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EXPORT DEMAND FUNCTION FOR BANGLADESH'S TEA

Narayan Chandra Das

ABSTRACT

This paper investigates the existence of long run export demand function for Bangladesh's tea and tests for three conditions of small country assumption, namely, one to one relation between export price of tea and price substitute goods in the world market, infinitely own price elasticity of export demand, and no significant effect of world income on tea export. For this purpose an inverse demand function (price dependent) is estimated using two-step Engle-Granger Procedure proposed by Engle and Granger (1987) and a third step of the Engle-Granger Procedure proposed by Engle and Yoo (1991). It has been found that there exists long-run export demand function (inverse) for Bangladesh's tea. There also exists short run Error Correction Model (ECM). The results suggest that none of the three conditions of small country assumption can be rejected. This is evident both from long run model and short run ECM. Earlier study found that price elasticity of export demand for tea is zero which suggests that Bangladesh can impose export tax on tea to maximize revenue earnings from tea exports. The finding of the present study rules out such a policy of export tax.

I. INTRODUCTION

Bangladesh's tea sector is export oriented and labor intensive and thus plays important role in Bangladesh economy through export earnings and employment generation. Tea sector accounts for about 0.8 percent of GDP in Bangladesh. Tea industry directly employs about 0.15 million people. During the 1990s Bangladesh's average export of tea amounted to US$ 3.6 million (2265 tons) which accounted for 7.37 percent of total agricultural exports (WTO defined).

Bangladesh is a small tea producing and exporting country. In the 1990s Bangladesh's tea production accounted for about 1.80 percent of world tea production while tea export accounted for about 2.04 percent of world tea exports. For a small country's exports demand function a number of conditions should hold. According to Athukorala and Riedel (1991):

If a small country assumption is valid for a given country, a number of conditions should hold: (1) there should be a one to one relation between the country's export prices and prices of competitive goods in the world market; (2) changes in the volume of country's export should exert no influence on the price of the country's exports; (3) fluctuation in world income should have no significant effect on country's exports; and (4) variations in the volume of exports should be explained by changing by conditions affecting domestic supply and demand for exportables.

In fact, the first three of these four conditions are the characteristics of a competitive firm's demand function. A small country having small share of world exports is like a

The author is Senior Research Associate, BRAC, BRAC Centre, Mahakhali, Dhaka 1212.

competitive firm in the world market. Thus, it is generally argued that a small country is a price taker in the world market. However, it is to be noted here that if the exported good is not homogeneous with the competitive goods in the world market then the characteristics may not hold.

The objective of the study is to examine the existence of long run export demand function for Bangladesh's tea and to tests for first three of the aforesaid conditions of small country assumption, namely, one to one relation between export price of tea and price substitute goods in the world market, infinitely own price elasticity of export demand, and no significant effect of world income on tea export. Since Bangladesh's tea export accounts for only small share of world tea export, it is intriguing to test for small country assumption. In the policy perspective export demand elasticity is crucial to look into relative importance of external demand and domestic supply factors in determining export. If export demand elasticity is absolutely less than or equal to one, government can impose export tax, because imposition of export tax on the item increases revenue earnings from exports. If export demand elasticity is greater than one, export incentive is justified to increase exports. Thus, if price elasticity of demand is not correctly estimated, it may create ambiguity in proper policymaking.

The paper is organized as follows. After the introduction in section I, section II briefly reviews empirical literature concerning tests for small country assumption. Analytical framework and data are discussed in section III while section IV discusses estimation strategy. Section V analyses the estimation results and concluding observations and Policy Implications are made in section VI.

II. LITERATURE REVIEW

A number of studies have tested for small country assumption. Athukorala and Riedel (1991) proposed an inverse (price dependent) demand function to test for small country assumption. The study estimated export demand function for Korean exports of machinery and equipment by using two stage least squares. It has been found that none of the conditions of small country assumption can be rejected.

Athukorala and Riedel (1996) estimated export demand function for total manufacturing exports from Korea (excluding ships and floating structures) and for two sub-categories: (1) textile, clothing and footwear and (2) machinery and equipment using quarterly data for the period 1977q1-1992q4. The study estimated both conventional (quantity dependent) and small country (price dependent) models by using Phillips-Hansen Fully Modified OLS method and Error Correction Modelling (ECM) procedure of Hendry. It has been found that fhe estimated inverse demand function for machinery and equipment supports the small country case under both estimation techniques, but for other categories of exports the small country case is rejected by both estimation techniques.

Muscatelli *et al* (1992) estimated export demand function for Newly Industrialized Economy (NIE) exports. The study did not estimate the long-run model but obtained the long

run elasticities by applying different dynamic specification. The study found that demand for exports is price inelastic and highly elastic to the growth in the world income.

Warr and Wollmer (1996) tested small country assumption for Thailand's rice export demand function. This study found that the small country assumption is rejected by the price normalization equation of export demand function (inverse demand function).

In the context of Bangladesh a number of studies such as Ahmed *et al.* (1993), Rahman *et al.* (1995), Choudhury (2001), Roy (1991), Shilpi (1990), Razzaque (2004) have estimated the export demand function.[1] Shilpi (1990) found that price elasticity of export demand for tea is zero. This implies that Bangladesh can impose export tax on tea to maximize revenue earnings from tea exports. Razzaque (2004) found that own price elasticity of export demand for Bangladesh's RMG to the EU is very high (ranging from -11.69 to -29.91). The study concluded that Bangladesh may be considered as a 'small' country, which is compatible with the policy of export promotion and trade liberalization.

III. ANALYTICAL FRAMEWORK AND DATA

Analytical Framework

Goldstein and Khan (1978) set out traditional framework for analyzing export demand. The specification of this model differs between studies, particularly with respect to dynamics and supply structure but the core of the underlying (long-run) framework is usually a demand equation for a particular country's exports of a given commodity, or group of commodities, defined as:

$$LX_d^t = \alpha_1 + \alpha_2 L(P_X^t / P_{Xw}^t) + \alpha_3 LY_w^t \qquad (1)$$

where, X_d^t is the quantity of export demanded at time t; P_X^t is the price of export at time t; P_{Xw}^t is the export price of competing goods in the world market at time t, Y_w^t is the weighted average of incomes of the country's trading partners at time t, and L is the natural logarithm operator.

But Athukorala and Riedel (1991) argued that:

"[I]f the country were truly a price-taker, [P_X^t and P_{Xw}^t] would be perfectly, or at least very highly, collinear. In this case, the relative price variable [P_X^t / P_{Xw}^t] would exhibit very little, if any, variability. Therefore, for a true small country, the coefficient on the relative price variable cannot be precisely estimated, and may turn out relatively low (and statistically insignificant) even though its true value is extremely high." (Athukorala and Riedel,1991, pp. 144).

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[1] Ahamed et al. (1993), Rahman et al. (1995), Choudhury (2001), Roy (1991) estimated export demand function at aggregated level. Shilpi (1990) estimated both at aggregate and disaggregated levels. Razzaque (2004) estimated at disaggregated level (RMG).

Athukorala and Riedel (1991) it is more apposite to test whether a country is a "small" player in the world market for a particular commodity by estimating (1) renormalized in the price of exports, such that:

$$LP_X^t = \beta_1 + \beta_2 LX_d^t + \beta_3 LP_{Xw}^t + \beta_4 LY_w^t \quad (2)$$

The present study attempts to estimate the inverse demand function (equation 2). Thus, the testable hypothesis is as follows: β_2=0 (infinitely own price elasticity of export demand), β_3=1 (one to one relation between export price tea and price of substitute goods in the world market) and β_4=0 (fluctuation of world income has no impact on exports).

Data

Measurement of World Income

As regards export of tea, Bangladesh's major partner countries are Pakistan, China, Japan, Russia, Egypt, Jordan, Netherlands, Sudan, United Kingdom and Iran. During 1990-98, on an average, Bangladesh exported about 98 percent of its total tea exports to these ten countries. Thus, world income is obtained by taking weighted average of these countries' GDP. Weight of each country is Bangladesh's average tea exports to that country during 1990-1998 as the share of Bangladesh's average tea exports to world during the same period. Data on GDP is measured at current US dollars and collected from UNCTAD. Ten countries' GDP and weights of respective country are shown in appendix 1.

Data on Other Variables

Sri Lankan auction price of tea is taken as the price of substitute goods in the world market.[2] Data on Sri Lankan auction price is collected from International Financial Statistics (IFS). Quantity and export price of tea are collected from Bangladesh Bureau of Statistics (BBS). Quantity of tea export is measured in tons. However, export price and quantity data are given in fiscal year (Bangladesh's fiscal year starts from July) while GDP and world price are given in calendar year. In the present study, each fiscal year's data is considered as the first calendar year data.[3] Both export price of tea and international price are measured in current US dollars. To covert export price tea (measured in Taka) in US dollars, period average exchange rate is considered and collected from International Financial Statistics (IFS). Data on export price of tea, Sri Lankan auction price of tea and quantity of tea exports are shown in appendix 2.

IV. ESTIMATION STRATEGY

Time Series Properties of Data

Whether a model should be estimated using OLS, it is necessary to consider the underlying properties of the process that generate time series. Most of the time series data are

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[2]It should be noted here that there is also London auction price of tea. However, with the consideration that Sri Lanka is a South Asian country, Sri Lankan Auction Price is taken as price of substitute goods in the world market.

[3] For example data for fiscal year 1991/92 is considered as the data for calendar year 1991.

non-stationary as indicated by high serial correlation between successive observations, particularly when the sample interval is small. In the case of stationary time series, its mean, variance and auto-covariance are independent of time. Thus, the present study will check for unit roots for the variables.

The most popular tests for unit roots are the Dickey Fuller (DF) and Augmented Dickey Fuller (ADF) tests. The DF test is based on the following equation:

$$DY_t = \delta + (\rho - 1)Y_{t-1} + \psi T + e_t \quad (3)$$

where Y_t is a time series, D is the first difference operator, T is the time trend and e_t is the white noise error term. The null hypothesis is that $(\rho - 1)=0$ (Y_t is non-stationary) against alternative hypothesis of $(\rho - 1)<0$ (Y_t is stationary). The t-ratio of the coefficient of the Y_{t-1} term provides the DF test statistic. In equation (3) the estimated t-ratio of the coefficient of Y_{t-1} term is non-standard. Thus, it requires that the computed test statistic is to be compared with the corresponding critical value to infer about the stationarity of the variable. In such case, corresponding critical values were first provided by Dickey and Fuller (1981).

However, if the error term in equation (3) suffers from serial correlation, the inference drawn will not be valid. Thus, in ADF test, DF equation is augmented by incorporating lag(s) of the dependent variable as additional regressor(s) to remove serial correlation problem. The ADF test is based on the following equation:

$$DY_t = \delta + (\rho - 1)Y_{t-1} + \psi T + \sum_{i=1}^{n} \theta_i DY_{t-i} + e_t \quad (4)$$

Likewise the DF equation, the t-ratio on the coefficient of the Y_{t-1} term provides the ADF statistic.

However, for small sample test for unit roots may be complicated. Again, ADF test has low power problem which arises because of over rejection of null hypothesis when it is true and under rejection of null hypothesis when it is false. According to Hall (1986), in the case of small sample, inspection of autocorrelation coefficient function and correlogram may be important tools to check for stationarity.[4] For non-stationary time series autocorrelation coefficients die down slowly. Again, one can test whether all autocorrelation coefficients are simultaneously equal to zero or not up to certain lags. The null hypothesis is that all the autocorrelation coefficients are simultaneously equal to zero (time series is stationary) against the alternative hypothesis of all the autocorrelation coefficients not simultaneously equal to zero (time series is non-stationary). This test can be performed using Ljung-Box and Box-Pierce statistics.

Box and Pierce (1970) developed Q statistic which is defined as:

[4] It is to be noted here that the number of lags for autocorrelation coefficient is important. A rule of thumb is to compute Autocorrelation Correlation Function (ACF) up to one-third to one-quarter length of the time series (Gujarati 2003).

$$Q = n\sum_{k=1}^{m} p_k^2 \tag{5}$$

where n is the sample size and m is the lag length. The Q statistic is approximately (i.e., in large sample) distributed as the chi-square distribution with *m* degrees of freedom.

A variant of the Box-Pierce Q statistic is the Ljung-Box (LB) statistic developed by Ljung and Box (1978) which is defined as:

$$LB = n(n+2)\sum(\frac{p_k^2}{n-k}) \tag{6}$$

LB statistic is also approximately (i.e. in large sample) distributed as the chi-square distribution with *m* degrees of freedom.

Although in large samples both the Q and LB statistics follow the chi-square distribution with *m* df, the LB statistic has been found to have better (more powerful, in the statistical sense) small-sample properties than the Q statistic (Gujarati 2003).

Once it is determined that the variables are non-stationary, estimation of the regression using Ordinary Least Square (OLS) will be no longer valid. In such case using some kind of cointegration technique is essential to infer about the long-relationship.

Testing for Cointegration: The Engle-Granger Procedure

Cointegration is a statistical concept designed to deal explicitly with the analysis of the relationship between (among) non-stationary time series. Generally if two time series Y_t and X_t both are ~I(d), then the linear combination of the time series will also be ~ I(d); that is, the residuals obtained from the regression model of Y_t on X_t will be ~ I(d). If, however, there exists a vector b such that the $U_t = Y_t - bX_t$ where U_t is residuals, is of lower order of integration, I(d-q), where q>0, then Engle-Granger (1987) define Y_t and X_t as cointegrated of order (d, q). Thus, if Y_t and X_t both are ~ I(1) variables and U_t is ~ I(0), the two variables would be cointegrated. This is known as the first step of Engle-Granger cointegration technique.

However, to test the null hypothesis of non-cointegration (U_t is ~I(1)) against the alternative hypothesis of cointegration (U_t is ~ I(0)), one can use ADF test. Likewise conventional ADF equation, ADF equation on residuals does not include the constant term, because by definition the mean of residuals should be zero. The ADF equation for residuals is as follows:

$$DU_t = \phi U_{t-1} + \sum_{i=1}^{n} \lambda_i DU_{t-i} + \omega_t \tag{7}$$

where D is the first difference operator, U_t is the residual obtained from static model and ω_t is white noise error term. However, Hansen (1992), on the basis of Monte Carlo

experimentation, showed that irrespective of whether U_t contains a deterministic trend or not, inclusion of a time trend in equation (7) results in a loss of power (i.e. leads to under-rejecting the null of no cointegration when it is false). Thus, trend term is not included in equation (7). The null hypothesis is that ϕ=0 against alternative hypothesis of ϕ<0. The critical values for the ADF test on the residuals were estimated by Engle and Granger (1987) and Mackinnon (1991). However, as in the case of small sample, test for stationarity may be complicated and ADF test may suffer from low power problem, the study will also check for autocorrelation coefficients of residuals.

According to Engle and Granger (1987) if the variables were cointegrated there would have existed an Error Correction Model (ECM). The error correction model will capture all the dynamics of the long run behavior. This is known as the second step of Engle-Granger procedure. The error correction model is constructed by regressing the dependent variable on its stationary form (i.e. first difference in the case of I(1) variable) on its own lagged values and current and lagged of independent variables on their stationary form and lagged of the residual term obtained from the long run model estimated at the first step.

There are two ways of estimating ECM, namely, general to specific method and specific to general method. The present study has estimated ECM following general to specific method of Hendry (1979, 1995). Suppose one has two variable regression Y on X where both Y and X are ~ I(1) then ECM of the Engle-Granger procedure can be written as:

$$DY_t = \beta_0 + \sum_{i=1}^{m} \beta_{1i} DY_{t-i} + \sum_{i=0}^{m} \beta_{2i} DX_{t-i} + \beta_3 U_{t-1} + \varepsilon_t \qquad (8)$$

where, U_{t-1} is the lag of residual obtained from the first step, ε_t is the white noise disturbance term and D is the first difference operator. Note that in equation (8), all variables are stationary implying that all test statistics are valid. A valid ECM requires that: (1) the sign of the coefficient of error correction term will be negative; (2) it should be highly significant, and (3) it should be absolutely less than one.

The primary problem in the two-step Engle-Granger cointegration technique is that it gives us the indication that whether there exists a log run relationship or not, but is not possible to test whether a particular variable is significant or not in the long run model. Thus, the study used third step of Engle-Granger procedure proposed by Engle and Yoo (1991) to draw valid inference on the parameters and thus to test the characteristics of the long run demand function.

Assuming that there is both a unique cointegration vector and weak exogeneity of the right hand side variables in the short-run ECM (equation 8), and then the third step provides a correction to the first-stage estimate of the coefficients and ensures that it has a normal distribution. Suppose one has a two variable regression model Y on X, then the third step estimation technique can be analyzed as follows.

After estimating the static long run model ($Y_t = b_1 + b_2 X_t + U_t$), residuals are obtained, which enter the second stage regression as one period lag. The coefficient of the

lagged residual term which is known as the speed of adjustment is used in a third step-regression:

$$\varepsilon_t = \alpha_1 + \alpha_2(\vartheta * X_{t-1}) \quad (9)$$

where ε_t is the residual obtained from the second stage regression (i.e. ECM) and ϑ is the speed of adjustment obtained at the second stage.

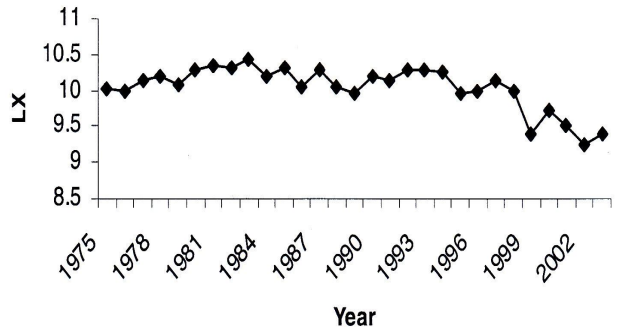
Then, adding α_2 and b_2 we get the final coefficient. The standard error on α_2 term will provide the correct standard error for the final coefficient ($(\alpha_2 + b_2)$.

V. RESULTS AND DISCUSSION

Examining the time series properties of data

The estimation begins with the identification of whether the variables are stationary or not. DF and ADF tests on the variables on their levels and first differences are provided in table 1. Since the data is annual in nature following the usual practice one lag is taken in the ADF regression. Figures 1-8 show the plot of LX (log of quantity of exports), LPX (log of export price of tea), LPW (log of Sri Lankan auction price) and LY (log of partner countries' weighted average income) and on their levels and first differences. It appears that all the level variables show trend in the data (figures1- 4). Thus, DF and ADF regressions for all the level variables include trend term. On the other hand, none of the first differenced variables show any clear trend (figures 5-8). Thus, DF and ADF regressions for all the first differenced variables exclude trend term. DF and ADF regressions for all level and first differenced variables include constant term. Table 2 shows the Box-Pierce and Ljung-Box test statistics of the autocorrelation coefficients on their levels and first differences up to lags 8 and 9. The null hypothesis is that all autocorrelation coefficients are simultaneously equal to zero (time series is stationary) against the alternative hypothesis of all the autocorrelation coefficients not simultaneously equal to zero (time series is non-stationary).

Figure 1: Plot of LX



It appears from table 1 that in the case of LX, DF and ADF tests accept the null hypothesis of non-stationary since absolute computed DF and ADF statistics are less than the critical values. Again, Box-Pierce and Ljung-Box tests reject the null hypothesis of all the

autocorrelation coefficients simultaneously equal to zero both up to lags 8 and 9 (table 2). Thus, LX is concluded to be a non-stationary variable. However, in the case of DLX, DF and ADF tests reject the null hypothesis of non-stationary and the null hypothesis of all the autocorrelation coefficients simultaneously equal to zero is accepted both up to lags 8 and 9. Therefore, LX is concluded to be ~ I(1) variable.

Table 1. DF and ADF tests for the variables on their levels and first differences.

Variable	DF	ADF
LX	-2.744	-1.519
LPX	-3.954	-3.352
LPW	-4.145	-5.250
LY	-4.391	-3.423
DLX	-7.524	-5.155
DLPX	-7.131	-5.316
DLPW	-5.649	-6.080
DLY	-7.431	-6.000

Note: 1. All the level variables include trend term. None of the first differenced variables include trend term.

2. D denotes first difference operator.

3. Critical values for DF and ADF tests at 5 percent level of significance without trend term and with trend term are -2.98 and -3.59 respectively.

For LPX, although DF rejects the null hypothesis of non-stationary, ADF test accepts the null hypothesis. But, Box-Pierce and Ljung-Box tests accept the null hypothesis of all autocorrelation coefficients simultaneously equal to zero both up to lags 8 and 9. Thus, it is difficult to check the stationarity of this variable. However, DLPX is stationary supported by DF and ADF tests, and test of autocorrelation coefficients. Therefore, LPX may be ~ I(1) or ~ I(0) variable.

Table 2. Box-Pierce and Ljung-Box statistics of autocorrelation coefficients.

Variable	Up to lag 8		Up to lag 9	
	Box-Pierce Statistic	Ljung-Box Statistic	Box-Pierce Statistic	Ljung-Box Statistic
LX	30.563[.000]	35.085[.000]	31.000[.000]	35.763[.000]
LPX	5.354[.719]	6.6413[.576]	7.374[.598]	9.771[.369]
LPW	12.849[.117]	15.810[.045]	16.502[.057]	21.472[.011]
LY	48.581[.000]	57.729[.000]	48.943[.000]	58.291[.000]
DLX	8.112[.423]	9.892[.273]	8.324[.502]	10.227[.332]
DLPX	5.973[.650]	7.1909[.516]	7.380[.598]	9.412[.400]
DLPW	12.865[.117]	16.988[.030]	14.754[.098]	19.970[.018]
DLY	6.396[.603]	7.447[.489]	6.489[.690]	7.594[.576]

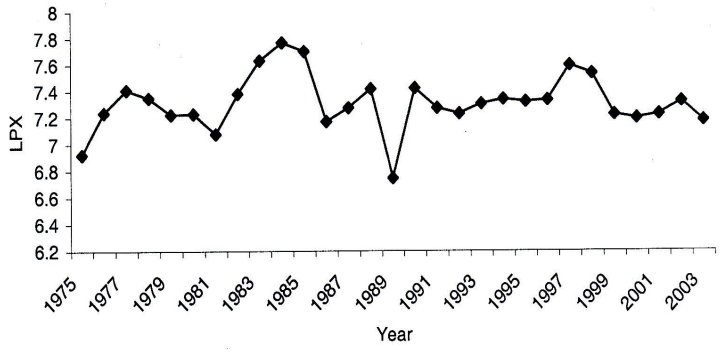
Note: figures in the parentheses denote p-value.

LPW is found to be stationary according to both DF and ADF tests. However, Ljung-Box tests reject the null hypothesis of all the autocorrelation coefficients simultaneously equal to zero both up to lags 8 and 9. Thus, it is problematic to conclude that this variable is stationary. Both DF and ADF tests suggest that DPW is stationary although null hypothesis of

all autocorrelation coefficients simultaneously equal to zero is rejected by Ljung-Box test at 5 percent level of significance both up to lags 8 and 9. However, plot of DPW (fig 7) seems to suggest that DPW may be stationary. Thus, it would not be inappropriate to infer that DPW is stationary. Therefore, LPW may be either ~ I(0) or ~ I(1) variable.

DF test rejects the null hypothesis of non-stationary in the case of LY, but ADF test accepts the null hypothesis. Again, test of autocorrelation coefficient suggests that LY is non-stationary. Since, in general, ADF test is considered to be better than DF test, LY can be considered as a non-stationary variable. But DLY is stationary according to both DF and DF tests and test of autocorrelation coefficients. Thus, LY is concluded to be ~ I(1) variable.

Figure 2: Plot of LPX



Engle-Granger Cointegration Technique
First step of Engle-Granger Cointegration Technique

OLS estimation of the static model (equation 2) is as follows:

$$PX = 0.563 + 0.244\,LX + 0.778LPW - 0.135LY \quad (8)$$

$R^2 = 0.517$, D.W. = 2.141

In the above estimated regression results, diagnostic test statistics and standard errors are not reported because they are not valid for non-stationary variables. In the model R^2 is less than Durbin-Watson statistic, implying that one cannot be sure whether it is spurious regression or not.[5] It should be noted here that the estimation results in equation (8) may be super consistence if there exists cointegration.[6] A quicker method of testing cointegration is

[5] If R^2 is greater than DW statistic, the estimated regression is spurious but if otherwise one can not infer whether the regression is spurious or not.

[6] The super consistency property states if all the variables are non-stationary ~ I(1), and U_t (residuals) is ~ I(0), then as sample size increases the OLS estimator of the coefficients converges to its true value at a much faster rate than the usual OLS estimator of I(0) variables.

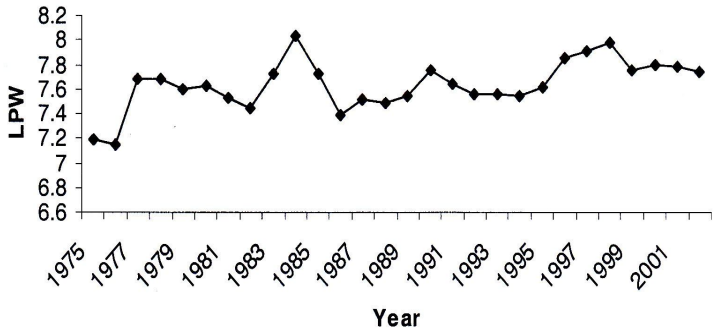
the cointegration Regression Durbin-Watson (CRDW) test proposed by Sargan and Bhargava (1983). In this method DW statistic is compared with the CRDW critical value.[7] If DW statistic is greater than CRDW critical value, the null hypothesis of non- co-integration is rejected. In our case CRDW critical value is 1.1. Since DW statistic is greater than CRDW critical value the null hypothesis of non-cointegration is rejected.

Table 3. Autocorrelation coefficients, and Box-Pierce and Ljung-Box statistics of residuals.

Order	Autocorrelation Coefficient	Standard Error	Box-Pierce Statistic	Ljung-Box Statistic
1	-0.085	0.186	0.208[.648]	0.231[.631]
2	0.039	0.187	0.254[.881]	0.283[.868]
3	0.013	0.187	0.259[.968]	0.289[.962]
4	-0.295	0.187	2.780[.595]	3.419[.490]
5	-0.221	0.203	4.203[.521]	5.253[.386]
6	-0.130	0.211	4.698[.583]	5.920[.432]
7	-0.096	0.214	4.963[.665]	6.293[.506]
8	0.008	0.215	4.964[.761]	6.295[.614]
9	0.202	0.215	6.153[.725]	8.137[.520]

Note: Figures in the parentheses denote p-value

Figure 3: Plot of LPW



However, the powerful test for cointegration is the test for stationarity of the residuals. The estimated ADF regression for the residuals is as follows:

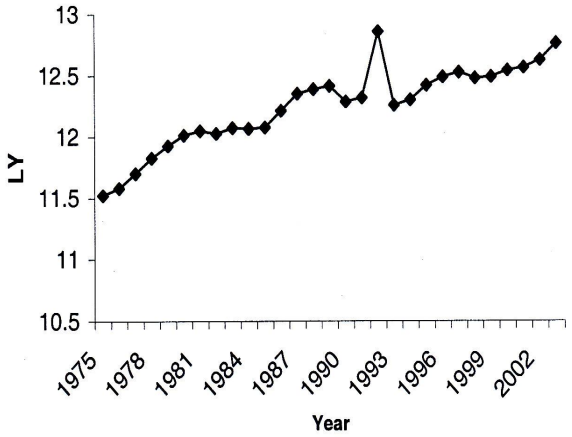
$$DR = -0.995R(-1) - 0.039DR(-1) \quad (9)$$

t-ratio –3.52 -0.20

[7] The critical values for CRDW test can be found in Maddala (1992).

At 5 percent and 10 percent level of significance, critical values for the ADF test of unit roots for residuals with 4 variables in the cointegrating regression are respectively -4.10 and -3.81.[8] The absolute computed t ratio is less than critical value, implying acceptance of non-cointegration. Thus, ADF test on residuals contradicts the CRDW test result. Let us check autocorrelation coefficients of the residuals. The autocorrelation coefficients of residuals along with Box-Pierce and Ljung-Box statistics are reported in table 3. It appears that null hypothesis of all autocorrelation coefficients simultaneously equal to zero is accepted by both Box-Pierce and Ljung-Box tests irrespective of order of the lags.

Figure 4: Plot of LY



According to Bartlett (1946), if a time series is purely random, that is, it exhibits white noise, the sample autocorrelation coefficients are approximately normally distributed with zero mean and variance 1/n, where n is the sample size. By following the properties of the standard normal distribution, 95% confidence interval for any sample autocorrelation coefficients will be ± 1.96($1/\sqrt{n}$). Given the sample size of 29, the interval is -0.364 to + 0.364. Thus if any autocorrelation coefficient falls inside the interval, one can accept the null hypothesis that the true autocorrelation coefficient is zero. The autocorrelation coefficients reported in table 3, fall insides the interval -0.364 to + 0.364, suggesting that residuals are stationary. Again the plot of residuals (figure 9) seems to suggest that it is stationary. Thus, it would not be inappropriate to conclude that residuals obtained from regression equation (8) are stationary which in turn suggest that the null hypothesis of non-cointegration is rejected.

Second Step of Engle-Granger Cointegration Technique

As evidence suggests that there exists cointegration, the next step is to estimate short run ECM. Using the notation of general to specific model, at first two lags of the first differences

[8] The critical values are taken from Davidson and MacKinnon (1993) and cited in Johnston and Dinardo (1997), p.265.

of the explanatory and dependent variables are taken. Estimated general model is presented at in second column of table 4. After estimating the general model, iteratively most insignificant variables are eliminated. The most parsimonious estimated model is shown in the last column of table 4.

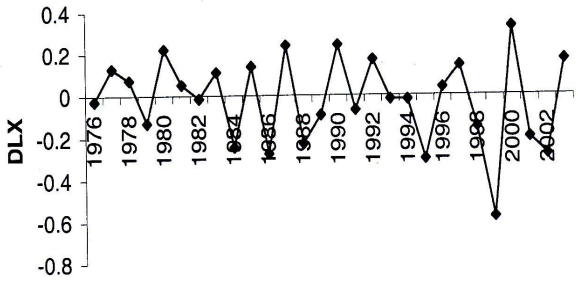
Table 4. Short run error correction model of Engle-Granger procedure.

Variable	Coefficient (1)	Coefficient (2)	Coefficient (3)
DLPX(-1)	-0.155 (.733)		
DLPX(-2)	-0.0407 (.894)		
DLPW	0.604 (.066)	0.704(.006)	0.754 (.000)
DLPW	0.180 (.646)	0.073(.693)	
DLPW(-1)			
DLPW(-2)	-0.167 (.597)		
DLX	0.235 (.357)	0.166(.369)	0.099(.540)
DLX(-1)	0.146 (.668)		
DLX(-2)	0.106 (.709)		
DLY	-0.242 (.429)	-0.246(.318)	-0.119(.526)
DLY(-1)	-0.254 (.440)	-0.189(.460)	
DLY(2)	-0.253 (.415)	-0.178(.466)	
R(-1)	-0.861 (.135)	-1.065(.000)	-1.124(.000)
Adjusted R^2	0.475	0.594	0.636
Serial correlation χ^2 (1)	0.237	0.110	1.920
Heteroscedasticity χ^2 (1)	0.037	0.054	0.035
Normality χ^2 (2)	5.878	4.795	5.894
Functional Form χ^2 (1)	0.004	0.012	0.003

Note: 1. Critical values of χ^2 (1) and χ^2 (2) at 5 percent level of significance are respectively 3.84 and 5.99.
2. Figures in the parentheses (in 2nd, 3rd and 4th columns) denote the p-values.
3. R is the residual obtained at the first stage regression.

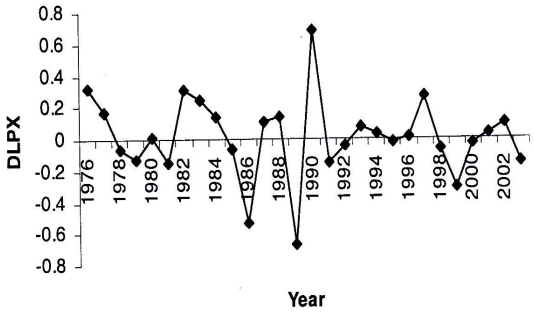
In the final short run model only the DLPW is significant which bears expected sign. Coefficients of DLX and DLY have unexpected sign, but they are not significant (table 4). The coefficient of the error correction term (R(-1)) bears expected sign and highly significant. As has been noted earlier, the coefficient of the error correction term should be absolutely less than one. But the estimated coefficient is not absolutely less than one. Let us impose restriction on the coefficient of error correction term as -1. This test is performed using Wald test which follows chi-square distribution with degrees of freedom (df) equal to the number of restrictions. The computed chi-square value is 0.158 against the critical chi-square value of 3.84 at 5 percent level of significance and degrees of freedom 1. Thus, null hypothesis is accepted. The restriction on the coefficient of R(-1) as -0.8 is further imposed. In this case computed chi-square is found to be 2.99 against the critical chi-square value of 3.84 at 5 percent level of significance indicating acceptance of null hypothesis. Thus, one cannot deny that the coefficient of the error correction term is absolutely less than one. Note that speed of adjustment is close to -1 which implies that it takes just over one year to return to the long run equilibrium path once there has been any deviation from equilibrium.

Figure 5: Plot of DLX



To check for diagnostics of the short run model, Godfrey's (1978) LM test for serial correlation, Ramsey's (1969) RESET test for functional form, White's (1980) general heteroscedasticity test for heteroscedasticity and Jarque-Bera (1987) test for normality were performed. All the test statistics follow chi-square distribution. It appears that in the final error correction model null hypothesis of no serial correlation, no heteroscedasticy, normally distributed residuals and no problem with functional form are accepted since each computed chi-square value is less than corresponding critical chi-square value at 5 percent level of significance.

Figure 6: Plot of DLPX



Before going to third step of Engle-Granger procedure where test for the characteristics of the long run export demand function is performed, effort has been made to test the characteristics for the short-run model. In the model coefficients of both DLX and DLY are insignificant, suggesting that in the short run export demand elasticity with respect to its price is infinite and variation in the income of the partner countries has no impact on tea export. To test whether the coefficient of the DLPW is one or not, Wald test is used. Estimated Wald test statistic is found to be 1.86 against the critical chi-square value of 3.84 at 5 percent level of significance. Thus it appears that the three conditions of small country assumption do hold for short run model.

Third Step of Engle-Granger Cointegration Technique

Results of the third step of Engle-Grange procedure are presented in table 5. Before interpreting the results, a word of caution should be noted here. Estimated speed of adjustment at the second step of Engle-Granger procedure was found to be -1.124. However, by imposing restriction, we have found the speed of adjustment less than 1 in absolute term (around -0.80). Third step results are obtained assuming speed of adjustment -0.80. It appears from table 5 that coefficient of LPW is significant at 1 percent level of significance. Coefficients of both LX and LY are not statistically different from zero even at 10 percent level of significance. However, the estimated coefficient of LPW is found to be 0.78. Let us test the hypothesis that the coefficient of LPW is equal to 1 using t-test. The estimated t-ratio is found to be only -1.06 against the critical t-ratio of 2.06 at 5 percent level of significance. Therefore, null hypothesis is accepted.

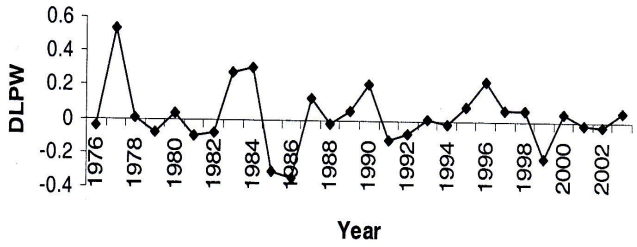
Table 5. Third step results of Engle-Granger procedure

Variables	Coefficient	Standard errors	t-ratio
C	0.630	2.001	0.315
LPW	0.781	0.204	3.829
LX	0.205	0.136	1.503
LY	-0.097	0.136	-0.713

Note: Critical t-ratios at degrees of freedom 25 and 1%, 5% and 10% level of significance are 2.79, 2.06 and 1.71 respectively.

Thus, there exists one to one relation between export price of tea and price substitute goods in the world market; elasticity of export demand is infinite and variation in income of partner countries has no impact on tea exports.

Figure 7: Plot of DLPW

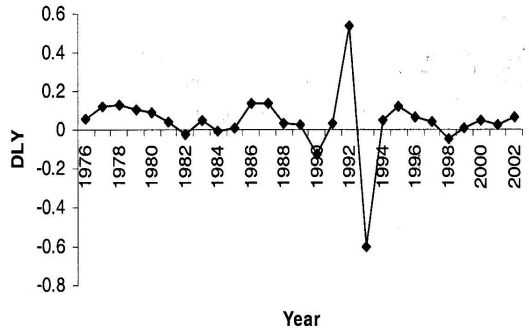


VI. CONCLUDING OBSERVATIONS AND POLICY IMPLICATIONS

The objective of this paper was to test for existence of a long run export demand function for Bangladesh's tea and test for three conditions of small country assumption, namely, one to one relation between export price of tea and price substitute goods in the world market, infinitely elasticity of export demand, and no significant effect of world income on tea export.

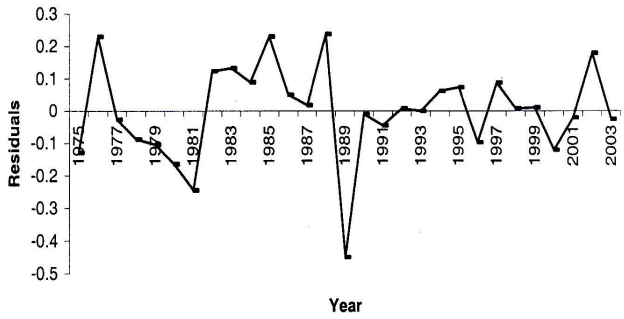
For this purpose this study has estimated an inverse demand function proposed by Athukorala and Riedel(1991) using two steps Engle-Granger cointegration technique proposed by Engle and Granger (1987) and third step of Engle-Granger technique proposed by Engle and Yoo (1991) where the later allows to draw valid inference on the parameters and thus to test for the characteristics.

Figure 8: Plot of DlY



It has been found that there exists a long-run export demand function for Bangladesh's tea. There also exists a short run error correction model. The result suggests that none of the three conditions of small country assumption can be rejected. This is evident not only from long run static model but also from Error Correction Model (ECM). Thus small country assumption holds for Bangladesh's tea exports.

Figure 9 : Residuals obtain from Long Run Reression



The findings of this study suggest a number of policy implications. Earlier study found that price elasticity of export demand for tea is zero which suggests that Bangladesh can impose export tax on tea to maximize revenue earnings from tea exports. We have found that price elasticity of export demand is infinite. Thus the possibility of export tax on tea is ruled out. The findings of the current study also suggest that domestic factors (such as conditions

those affect domestic production and consumption of tea) may the important determinants of tea export. Therefore, to increase export of tea, Bangladesh will need to enhance its supply side capacity by expanding area under cultivation, and using high yielding varieties.

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