconductor industry accounted for US$20.7 billion or 42.8% of the country's total electronics exports, which is Malaysia's leading non-resource-based export-oriented industry (MIDA, 2004). Besides, Malaysia is among the world's largest exporters of semiconductor devices (e.g. linear and digital integrated circuits, memories and microprocessors, opto-electronics, discrete devices, hybrids and arrays). A large part of the industry has been dominated by multinational corporations (MNCs), which use Malaysia as a production base for exports to their home countries or third markets. Among the MNCs, are Intel, Motorola, Agilent, AMD, National Semiconductor, Fairchild Hitachi, NEC, Fujitsu. Toshiba, Infineon and STMicroelectronics, lie in assembly, testing and packaging of semiconductors (Matrade, 2002).

Despite the semiconductor industry is highly internationally linked, it is found to be vulnerable to external shocks (or influences) such as the economic slowdown of Malaysia's major trading partners, a rapid technological change in the global markets for electrical and electronic products, and an increase in Ringgit exchange rate variability, which could increase the exchange rate uncertainty associated with international transactions in semiconductor exports. In this context, a study by Doraisami (2004) on the factors which caused the slowdown in export growth occurred in all East Asian economies that were affected by the Asian currency crisis, found that misaligned exchange rates, and the vulnerability of the downturn in the electronic cycle could also be major factors leading to poor Malaysia's export performance. She found a unique long-run relationship exists among the electronic cycle, US/yen dollar rate, and US total new orders for electronics. Even though there is empirical consensus on the effects of exchange rate variability on export and import flows (see Asseery and Peel, 1991; Bahmani-Oskooee and Ltaifa, 1992; Arize, 1996a and 1996b; Arize *et al*., 2000; Abbott *et al*., 2001), the effects of exchange rate variability on semiconductor exports should not be overlooked, especially for Malaysia, otherwise the export demand model can be subject to misspecification error due to 'omitted variable'. In line with this concern, Bahmani-Oskooee and Ltaifa (1992) found that the exchange rate uncertainty is unfavorable to the exports of both developing and developed countries. And, they found that the developed countries' exports are less sensitive to exchange risk than that of developing countries.

The main objective of this paper is to examine the effects of exchange rate variability on export demand for semiconductors, which is the largest sub-sector of electronics industry in Malaysia. The empirical evidence on the impact of exchange rate variability on trade flows is mixed. The results of various studies supporting the proposition that trade can be impeded by exchange rate variability can be found in Coes (1981), Cushman (1983; 1986; 1988), Akhtar and Hilton (1984), Kenen and Rodrik (1986), Thursby and Thursby (1987), Brada and Mendez (1988), Caballero and Corbo (1989), Koray and Lastrapes (1989), Pere and Steinherr (1989), Pozo (1992), Bahmani-Oskooee and Ltaifa (1992), Arize (1996a; 1996b), Hassan and Tufte (1998) and Arize *et al*. (2000). On the other hand, empirical findings by Hooper and Kohlhagen (1978), Gotur (1985), Needham (1986), Bailey *et al*. (1987), Asseery and Peel (1991) and Abbott *et al.* (2001) failed to detect a significant association between exchange rate variability and trade flows. As pointed out by Arize (1996b), the conflicting evidence was partly due to the differences in methodology, sample period and estimation techniques. For example, in Needham's (1986) econometric study, the estimation period covered was the first quarter of 1975 to the second quarter of 1985. Australia only floated her dollar in December 1983 and hence, very few of the more volatile post-float exchange rate's experience was included in the sample. Another possible explanation is that most empirical studies only look at export flows at the aggregate level and therefore, may fail to capture the effects at the individual export industry level.

Broadly speaking, the contributions made by this paper are follows. Firstly, it attempts to estimate the demand for Malaysia's semiconductor exports by SITC (Standard International Trade Classification) product group in the long -run and short run by using cointegration and ECM techniques. A literature survey shows that the available evidence is limited for the case of Malaysia's semiconductor industry. Secondly, the empirical study, which considers exchange rate variability as a potential determinant of export demand at disaggregated level (by industry or product group), is not available. Thus, this paper provides an extra dimension to the existing literature. Thirdly, estimating disaggregated export demand model can reduce the risk of aggregation bias in its regression estimates. Lastly, the estimated regression models can provide important implications for policy formulation and analysis for the semiconductor industry.

The rest of this paper is organized as follows. Section II discusses the specification of the export demand model with theoretical considerations of key determinants of export demand. It also addresses the use and availability of the data followed by the application of appropriate econometric methods and procedures to undertake the empirical study. Section III reports the results. The main conclusions and the policy implications are presented in Section IV.

II. MODEL SPECIFICATION, DATA AND METHOD

Determinants of Export Demand

The standard export demand function relates the rest of the world demand for a country's exports to the world income positively, and to the relative price of a country's export price over world export prices, negatively. This paper extends the standard specification of the demand for exports by further incorporating exchange rate variability. Bahmani-Oskooee and Kara (2003, p.296) included the nominal effective exchanger rate as a potential determinant of export demand; if a currency depreciation is to stimulate exports, a negative sign of its estimated parameter is expected, and *vice versa*.

However, it is interesting to note that the theoretical question of whether exchange rate variability has an impact on export flows is still contentious. There are situations in which the exchange rate variability could have negative or positive effects on exports. The outcome of these effects is basically dependent on the availability of hedging mechanism (Sercu and Vanhulle, 1992), the limitations and costs of forward exchange markets (Medhora, 1990; Caporale and Doroodian, 1994), and the degree of risk aversion (De Grauwe, 1988). For instance, if exporting firms are risk-averse and hedging is either expensive or impossible, increases in exchange rate variability could hamper export flows. This is because unpredictable exchange rate movements can create uncertainty about future profits from export trade and, as a result of risk aversion and future profit uncertainty, these firms are inclined to shift away from more risky export markets. Hence, this would result in lower volume of trade. On the other hand, these firms can reduce or avoid the exchange rate uncertainty in the short -run by using the forward exchange markets to manage the timing of their international transactions. However, hedging against adverse exchange rate movements is not sufficient to eliminate the effect of variability in the long -run because the maturity of the forward exchange contracts is relatively short and the exchange rate variability in the medium and longer term creates uncertainty which exporters and importers cannot predict the receivables and payables of their foreign exchange transactions over an extended period of time. Another limitation of the forward exchange markets is the size of the contract. Caporale and Doroodian (1994) pointed out that in the case of US-Canadian trade, they must have an average of US1 million per contract before hedging can take place. So, these limitations of the forward exchange markets indicate the difficulties for international trade firms in planning the magnitude and timing of foreign exchange receipts and payments with use of foreign exchange markets. Moreover, De Grauwe (1988) argued that the effects of exchange rate variability on exports depend on the degree of risk aversion. If the degree of risk aversion increases, an exporting firm will raise exports in response to the depressing effect of a decline in export earnings due to higher exchange rate variability. However, Baldwin and Krugman (1989) and Dixit (1989a, 1989b) had shown using hysteretic models of international trade that exchange rate variability could also influence international trade, in particular if significant sunk costs (i.e. high fixed costs associated with establishing an export market) were involved in international transactions, even when exporters are risk neutral.

On the basis of the conceptual framework of exchange rate variability discussed above, and following the existing studies (Arize, 1996a and 1996b; Arize *et al*., 2000; Abbott *et al*. 2001; Bredin *et al.* 2003), the long-run export demand equation for semiconductors can be written in log-linear form as:

$$lnQ_t = \beta_0 + \beta_1 lnRP_t + \beta_2 lnFY_t + \beta_3 EV_t + e_t \quad (1)$$

where Q_t is the export quantity of semiconductors demanded based on SITC product group (i.e. SITC 776) from the electronics sector, RP_t represents relative price (i.e. the unit price of semiconductor exports at SITC 776 deflated by the price index of similar products of Malaysia's major export competitors), FY_t is foreign real income, and EV_t is the exchange rate variability, which is a proxy for exchange rate uncertainty. It was constructed based on the moving-sample standard deviation of real effective exchange rate (REER) (see Appendix A). The parameters of β_1, β_2 and β_3 are price, income and exchange rate variability elasticities, respectively.

It is a conventional practice to assume equation (1) is homogeneous of degree zero in prices in the long-run by specifying it as a function of relative price rather than two separate price terms. This assumption is further applied in this paper because the price homogeneity hypothesis of whether Malaysia's semiconductor export prices match its export competitors' export prices one for one in the long-run cannot be tested due to the data unavailability of the latter in published sources. However, the imposition of the price homogeneity restriction on equation (1) could help to reduce multicollinearity between the price terms and conserve the degrees of freedom.

Furthermore, it is expected that the estimated parameter of relative price, β_1 has a negative sign. By assuming all else constant, as price of semiconductor exports rises relative to prices of similar goods produced by Malaysia's major export competitors, the lower price competitive is Malaysia's exports in the world markets, the lower is the quantity demanded for Malaysia's semiconductor exports. Tentatively, equation (1) ignores non-price competitive factors such as product innovations and product improvements because the data cannot be readily captured at both aggregated and disaggregated levels. Despite, to certain extent, non-price competitive factors can be jointly 'proxied' by a linear time trend variable, it also capture other factors such as the progress of technology, financial liberalization, and so on. This paper does not further incorporate these 'qualitative' factors by a fact that the effects of these factors have been channeled to the demand for exports through the relative price and the exchange rate variability.

In accordance to theory, β_2 is expected to be positive because other things being equal, if FY_t increases, the greater is the demand by foreign consumers for Malaysia's made goods. For example, an economic boom in the economies of Malaysia's major trading partners tends to increase their quantity demanded for Malaysia's exports of semiconductors, which in turn means Malaysia can generate more output, income and employment. This is in line with the argument that export trade of semiconductors can be an "engine of growth" for Malaysia. In the light of theory that cannot determine the sign of β_3 (which can explain the relation between semiconductor exports and exchange rate variability), this paper will examine the effects of exchange rate variability on export demand for semiconductors in long -run and short -run using Johansen's multivariate cointegration and ECM techniques.

Data Sources and Measures

The data used in this paper are expressed in logarithmic terms for the estimation regressions. The data are measured in indices (i.e.1995 = 100) to ensure all variables are unit free except for the exchange rate variability (EV) (see, Adler, 1946; Houthakker and Magee, 1969; Goldstein and Khan, 1978; Muscatelli *et al.,* 1992 and 1995; Hassan and Tufte, 1998; Abbott and DeVita, 2002). The sample period is from 1990 to 2001, which gives effectively 48 observations. The choice of this sample period is based on the availability of data, especially the disaggregated series for semiconductor exports by SITC product group. These domestic series, which are unpublished, were collected from the *Department of Statistics*, Malaysia. The data for foreign real incomes and the proxies for foreign unit export prices for semiconductor products, which have already been deseasonalized, were obtained from the

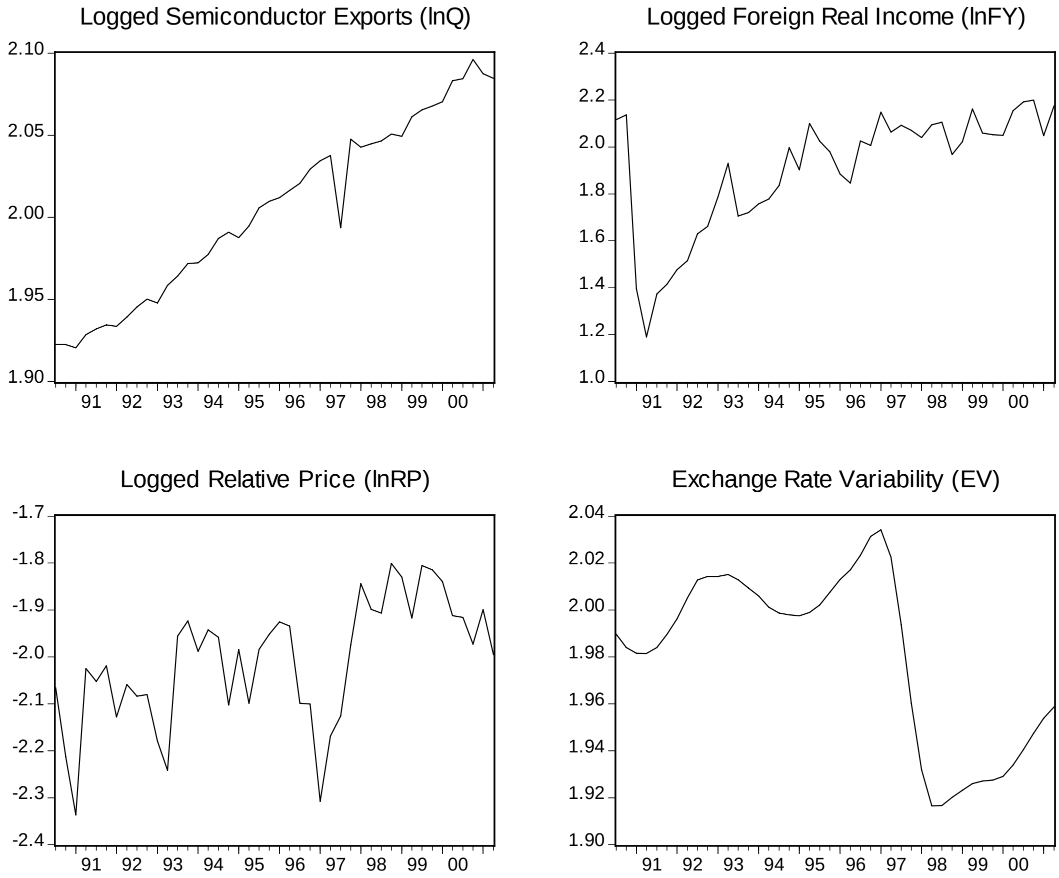
International Monetary Fund's (IMF) *International Financial Statistics*. The definition and transformation of each variable is briefly explained in Appendix A.

Econometric Methods and Procedures

The initial step in time series modeling is to identify the degree of integration of the variables in interest. Most of the macroeconomic series are found to have unit roots i.e. they are not stationary or their variances increase with time (Nelson and Plosser, 1992; Arize, 1996a and 1996b). If unit roots are presented in each time-series variable, 'spurious' (invalid) regression may arise if we regress these time-series variables (in levels) that contain a trend component using Ordinary Least Squared (OLS) estimator (see Engle and Granger, 1987). Hence, it is necessary to pre-test each variable of interest for unit roots before the OLS regression is estimated. The conventional unit root tests can be carried out by using Augmented Dickey-Fuller (ADF), Phillips and Perron (PP), and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests. However, it is important to note that these specifications assume no structural break(s).

Figure 1 illustrates the time-series plots of logged variables in levels. A virtual inspection on the plots suggests that the logged semiconductor exports, logged relative price, logged foreign real income, and exchange rate variability have structural break(s) in the third quarter of 1997, second quarter of 1997, second quarter of 1991 and first quarter of 1997. respectively. In this context, as warned by Perron (1989) that the conventional unit root tests were not appropriate for variables that have undergone structural changes, and the power to reject the unit root null declines if the data contains a structural break that is ignored. In this relation, a specification takes into account a structural break is necessary in empirical analysis whether the demand for semiconductor exports and the key determinants are nonstationary, or in $I(1)$ processes.

Figure 1: Time-series plots of logged export quantity, logged relative price, logged foreign real income and exchange rate variability in levels.



Therefore, this study applies the recently developed unit root tests with an unknown level shift which are proposed by Lanne *et al.* (2002), and Saikkonen, and Lütkepohl (2002) in order to examine whether the variables of export demand function are nonstationary. The proposed equation for unit root tests is $y_t = \mu_0 + \mu_1 t + f_t(\theta)'\gamma + x_t$ which is based on estimating the deterministic term first by a Generalised Least Squares (GLS) procedure under the unit root null hypothesis and subtracting it from the original series. Following this, an ADF type test is applied on the adjusted series which also includes terms to correct for estimation errors in the parameters of the deterministic part. As in the case of the ADF statistic, the asymptotic null distribution is non-standard. The relevant critical values are tabulated in Lanne *et al*. (2002). Since the break date is unknown, Lanne *et al.* (2002) recommended choosing a reasonably large autoregressive order (AR) in the first step then selecting the break date which minimizes the GLS objective function used to estimate the parameters of the deterministic part. A shift function, which is here denoted by $f_t(\theta)'\gamma$, may be added to the deterministic term μ_t of the data generation process. Hence, a model $y_t = \mu_0 + \mu_1 t + f_t(\theta)'\gamma + x_t$ is considered where θ and γ are unknown parameters or parameter vectors and the errors x_t are generated by an AR(*p*) process with a possible unit root. In general, three possible shift functions can be implemented which are:

1. A simple shift dummy variable with shift date T_B, $f_t^{(1)} = d_{1t} := \begin{cases} 0, t < T_B \\ 1, t \geq T_B \end{cases}$. This function does not involve an extra parameter θ. In the shift term $f_t^{(1)}\gamma$, the parameter γ is a scalar. Differencing this shift function leads to an impulse dummy.
2. The second shift function is based on the exponential distribution function which allows for a nonlinear gradual shift to a new level starting at time T_B, $f_t^{(2)}(\theta) = \begin{cases} 0, t < T_B \\ 1 - \exp\{-\theta(t - T_B + 1)\}, t \geq T_B \end{cases}$. In the shift term $f_t^{(2)}(\theta)\gamma$, both θ and γ are scalar parameters. The first scalar parameter is confined to the positive real line (θ>0), whereas the second scalar parameter may assume any value.
3. The third shift function can be expressed as a rational function in the lag operator applied to a shift dummy d_{1t}, $f_t^{(3)}(\theta) = \left[\frac{d_{1,t}}{1-\theta L} : \frac{d_{1,t-1}}{1-\theta L} \right]'$. The actual shift term is $\left[\gamma_1 (1-\theta L)^{-1} + \gamma_2 (1-\theta L)^{-1} L\right] d_{1t}$, where θ is a scalar parameter between 0 and 1 and $\gamma = (\gamma_1 : \gamma_2)'$ is a two-dimensional parameter vector. Note here that both $f_t^{(2)}(\theta)\gamma$ and $f_t^{(3)}(\theta)'\gamma$ can generate sharp one-time shifts at time T_B for suitable values of θ. Thus $f_t^{(2)}(\theta)\gamma$ and $f_t^{(3)}(\theta)'\gamma$ are more general than $f_t^{(1)}\gamma$.

However, as based on Monte Carlo simulations, Lanne *et al.* (2002, p.682) found that the performance of the tests tends to be inferior if one of the more complicated shift functions $f_t^{(2)}$ or $f_t^{(3)}$ is employed. Where the tests give different results, the findings from $f_t^{(1)}$ are preferable. Thus, this paper only applies the findings from $f_t^{(1)}$.

Once the variables of interest are found to be nonstationary, or *I*(1), cointegration tests can be performed to investigate whether a long run relationship among the variables exits (i.e. stable long-run equilibrium among the variables in equation (1)). To do so, the Johansen trace test can be applied to test for the presence of a cointegrating vector or vectors among the non-stationary series as suggested by Johansen (1988, 1991) and Johansen and Juselius (1990, 1992, 1994). The assumption imposed on the cointegration equations is linear deterministic trend and intercept in data without structural break(s). However, as indicated by the time-series plots in Figure 1, all the logged variables have structural breaks in levels. Hence, the specification of Johansen test involves deterministic trend and intercept with structural break date(s) becomes more appropriate. In fact, this modeling option can be done in the econometric software package JMulTi[‡] and the p-values are based on Johansen *et al.*'s (2000) critical values which take into account of break date(s).

If the economic time series are found to be cointegrated, an econometric framework for an ECM representation can be specified. The error-correction process can reconcile the long-run equilibrium with disequilibrium behavior in the short-run. The advantage of using this modeling strategy is that it minimizes the least squares errors and increases the convergence speed of the estimates to its true values (Engle and Granger, 1987). Thus, the ECM specification of equation (1) can be written as follows:

$$\Delta ln\mathrm{Q}_t = \mathrm{a} + \sum_{j=o}^{p} b_j \Delta ln\mathrm{RP}_{t-j} + \sum_{j=o}^{p} c_j \Delta ln\mathrm{FY}_{t-j} + \sum_{j=o}^{p} d_j \Delta \mathrm{EV}_{t-j} + \sum_{j=1}^{p} e_j \Delta ln\mathrm{Q}_{t-j}$$

$$-\lambda \mathrm{EC}_{t-1} + \varepsilon_t \qquad (2)$$

where Δ is first-order differencing operator (e.g. $ln\mathrm{Q}_t - ln\mathrm{Q}_{t-1}$) and EC_{t-1} stands for the previous period's error correction term generated from a cointegrating equation using OLS estimator.[§]

Given that this paper has only 48 observations and to save the degrees of freedom, a maximum lag length of 4 will be imposed on equation (2). As a general rule, an optimal lag length of 4 quarters is sufficient in empirical study when quarterly data are being used. Then we can narrow the general model down by looking for simplification that is acceptable to the data on the basis of "general to specific" modeling paradigm using individual *t*-test. In selecting the specific model, all those variables that have relatively small absolute *t*-value (less than one) were dropped sequentially. The rationale of this model specification search strategy is that empirical researchers do not know the actual data generating process and

[‡] The software is available from http://www.jmulti.de/

[§] According to Abeysinghe and Tan (1999), in small samples OLS may still be the best choice among the six estimation techniques viz. OLS, unrestricted error correction model or autoregressive distributed lag model (ARDL), a fully modified least square, 3-step estimator, OLS regression augmented by leads and lags of the differenced explanatory variables, and Johansen's estimator.

hence, are not possible to deduce in advance what model specification actually represents the data generation process (Davidson *et al*. 1978 and Hendry 1980).

III. EMPIRICAL RESULTS

Before the long run relation regression is estimated, each individual series in levels is tested for unit roots using unit root tests for processes with unknown level shift developed by Lanne *et al.* (2002), and Saikkonen, and Lütkepohl (2002). Table 1 reports that the test statistics do not reject the null hypothesis of unit root for all variables included in the export demand function which suggest that they are integrated of order one, or in *I*(1) processes. However, the result of KPSS suggests that EV is stationary in levels when the test statistic is 0.145 with the 5 percent critical value of 0.146, the null hypothesis of trend stationary is accepted. This strongly supports previous empirical work that most economic time series have a unit root.

Table 1: Results of unit root tests for processes with level shifts, $f_t^{(1)}$

Regressor:		Test statistic	Suggested break date[1]
$ln\text{Q}_t$		-1.9826	1997q3
$ln\text{RP}_t$		-3.228[2]	1997q1
$ln\text{FY}_t$		-2.1432	1997q1
EV_t		-2.3452	1997q2
Critical values (T=50)	1%	-3.81	
	5%	-3.15	
	10%	-2.86	

Notes: The critical values are from Lanne *et al*. (2002). Shift function – impulse dummy; time tend and seasonal dummies included. The lag order is suggested by AIC (Akaike Information Criterion). The null hypothesis is a unit root with unknown with level shift(s).

[1] As Lanne *et al.* (2003) recommend choosing a reasonably large AR order as a first step, and then picking the break date which minimizes the GLS objective function used to estimate the parameters of the deterministic part.

[2] The test statistic is based on a unit root equation without trend. The 10% critical value for this specification is -2.67 (T=50). The test statistic, which is based on a unit root equation with trend (and seasonal dummies) is -3.6459, rejects the null hypothesis at 5%.

The next step is to apply the Johansen multivariate cointegration procedure to test whether there is a cointegrating vector or vectors among the non-stationary series. As a rule of thumb, a four quarters lag length is considered for quarterly data. Table 2 and Table 3 report the estimated trace test statistics without and with break dates (without and with exchange rate variability variable), respectively. Overall, the cointegration test results from both tables confirm that there exists at least one cointegrating relationship among the four variables, export demand for semiconductors (at SITC level), relative price, foreign real income, and exchange rate variability. However, in practice, even if the *F*-test rejects cointegration, we still cannot draw any conclusion because a more powerful test for cointegration in this set-up is the coefficient obtained for lagged error correction term (see Kremers *et al.* 1992). As noted by Bahmani-Oskooee and Brooks (1999), if the lagged error-correction term turns out to be negative and significant, cointegration is supported.

Table 2: Results of Johansen Trace test without break date (p-value)

lnQ_t, $lnRP_t$, $lnFY_t$	r0 = 0	1	2	3
Intercept	0.000	0.054	0.054	
Trend and intercept	0.004	**0.243**	**0.201**	
lnQ_t, $lnRP_t$, $lnFY_t$, EV_t	r0 = 0	1	2	3
Intercept	0.000	0.000	0.002	0.035
Trend and intercept	0.000	0.007	0.051	**0.163**

Note: 4 lags were included since the data are quarterly data.

Table 3: Results of Johansen Trace test with break date (p-value)

lnQ_t, $lnRP_t$, $lnFY_t$	r0 = 0	1	2	3
Intercept				
Break date: 1997q1	0.000	0.000	0.088	
1997q2	0.000	0.000	0.012	
1997q3	0.000	0.000	0.006	
Trend and intercept				
Break date: 1997q1	0.000	0.006	**0.173**	
1997q2	0.000	0.001	0.019	
1997q3	0.000	0.000	0.014	
lnQ_t, $lnRP_t$, $lnFY_t$, EV_t	r0 = 0	1	2	3
Intercept				
Break date:- 1997q1	0.000	0.000	0.0003	**0.1103**
1997q2	0.000	0.000	0.0006	0.0751
1997q3	0.000	0.000	0.0003	0.0924
Trend and intercept				
Break date:- 1997q1	0.000	0.000	0.0008	0.0212
1997q2	0.000	0.000	0.0011	0.0146
1997q3	0.000	0.000	0.0008	0.0276

Note: The break dates are based on the dates suggested in the unit root tests. The p-values are based on Johansen, *et al.*'s (2000) critical values which taking into account of break date(s).

The estimated long-run elasticities are reported in Table 4. Both the estimated long-run elasticities of relative price and foreign real income have expected sign i.e. negative and positive, respectively while the exchange rate variability is positively related to the demand for exports. First of all, the long-run relative price elasticity is less than one in absolute terms suggesting that the demand for Malaysia's exports of semiconductors is inelastic to price changes. In addition, the estimated parameter for long-run foreign income is inelastic with a magnitude of 0.114 implying that the semiconductor exports are not responsive to changes in foreign real income during the sample period considered. This is plausible for semiconductor exports, and it can be intuitively linked to (1) intensified efforts to deal with Year 2000 (Y2K) problem; (2) the more widespread usage of internet, e-commerce, cellular phones, and telecommunications by foreign business firms and consumers; (3) an increased investments in technology; and (4) the need to upgrade technology for competitive reasons. On the other hand, the estimated long-run elasticity for exchange rate variability is found to be statistically significant and has a magnitude of 0.16 (inelastic), implying that the variability of exchange rate has some positive effect on Malaysia's semiconductor exports in the long-run. This evidence corroborates De Grauwe's (1988) view that semiconductor exporting firms, which are mostly MNCs, are very risk-averse and have a tendency to raise exports in response to an increase in exchange rate uncertainty in the long -run because they worry about the decline in export earnings. Bahmani-Oskooee and Ltaifa (1992) found that within the developing countries, those who fixed their exchange rates to one major currency were found to be subject to less risk than the other developing countries.

It is worth noting that when the exchange rate variability is omitted from the long-run regression lnQ_t (1), the dummy variable which is used to capture the possible structural break is not significantly different from zero. Moreover, when it is included in the estimation such as equation lnQ_t (2), the dummy variable is statistically different from zero explaining the suggested break dates could be linked to the sharp depreciation of the Ringgit exchange rate before the Asian currency crisis i.e. 1997q1, 1997q2, and 1997q3. The structural breaks have a significant positive impact on exports. Its elasticity magnitude of 0.039 suggests that the Ringgit's slide against the US dollar and most of other major currencies before the crisis period have been quite instrumental in raising the external demand for Malaysia's semiconductors despite the uncertain outlook in the currency markets.

Table 4: Ordinary least squares estimation of long-run elasticity parameters

Regressor:	$ln\mathrm{Q}_t$ (1)	$ln\mathrm{Q}_t$ (2)
RP_t	0.208***	0.162***
	(5.258)	(3.477)
FY_t	0.124***	0.114***
	(6.534)	(5.887)
EV_t		-0.287*[1]
		(-1.75)
Dummy	0.035*	0.039*
	(1.749)	(1.973)
Constant	2.184***	2.679***
	(22.275)	(8.971)
R-squared	0.737	0.756
Durbin-Watson statistic	0.807	0.734
F-statistic [p-value]	37.39 (0.000)	30.259 [0.000]

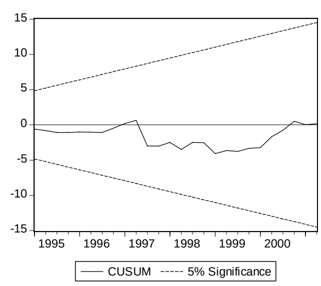
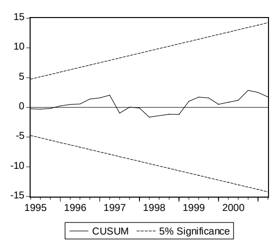
Note: *, **, and *** denote significant at 10%, 5% and 1%, respectively; (.) is t-statistic; [1] the estimated parameter of *ln*RP is -0.251 (t-ratio is -1.492) without dummy variable; Dummy variable takes on the value 1 from 1997:1 – 1997:3, and 0 for other quarters; (1) denotes estimation without EV_t and (2) refers to estimation with EV_t.

Table 5 provides the estimation results of the ECM with and without exchange rate variability. The diagnostic test statistics do not reject the null hypotheses of the OLS assumptions relating to error processes such as white noise and normality (except for *ln*Q(1)). In addition, both models are free from ARCH effect. The Ramsey's regression specification error tests (RESET) do not detect any misspecification in the estimated regressions $ln\mathrm{Q}_t$ (1) and $ln\mathrm{Q}_t$ (2) while the plots of CUSUM and CUSUMSQ tests suggest both the regressions are stable at 5 per cent significance level.** However, the negative EC term's coefficient is only statistically significant in regression $ln\mathrm{Q}_t$ (2) i.e. when the variability of exchange rate variable is included in the estimation. This evidence advocates the results of Johansen's multivariate cointegration tests that at least one cointegrating vector for both models (with and without exchange rate variability) that an equilibrium relationship is from $ln\mathrm{Q}_t - \beta_0 - \beta_1 ln\mathrm{RP}_t - \beta_2 ln\mathrm{FY}_t - \beta_3\mathrm{EV}_t$ rather than $ln\mathrm{Q}_t - a_0 - a_1 ln\mathrm{RP}_t - a_2 ln\mathrm{FY}_t$. Furthermore, the estimation results

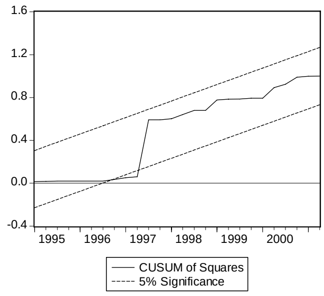
** However, this is not a serious concern since the instability is observed in 1996q2-1997q2 and the CUSUM test rejects instability.

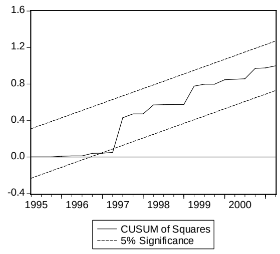
from Table 5 also show the relative price, foreign real income and the variability of exchange rate have short-run effects on semiconductor exports. However, there is not much interpretation that could be attached to the short-run coefficients; perhaps, all they show are the dynamic adjustment of all variables (see Bahmani-Oskooee and Rehman, 2005, p. 775).

Table 5: Regression results for ECM

Regressor:	$\mathit{ln}Q_t$ (1)	$\mathit{ln}Q_t$ (2)
EC_{t-1}	-0.128 (-1.392)	-0.284 (-3.119)**
$\Delta \mathit{ln}Q_{776\ t-1}$	-1.136 (-5.59)***	-1.089 (-6.175)***
$\Delta \mathit{ln}Q_{776\ t-2}$	-1.31 (-4.75)***	-1.331 (-6.0145)***
$\Delta \mathit{ln}Q_{776\ t-3}$	-0.955 (-3.719)***	-0.927 (-4.253)***
$\Delta \mathit{ln}Q_{776\ t-4}$	-0.36 (-1.90)*	-0.316 (-2.076)**
$\Delta \mathit{ln}RP_t$	-0.099 (-2.56)**	0.603(3.663)***
$\Delta \mathit{ln}RP_{t-1}$	-0.038 (-2.16)**	
$\Delta \mathit{ln}RP_{t-2}$	0.053 (1.52)	
$\Delta \mathit{ln}RP_{t-3}$		-1.196 (-4.16)***
$\Delta \mathit{ln}RP_{t-4}$	0.032 (2.168)**	1.172 (4.214)***
$\Delta \mathit{ln}YF_t$	-0.086 (-2.327)**	-0.0554 (-1.915)*
$\Delta \mathit{ln}YF_{t-1}$		-0.0434 (-1.497)
$\Delta \mathit{ln}YF_{t-2}$	0.053 (1.46)	
$\Delta \mathit{ln}YF_{t-3}$		
$\Delta \mathit{ln}YF_{t-4}$	0.0197 (1.676)	
ΔEV_t		-0.065 (-2.066)**
ΔEV_{t-1}		-0.070 (-2.1) **
ΔEV_{t-3}		-0.0244 (-2.084)**
Constant		0.023 (6.658)***
R-squared	0.691	0.801
Durbin-Watson stat	2.266	2.233
F-stat [p-value]	4.84 (0.000)	7.763 [0.000]
Jarque-Bera [p-value]	30.05 [0.000]	0.740 [0.691]
ARCH: F-stat [p-value]	0.271 [0.764] (2 lags)	0.950 [0.397] (2 lags)
LM test: F-stat [p-value]	0.766 [0.476] (2 lags)	0.604[0.555] (2 lags)
Ramsey's RESET: F-stat [p-value]	0.564 [0.46] (1 lag)	3.69 [0.067] (1 lag)
CUSUM		

CUSUMSQ





Note: *, **, and *** denote significance at 10%, 5% and 1%, respectively; (.) is *t*-statistic; (1) denotes estimation without EV_t and (2) refers to estimation with EV_t.

IV. CONCLUSIONS

In this paper, we have found that all the variables in the export demand equation i.e. quantities of export, relative price, real foreign income, and real exchange rate variability have structural breaks based on the time-series plot of each variable. The results of the unit root tests proposed by Lanne *et al.* (2002), and Saikkonen, and Lütkepohl (2002), which take into account of structural break, suggest break dates of 1997q1, 1997q2 and 1997q3, confirm all variables are nonstationary ($I(1)$). The application of Johansen trace tests, which also take into account of structural break, suggests that there is at least a cointegrating relationship among the variables. The validity of this cointegrating relationship is supported by the negative statistical significance of the error-correction term in ECM.

A brief discussion is made on the policy implications which are based on the findings of this paper. The estimated elasticity for long-run relative price elasticity is less than unity (in absolute terms), implying the competition in semiconductor exports is imperfect i.e. price is a less sensitive factor to determine the level of semiconductor exports. This empirical evidence implies that Malaysia's semiconductors exporters can adopt various forms of non-price competition to enhance their exports' growth performance. For instance, they should consider upgrading their production activities towards more technological advanced products in order to have a non-price competitive edge on their export competitors in the global markets. The semiconductor industry is continually evolving because of ongoing R&D activities that can lead to new and improved products. Therefore, the core activities of semiconductor manufacturing such as silicon ingot growing, cutting and polishing of silicon wafers, chip design and wafer fabrication become important non-price competitive activities. Despite the industry is not fully equipped with the necessary technical skills to produce the latest generation of integrated circuit (IC) and the state-of-the-art product design and development, the semiconductor firms should invest in the production of high value-added products or activities in order to sustain the growth momentum and competitiveness in the industry. Currently, Malaysian government is attempting to promote these activities in order have non-price competitive advantage in the manufacturing of semiconductors. To support the efforts of semiconductor firms as well as policies and strategies to shift towards the production of higher value-added semiconductor devices and activities, the government should continue to upgrade the physical and technology infrastructure, and develop programs focusing on critical skills needed by the industry for manpower training.

With reference to inelastic long-run foreign income elasticity of export demand for semiconductors, it implies that Malaysia's semiconductor exports have a low degree of exposure to its major export markets such as the U.S. and Japan because the industry is dominated by wholly owned subsidiaries of foreign corporations mainly from U.S. and Japan, which provide both backward and forward linkages for global electronics markets. For example, a backward linkage is achieved when inputs are being imported from abroad or home countries of MNCs for value added in Malaysia while forward linkage is achieved when semifinished products e.g. semiconductor devices are being exported back to their home countries or affiliates elsewhere for assembly or distribution (Sieh-Lee, 2000).

The findings also show that the long-run exchange rate variability elasticity is positive and is statistically significant. It suggests that exchange rate variability has a positive influence on semiconductor exports in the long -run. This in no doubt supports the theoretical argument developed by De Grauwe (1988) that a very risk-averse exporter, who is concerned with depressing effect on export earnings, may export more when the exchange rate uncertainty increases. By the same token, as proposed by Doraisami (2004), exchange rate monitoring and export diversification could be used to enhance the performance of the Malaysian exports sector.

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Appendix A. Variable Definition

(1) Q_t = Quantity of semiconductor exports based on SITC 776 product group.

(2) RP_t = $\dfrac{P_{SITC776}}{PW_t}$

where P_{SITC} is the unit price of semiconductor exports, which is constructed based on the followings:

$$\text{Unit price of SITC 776 exports} = \frac{\text{Current FOB value (RM) of SITC 776 exports}}{\text{Quantity of SITC 776 exports}}$$

where FOB is the abbreviation for free on board, and the export quantity of SITC 776 product group is measured in common unit.

$$PW_t = \sum_{n=1}^{3} w^n P^n_t$$

where $w^n = x^n / \sum_{n=1}^{3} x^n$, trade share of Malaysia's n^{th} major trading partners i.e. U.S. Japan and Singapore, x^n = Malaysia's electrical exports to the n^{th} trading partners, and P^n_t = wholesale price index of Malaysia's n^{th} major trading partners.

(3) FY_t = $\sum_{n=1}^{3} w^n FY^n_t$

where FY^n_t = real gross domestic product of Malaysia's n^{th} trading partner, and the weights used for constructing foreign income variables are similar to those used in the construction of the relative price variables.

(4) EV_{t+m} = $[\frac{1}{m}\sum_{i=1}^{m} (REER_{t+i-1} - REER_{t+i-2})^2]^{½}$

where REER = index of real effective exchange rate (1995=100), and m = 4, which is the order of the moving average.