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# Structural change in the export demand function for Indonesia: Estimation, analysis and policy implications

Akhand Akhtar Hossain *

*School of Economics, Politics and Tourism, The University of Newcastle, Newcastle, NSW 2308, Australia*

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**Abstract**

This paper investigates the aggregate export demand behaviour in Indonesia with annual data for the period 1963–2005. Both the Pesaran bounds testing and the Johansen cointegration tests results suggest that there exists a long-run relationship between real exports, world income and the relative export prices in Indonesia. The long-run income elasticity of the demand for Indonesia's exports is significantly greater than one and the long-run relative export price elasticity of the demand for its exports is significantly lower than one. The recursive and rolling regressions and the Hansen–Johansen stability test results suggest that the export demand function for Indonesia has undergone a significant structural change since the late-1990s, which is reflected in the decrease of the income elasticity, and an increase in the relative export price elasticity, of demand for Indonesian exports. These results explain the relatively slow growth of Indonesian exports since 2000 that has slowed the country's economic recovery from the financial crises of 1997–1998.


*JEL classification:* C32; F11

*Keywords:* Export demand; Cointegration; Bounds testing; Indonesia

## 1. Introduction

Indonesia has achieved significant economic and social development since the late-1960s by adopting an export-oriented development strategy.[1] The military government under General Soeharto provided the much needed economic and political stability during 1966–1996. During this

* Tel.: +61 2 49218842; fax: +61 2 49216911.
*E-mail address:* akhtar.hossain@newcastle.edu.au.

[1] For all intents and purposes the export-oriented industrialisation in Indonesia started in the aftermath of the decline in oil prices in the mid-1980s (Aswicahyono & Pagestu, 2000; Hossain, 2006).

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period the traditional agricultural, albeit natural resource-rich, economy of Indonesia underwent an impressive transformation into a balanced, outward-oriented industrialising economy. The 1997–1998 currency-cum-financial crises in Indonesia were unexpected given its strong economic fundamentals (IMF, 2007). Yet, unlike South Korea, Malaysia and Thailand, the Indonesian crises lasted longer. However, contrary to the prediction of a 'backlash protectionism' after the crises, Indonesia has maintained its outward-oriented development strategy. This has been appreciated by international investors and financiers. Indonesia has slowly recovered and moved to a higher growth path since 2004. Having improved economic, social and political institutions, macroeconomic policies have lately been geared successfully towards price stability while supporting high growth (IMF, 2007). This is attributed to an improvement in the international trade environment. Indonesia encountered difficulties in raising its export growth to the level that might have pushed its economic growth rate to the pre-crisis path (IMF, 2004; Nasution, 2002; Soesastro & Basri, 2005). IMF (2007, p. 6) has estimated that Indonesia in fact lost market share from 0.97% of total trade in 2000 to 0.95% in 2006.[2]

The main objective of this paper is to examine the possibility of a structural change in the export demand behaviour in Indonesia. Following the trade literature, a conventional export demand equation is specified and both the Pesaran bounds testing[3] and the Johansen cointegration tests are conducted to establish a long-run relationship between real exports, world income and the relative export prices.[4] The recursive and rolling regressions and the Hansen–Johansen stability test are then conducted to investigate whether the export demand function in Indonesia has undergone a structural change. The paper draws policy implications from empirical findings.

## 2. Model specification

Following the trade literature,[5] Indonesia's exports are assumed to be demand determined. Accordingly, real exports are an increasing function of world income and a decreasing function of relative export prices.[6] For estimation purposes, the aggregate export demand function in Indonesia can then be specified in a log-linear form, such that

$$\ln \mathrm{REX}_t = \beta_0 + \beta_1 \ln Y_t^{\mathrm{w}} + \beta_2 \ln 100 \left( \frac{\mathrm{EPI^i}}{\mathrm{EPI^a}} \right)_t + U_t \tag{1}$$

[2] Indonesia has lost market shares in both non-oil exports and manufactured products.

[3] Pesaran and Shin (1996), Pesaran and Pesaran (1997), Pesaran and Smith (1998), and Pesaran, Shin, and Smith (2001) have developed the autoregressive distributed-lag (ARDL) bounds testing approach to cointegration. This test remains valid irrespective of whether the regressors are purely $I(0)$, purely $I(1)$ or mutually cointegrated.

[4] The study uses annual data for the period 1963–2005. All data are drawn from various issues of the *International Financial Statistics* of the IMF.

[5] The trade literature is voluminous. The review and recent empirical studies include Goldstein and Khan (1985), Arize (1990, 2001), Roy (1991), Muscatelli and Stevenson (1995), Bahmani-Oskooee (1986), Bahmani-Oskooee and Niroomand (1998), Masih and Mashi (2000), Hamori and Matsubayashi (2001), and Singh (2002).

[6] An implicit assumption behind this demand function is that supply is not a constraint on exports. Indonesia has the productive capacity to increase the supply of most of its exportable products as their demand increases. Given that Indonesia is a labour-surplus country and is endowed with natural resources, this assumption is realistic at the early stage of its industrialisation.

where REX is Indonesia's volume of exports,[7] $Y^w$ the world real GDP, $EPI^a$ the export unit value index in Asia,[8] and $U$ is a random error term with zero mean and a constant variance. In this model an increase in global income increases the demand for Indonesian products. Depending on the value of $\beta_1$, Indonesia's exports can be income-elastic ($\beta_1 > 1$) or income-inelastic ($\beta_1 < 1$). Similarly, an increase in the export price in Indonesia relative to that in Indonesia's competitors decreases the demand for its products. Depending on the absolute value of $\beta_2$, Indonesia's exports can then be price-elastic ($|\beta_2| > 1$) or price-inelastic ($|\beta_2| < 1$.

## 3. Testing for the long-run export demand relationship in Indonesia

This section conducts both the Pesaran bounds testing and the Johansen cointegration tests to establish any long-run relationship between real exports, world income and the relative export prices in Indonesia.

### 3.1. The ARDL bounds testing approach to cointegration

The error-correction form of the ARDL model in the variables real exports (REX), world income ($Y^w$) and the relative export prices (REP) is specified as follows:

$$\begin{aligned}\Delta \ln REX_t = {} & \alpha_0 + \alpha_1 T + \sum \beta_i \Delta \ln REX_{t-i} + \sum \gamma_i \Delta \ln Y^w_{t-i} \\ & + \sum \phi_i \Delta \ln REP_{t-i} + \delta_1 \ln REX_{t-1} \\ & + \delta_2 \ln Y^w_{t-1} + \delta_2 \ln REp_{t-1} + U_t \end{aligned} \tag{2}$$

where the coefficients $\beta_i$, $\gamma_i$ and $\phi_i$ represent the short-run dynamics of the underlying variables in the ARDL model and the coefficients $\delta_s$ represent the long-run relationship. This specification is based on the maintained hypothesis that the time series properties in the export demand relationship can be approximated by a log-linear VAR(p) model, augmented with deterministic intercepts and (probably) trends (T). Although in the specification the value of $i$ can be infinity, the model is estimated sequentially with one to the maximum lag 3.[9]

#### 3.1.1. Testing for the hypothesis that $\delta_1 = \delta_2 = \delta_3 = 0$

Eq. (2) is estimated first in a restricted form by excluding the level form lag variables and then is tested for the significance of the lagged level variables through a variable addition test ($F$-test). The estimated $F$-statistic for the restriction that $\delta_1 = \delta_2 = \delta_3 = 0$ in the specification with (log) real exports (LREX) as dependent variable is denoted by $F(\text{LREX}|\text{LY}^w;\text{LREP})$, where $\text{LY}^w$ is the log of world income and LREP is the log of relative export prices. This process is repeated for the specification with LREP or $\text{LY}^w$ as dependent variable. The estimated $F$-statistic for the restriction that $\delta_1 = \delta_2 = \delta_3 = 0$ in the latter specifications are denoted by $F(\text{LY}^w|\text{LREX};\text{LREP})$ or $F(\text{LREP}|\text{LREX};\text{LY}^w)$. The estimated $F$-statistics are compared with critical values to determine

[7] The export volume is defined as Indonesia's total export earnings in US$ million deflated by the export unit value index (based on US$) in Indonesia $EPI^i$.

[8] $EPI^a$ is based on US$. In the text REP is used to represent the expression: $100 \times EPI^i/EPI^a$.

[9] Given the annual data, three lag terms are considered adequate. Pesaran and Pesaran (1997) suggest that a lag length of one period can be a reasonable choice in case of annual data.

Table 1
Testing for the long-run relationship between LREX, $LY^w$ and LREP ($F$-test)

| Lags | F-statistics | | | | | |
|---|---|---|---|---|---|---|
| | $F$(LREX\|$LY^w$;LREP) | | $F$($LY^w$\|LREX;LREP) | | $F$(LREP\|LREX;$LY^w$) | |
| | Full sample: 1965/1967–2005 | Pre-crisis: 1965/1967–1996 | Full sample: 1965/1967–2005 | Pre-crisis: 1965/1967–1996 | Full sample: 1965/1967–2005 | Pre-crisis: 1965/1967–1996 |
| With constant | | | | | | |
| 1 | 2.25 | 2.52 | 1.74 | 1.73 | 4.96 | 0.3 |
| 2 | 1.6 | 2.18 | 3.25 | 2.78 | 4.43 | 0.51 |
| 3 | 1.17 | 2.05 | 2.99 | 2.13 | 2.48 | 0.19 |
| With constant and trend | | | | | | |
| 1 | 2.34 | 2.68 | 1.75 | 1.15 | 5.81 | 4.96 |
| 2 | 1.58 | 2.03 | 2.42 | 1.58 | 6.72 | 4.43 |
| 3 | 1.29 | 1.66 | 2.69 | 1.86 | 5 | 2.48 |

*Note*: The critical value bounds of the $F$-statistic with constant are {3.17–4.14} and {3.79–4.85} at the 90% and 95% respectively and those with constant and trend are {4.19–5.06} and {4.87–5.85} at the 90% and 95%, respectively.

whether there exists a long-run relationship between real exports, world income and the relative export prices. In addition, the estimated $F$-statistics provide information on whether any of these variables can be considered a long-run forcing variable in determining others.

Table 1 reports the $F$-statistics with one up to three lags in the specification. The model is estimated for two cases: with both an intercept and trend and only intercept. The results are sensitive to the inclusion of trend in the specification for real exports as dependent variable. In the specification with both unrestricted intercept and unrestricted trend, the critical value band for $k=2$ is {4.19–5.06} and {4.87–5.85} at the 90% and 95%, respectively. In the specification with intercept only, the critical value band for $k=2$ is {3.17–4.14} and {3.79–4.85} at the 90% and 95%, respectively. Since $F$(LREX|$LY^w$;LREP) is below the upper limit of the critical band irrespective of the order of lag, the null hypothesis of no long-run relationship between LREX, $LY^w$ and LREP cannot be rejected. However in the specification LREP as dependent variable (with trend), the statistic F(LREP|LREX;$LY^w$) exceeds the upper bound of the band and therefore the null hypothesis of no long-run relationship between LREP, LREX and $LY^w$ is rejected. These results suggest that LREP cannot be treated as a 'long-run' forcing variable for the explanation of LREX.

### 3.1.2. *Testing for the hypothesis that $\delta_1 = 0$*

In addition to the Wald test for the joint hypothesis, the presence of a long-run relationship can be examined via a $t$-test using the specified error-correction model (2). This can be done by testing for the significance of the coefficient on $LREX_{t-1}$.[10] Pesaran et al. (2001) have provided the lower and upper bound critical values for this statistic. Table 2 reports the test results for both the full and pre-crisis sample periods. The results remain sensitive to the order of lag but weakly support the presence of a long-run relationship between LREX, $LY^w$ and LREP.

[10] Banerjee, Dolado, and Mestre (1998) also suggest that the presence of a long-run relationship can be examined via the $t$-statistic on the coefficient on $LREX_{t-1}$.

Table 2
Testing for the long-run relationship between LREX, $LY^w$ and LREP ($t$-test)

| Lags | Coefficient on $LREX_{t-1}$ ($t$-statistics are in parentheses) | |
|---|---|---|
| | Full sample: 1965/1967–2005 | Pre-crisis: 1965/1967–1996 |
| With constant | | |
| 1 | −0.36 (−2.48) | −0.38 (−2.71) |
| 2 | −0.36 (−2.12) | −0.41 (−2.39) |
| 3 | −0.38 (−1.80) | −0.50 (−2.21) |
| With constant and trend | | |
| 1 | −0.36 (−2.68) | −0.42 (−2.75) |
| 2 | −0.37 (−2.17) | −0.44 (−2.29) |
| 3 | −0.41 (−1.86) | −0.52 (−2.18) |

*Note*: The critical value bounds of the $t$-statistic with constant are {−2.57 to −3.21} and {−2.86 to −3.53} at the 90% and 95% respectively and those with constant and trend are {−3.13 to −3.63} and {−3.41 to −3.95} at the 90% and 95%, respectively (Pesaran et al., 2001).

Table 3
Long-run coefficients on world income and relative export prices

| Regressors | Model: ARDL (2,0,1) (full sample: 1965–2005) coefficient ($t$-ratio) | Model: ARDL (2,0,1) (shorter sample: 1965–2000) coefficient ($t$-ratio) |
|---|---|---|
| Intercept | 3.48 (4.70) | 2.77 (6.73) |
| $LY^w$ | 1.86 (10.68) | 2.11 (17.22) |
| LREP | −0.22 (−1.38) | −0.31 (−3.17) |
| Wald test | | |
| $H_O$: coefficient on $LY^w = 1$ | $\chi^2_{(1)} = 24.35$ | $\chi^2_{(1)} = 82.25$ |
| $H_{O:}$ coefficient on LREP = 0 | $\chi^2_{(1)} = 1.90$ | $\chi^2_{(1)} = 10.04$ |

### *3.1.3. Estimating the coefficients of the long-run relationship*

The second stage of the ARDL modeling involves estimating the coefficients of the long-run relations and making inference about their values. In general, in estimating the long-run coefficients, the ARDL technique estimates $(p+1)^k$ number of regressions in order to obtain the optimal lag length for each variable, where $p$ is the maximum number of lags and $k$ is the number of variables in the equation. This paper uses the Schwartz–Bayesian criterion (SBC) to select the optimal order of lag.

Table 3 reports the estimated long-run coefficients on world income and the relative export prices with real exports as dependent variable. The export equation has been estimated with the order of lag two in the variables and a trend. The model is also estimated for two sample periods: 1965–2000 and 1965–2005. The estimated model for the shorter sample period 1965–2000 is used for ex post forecasting.[11]

In the estimated model, the coefficients on world income and the relative export prices bear their expected signs and are significant. The Wald test rejects the proposition of unit elasticity on

[11] Forecasting results are not reported but are available from the author upon request.

Table 4
The Johansen cointegration test results

| Null | Alternative | Maximum eigenvalue statistic | 0.05 critical value |
|---|---|---|---|
| Hypothesised number of cointegrating vector(s) | | | |
| $r=0$ | $r=1$ | 19.41* | 18.88 |
| $r\leq 1$ | $r=2$ | 7.4 | 12.45 |
| | | Trace statistic | |
| $r=0$ | $r\geq 1$ | 26.81* | 25.23 |
| $r\leq 1$ | $r=2$ | 7.41 | 12.45 |

| Variable | Coefficients | Standard errors |
|---|---|---|
| Normalised cointegrating vectors (estimation period: 1965–2005) | | |
| LREX | 1 | |
| $LY^w$ | −1.85 | −0.27 |
| LREP | 0.36 | 0.28 |

| Variable | Coefficient | $t$-Ratio |
|---|---|---|
| Error correction model: $\Delta LREX_t$ (estimation period: 1965–2005) | | |
| Error-correction term: $EC_{t-1}$ | −0.04 | −3.69 |

| Diagnostics | |
|---|---|
| Adjusted $R^2$ | |
| Serial correlation | $F(1,39)$: 0.84 |
| Functional form | $F(1,39)$: 0.10 |
| Normality | $\chi^2_{(2)} = 0.92$ |
| Heteroskedasticity | $F(1,39) = 0.37$ |

*Note*: *suggests that the corresponding null hypothesis is rejected at the 5% significance level. *Additional information*: restricted intercepts and no trends and VAR = 1. List of $I(1)$ variables: LREX, LREP and $LY^w$ (with intercept). $LY^w$ is assumed an exogenous variable. List of eigenvalues in descending order: {0.377, 0.165, 0.000, 0.000}.

both world income and the relative export prices. The income elasticity of demand for exports is 1.86, which is statistically greater than one. The relative export price elasticity of demand is −0.22, which is statistically lower than one. These results suggest that Indonesia's exports are income-elastic but price-inelastic.

### 3.2. The Johansen approach

The Johansen procedure is appropriate for determining multiple cointegrating relationships among real exports, world income and the relative export prices that are $I(1)$.[12] The weak-evidence[13] on the existence of a long-run export demand relationship can be confirmed by the Johansen approach. Accordingly, the Johansen cointegration tests have been conducted. Table 4 reports the summary results.

The results suggest that there is a unique cointegrating vector among real exports, world income and the relative export prices. The fact that the maximum eigenvalue test results accord

[12] Both the ADF and the KPSS tests results suggest that all the series have a unit root in the level form but are stationary in the first-difference form. These tests results are available from the author upon request.
[13] This is suggested by the bounds testing approach.

well with those of the trace test is noteworthy because the power of the trace test is relatively low (Johansen & Juselius, 1990). Further, the eigenvalue associated with the first cointegrating vector is considerably higher than that for other vectors, indicating the presence of a unique cointegrating vector in the system (Valadkhani, 2005). With the benefit of economic theory, the cointegrating vector can be interpreted as a long-run export demand relationship (Maddala, 2003).

The lower panel of Table 4 reports the implied long-run income and price elasticities of demand for exports, as obtained by normalising the cointegrating vectors. The income elasticity of demand for exports is 1.85. This value is consistent with that obtained by the bounds testing approach. An implication of a greater than unity income elasticity is that *ceteris paribus*, a rise in world income raises the demand for Indonesian exports more than proportionately. Indonesian exports are therefore considered income-elastic. The relative export price elasticity of demand for exports is −0.36. The likelihood ratio test rejects the condition that this coefficient value is not statistically different from one. This result confirms that Indonesian exports are price inelastic.

The significant error-correction term in the short-run model confirms the presence of a long-run export demand relationship, as suggested by a unique cointegrating vector.

## 4. Structural change in the export demand function

### *4.1. The recursive and rolling regression techniques*

Figs. 1 and 2 plot the recursive and rolling regression coefficients in the export demand function for Indonesia. The sample size for the rolling regression is 30. The figures show that while the income elasticity of demand for exports remains relatively stable, the relative export price elasticity has increased significantly since the mid-1990s. The estimated export price elasticity for the

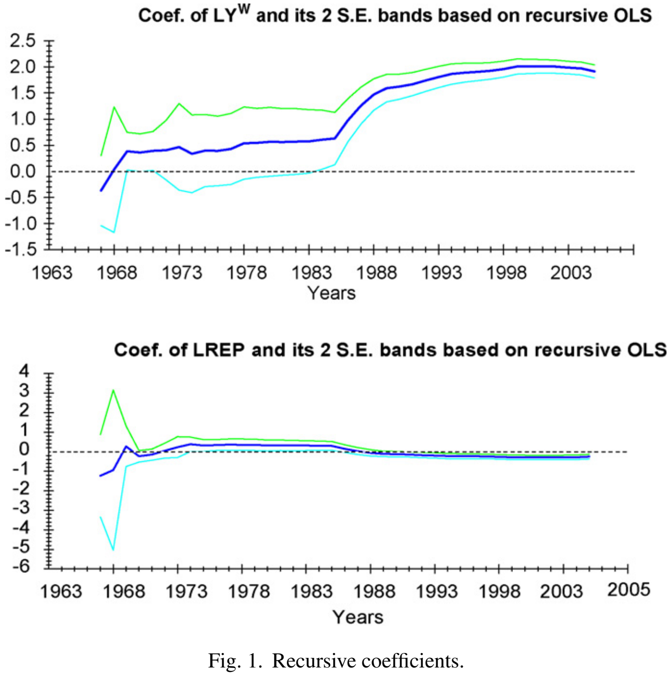


Fig. 1. Recursive coefficients.

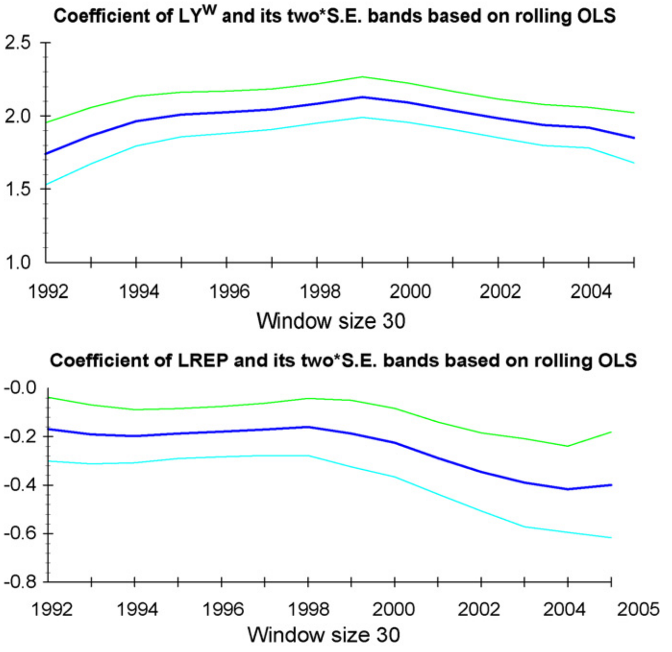


Fig. 2. Rolling regression coefficients.

sample 1975–2005 is −0.40 while the income elasticity for this period is 1.85. This indicates that the relative export prices are gradually becoming important in the export trade of Indonesia. This was one contributing factor to the recent slowdown of Indonesian exports given that the real exchange rate in Indonesia continued to appreciate after the sharp depreciation of the exchange rate during the currency crises.

### 4.2. *Stability of the cointegrating space: the Hansen–Johansen procedure*

Arize (2001) has used the Hansen and Johansen (1993) procedure to examine constancy of the cointegrating export demand relationship in Singapore. This test is conducted by estimating the cointegral relationship for the full sample and then recursively for sub-samples. To test the constancy of the cointegral relationship, Hansen and Johansen (1993, p. 8) have proposed to carry out a sequence of likelihood ratio tests based on the following statistic: $t\sum \ln((1-\lambda_i^*(t))/(1-\lambda_i(t)))$ which has a $\chi^2$ distribution with $(p-r)r$ degrees of freedom, where $p$ is the dimension of cointegrating space and $r$ is the number of cointegrating vectors. In the above equation, $\lambda_i^*$ is the largest eigenvalue obtained from the restricted sample while $\lambda_i$ is the largest eigenvalue obtained from the full sample.

Table 5 reports the test results, which show that the number of cointegrating vector remains one. The long-run income and price elasticities for the sub-samples are much higher than those for the full sample. For example, the income elasticity of demand for exports is 2.26 for the sub-sample 1965–1999 while the corresponding value is 1.84 for the full sample. Likewise, the price elasticity of demand for exports is −0.40 for the sub-sample 1965–1999 while the corresponding value is −0.36 for the full sample. It appears that the income elasticity of demand for exports has

Table 5
Temporal stability of the cointegrating equation: the Hansen–Johansen stability test

| Estimation period | Sample size | Maximum eigenvalue | | | Trace test | Cointegrating vector | Long-run elasticities | | Hansen–Johansen stability test | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | $H_O$: $r=0$; $H_a$: $r\leq 1$ | $H_O$: $r=1$; $H_a$: $r=2$ | $H_O$: $r=0$; $H_a$: $r\leq 1$ | $H_O$: $r\geq 1$; $H_a$: $r=2$ | | Real world income | Relative export price | Largest eigenvalue | HJ: $\chi^2=5.99$ (d.f. = 2) |
| 1965–1995 | 31 | 24.18 | 4.57 | 28.75 | 4.57 | 1 | 2.11 | −0.33 | 0.54152 | −9.499553622 |
| 1965–1996 | 32 | 25.34 | 4.71 | 30.06 | 4.71 | 1 | 2.11 | −0.33 | 0.54707 | −10.19572146 |
| 1965–1997 | 33 | 27.09 | 4.88 | 31.97 | 4.88 | 1 | 2.14 | −0.34 | 0.55997 | −11.46786226 |
| 1965–1998 | 34 | 28.81 | 4.99 | 33.8 | 4.99 | 1 | 2.18 | −0.36 | 0.57148 | −12.71656006 |
| 1965–1999 | 35 | 30.34 | 5.07 | 35.41 | 5.07 | 1 | 2.26 | −0.4 | 0.57977 | −13.77430962 |
| 1965–2000 | 36 | 30.14 | 5.56 | 35.7 | 5.56 | 1 | 2.17 | −0.37 | 0.56711 | −13.09932827 |
| 1965–2001 | 37 | 29.75 | 5.73 | 35.47 | 5.73 | 1 | 2.12 | −0.36 | 0.55246 | −12.23175469 |
| 1965–2002 | 38 | 29.4 | 5.68 | 35.07 | 5.68 | 1 | 2.09 | −0.36 | 0.53868 | −11.40995366 |
| 1965–2003 | 39 | 28.61 | 5.53 | 34.14 | 5.53 | 1 | 2.05 | −0.37 | 0.51984 | −10.14914558 |
| 1965–2004 | 40 | 29.09 | 4.89 | 33.98 | 4.89 | 1 | 2.03 | −0.39 | 0.51672 | −10.15030752 |
| 1965–2005 | 41 | 19.41 | 7.4 | 26.81 | 7.4 | 1 | 1.84 | −0.36 | 0.37712 | |

declined significantly over time and this has made the export demand relationship unstable. The Hansen–Johansen test statistic rejects the null of parameter constancy for all sub-samples.

## 5. Policy implications

The empirical results reported above have policy implications. Since the 1960s there has been a significant structural change in the composition of exports in Indonesia. Manufacturing products have gradually become more significant in the Indonesian export basket. For example, the share of machinery and manufactures has increased from 13% in the early 1980s to about 43% in 2006. Correspondingly, the share of mineral fuels has decreased from about 77% during 1981–1985 to 28% in 2006. Manufacturing products presently contribute about 50% of total exports. This makes exports sensitive to the relative export prices. The recent IMF studies show that one of the contributing factors to the slow growth of exports in Indonesia has been the appreciating real exchange rate, which was partly due to the inflows of portfolio capital (IMF, 2004, 2007). An implication is that Indonesia needs to give importance to its exchange rate policy. The aim of its managed exchange rate policy should be to maintain the real exchange rate at the optimal level.[14] In particular, to avoid creating overvaluation of the real exchange rate, Indonesia needs to maintain disciplined economic policies[15] and contain inflation through monetary policy.[16] The empirical results suggest that the global demand for Indonesian products has also declined. This is reflected in the falling income elasticity of demand for Indonesian exports, which were earlier dominated by highly income-elastic primary and resource products. Indonesia no longer enjoys the luxury of exporting such products rather are exporting standardised manufactured products that compete with other countries in the Asia-Pacific (IMF, 2004, 2007; James, Ray, & Minor, 2003). Global cyclical economic factors have also become significant for Indonesia.[17] Given such structural change in the export trade and international trade environment, Indonesia needs to devise a long-term strategy to raise the quality of its exportables through adaptation of better technology and encouragement of foreign investment in moderately high-tech industries. The focus should be to move towards human-capital based products and remain less dependent on resource-based and/or standardised manufactured products which face tremendous competition from relatively low-cost countries such as China, India and Vietnam.[18]

[14] The optimal or equilibrium real exchange rate is not constant but changes as the fundamentals of the real exchange rate change. The fundamentals of the real exchange rate include the international terms of trade, the level and composition of government expenditure on non-tradables, controls over trade and capital flows, technological progress and capital accumulation (Edwards, 1989).

[15] The IMF (2007, p. 28) reported that although the exchange rate of the *Rupiah* has shown considerable flexibility in nominal terms, it has appreciated by more than 50% in real effective terms in the last 6 years and that much of the real appreciation reflected the above-average inflation.

[16] The major reason for the surge of inflation after the 1997 currency crisis was excessive money supply growth (Siregar & Rajaguru, 2005). The low credibility of Bank Indonesia's monetary policy also made it difficult for Indonesia to keep its inflation rate low.

[17] Aswicahyono and Pagestu (2000, p. 485) have pointed this out to explain why Indonesia's exports slowed prior to the crisis: "... just prior to the crisis, Indonesia was in a transition stage of moving away from comparative advantage based on its abundant resource endowment, mainly natural resource-intensive exports, such as wood, and unskilled labor-intensive exports, such as garments and textiles, and moving to more technology- and capital-intensive exports, especially electronics (mainly consumer electronics). This was driven by foreign direct investment and to some extent lower labor costs".

[18] For detailed discussion on structural reforms for raising export growth and higher trade integration, see Aswicahyono and Pagestu (2000) and IMF (2007).

## 6. Summary and conclusions

This paper has investigated the aggregate export demand behaviour in Indonesia with annual data for the period 1963–2005. Both the Pesaran bounds testing and the Johansen cointegration tests results suggest that there exists a long-run relationship between real exports, world income and the relative export prices in Indonesia. The long-run income elasticity of demand for Indonesia's exports is significantly greater than one and the long-run relative export price elasticity of demand for its exports is significantly lower than one. The recursive and rolling regressions and the Hansen–Johansen stability test results suggest that the export demand function for Indonesia has undergone a significant structural change since the late-1990s, which is reflected in the decrease of the income elasticity, and an increase in the relative export price elasticity, of demand for Indonesian exports. These results explain the relatively slow growth of Indonesian exports since 2000 that has slowed the country's economic recovery from the financial crises of 1997–1998. Given the structural change in the composition of exports in favour of price-sensitive manufacturing products, Indonesian products can remain competitive in the international markets if it avoids overvaluation of the real exchange rate, lowers production costs by reducing inefficiencies and moves towards human-capital based products.

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