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A Macroeconometric Model of Bangladesh: Specification and Estimation

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1. Introduction

Macroeconometric modelling has a long tradition in economics, which originated with the work of Jan Tinbergen in the 1930s. Attempts to construct macroeconometric models of the Bangladesh economy have been rather limited and policy evaluations with the help of such models are rarely undertaken. Therefore, analytical exercises with regard to the effectiveness of different alternative options have not contributed much to the macro policy discourse of the country. It is in this context that the present paper develops a small macroeconomic model for Bangladesh to account for a number of important macro linkages in the economy and thus seeks to provide an operational model that can be used for policy analysis and for tracing out the impacts of various exogenous changes on the economy.

The model developed in this paper utilizes the small open-economy framework and attempts to take into account the unique features of the Bangladesh economy. It has 30 equations of which 18 are behavioural and the rest are identities. For convenience, the model can be shown to have 6 blocks, viz., (i) production, (ii) investment, (iii) private consumption expenditure, (iv) foreign trade, (v) government and (vi) monetary and price blocks. It captures, amongst others, the nexus between output, government deficit, foreign trade, money supply, and price level and can be used to examine the effects of both domestic and external shocks to the economy. The model is also suitable for examining the effects of monetary, fiscal and exchange rate policies on the overall economy.

The modelling exercise undertaken in this paper has several distinguishing features leading to important contributions to the applied macroeconomics literature on Bangladesh. First of all, very recently the Bangladesh Bureau of Statistics (BBS) has published new national income estimates for the country by incorporating extensive methodological and data improvements. This has resulted in an increase in national income by 26-43 per cent (in nominal terms) for every year between 1980 and 2000.[1] If the new national income estimate is to be a true reflection of Bangladesh's economy, the previous empirical research works using the old estimates of GDP must have encountered measurement errors, which the present paper overcomes.[2]

Second, most macroeconomic models tend to rely on a Keynesian framework emphasising only on the demand constraint of the economy. According to the Keynesian theory any policy that can stimulate effective demand would prove effective in terms of enhancing the economic growth performance. However, it should be kept in mind that in many developing countries including Bangladesh supply side problem is also important. It has been argued that in a supply constrained economy increasing effective demand through increased government spending gives rise to inflation instead of increasing employment and output (Mallick, 1999). As a result, recent macroeconometric modelling exercises have attempted to incorporate the supply side of the economy. In most cases, this is done by specifying an aggregate production function (e.g., Musila, 2002). In the present paper, in order to gain further insight into the supply constraint we specify production functions for agriculture, manufacturing and services sectors separately. There is no denying that production dynamics in these three sectors are quite different and hence the use of different specifications may be considered to be more appropriate. Therefore, along with the demand side of the economy the model in this paper takes into consideration disaggregated production functions.

Third, in macroeconometric modelling, the sector-specific equations are often chosen by using some kind of trial and error process. It involves running several regression equations and then selecting 'the best' one on the basis of such criteria as the goodness of fit and sign and statistical significance of the coefficients. This practice is susceptible to reflect the model builders' prejudice about what is to be expected (rather than the actual outcome) and might result in empirical specifications that lack theoretical justification. This method of running a number of equations and then choosing one of them is no longer considered as a good

[1] The revised national income estimates by the BBS are available only since 1980.

[2] If the errors of measurement are associated with the dependent variables, the ordinary least squares (OLS) estimates, which are usually used to estimate the model equations, are unbiased as well as consistent but they are less efficient. However, if there are measurement errors in the explanatory variables, the OLS estimators are biased as well as inconsistent.

practice in 'modern econometrics'.[3] In light of the shortcomings of the traditional econometric methodology in this paper we specify just one single equation for each sector (or sub-sector) and remain stick to it in the estimation stage. Consequently, the danger of specification mining is eliminated. Moreover, plausible signs on the parameters and their statistical significance in the estimated equations seem to suggest appropriateness of our specifications.

Fourth, another important dimension of the present modelling exercise has been the explicit consideration of the time series properties of the data and use of estimation techniques suitable for dealing with the non-stationary data. Previous studies on Bangladesh (and for many other countries) failed to recognise the problem while in this paper all variables are tested to determine whether they are stationary or not. Since most of the variables appear to be non-stationary on their levels, cointegration techniques, which can only provide for valid long-run relationship under such circumstances, have been used for empirical estimation of the models. In most cases, the short-run dynamic equations are estimated following the error-correction modelling technique, which is compatible with the cointegration methodology. The estimation of the models while vividly illustrates the difficulties associated with the determination of the time series properties of the variables involved, convincing evidence for valid long-run relationships are obtained for a majority of the equations. It is also to be noted that unlike the previous studies, serious attention has also been given to various diagnostic tests concerning the estimated models.[4]

Fifth, one interesting finding of the current exercise is that in some equations there is evidence of structural breaks in the parameters. This includes the equations for private consumption expenditure, investment in agriculture, investment in manufacturing, and investment in services. It was revealed that in all cases the structural breaks were due to apparently different behaviour of the model for the 1980s and 1990s. Without inferring about the direction of causality, it seems that those structural breaks are associated with the post reform period as Bangladesh implemented a number of macro and trade policy reforms in the 1990s.

[3] Discussions on the traditional and modern approaches to econometric methodologies may be found in Charemza and Deadman (1994).

[4] Previous studies only considered the serial correlation problem. Other problem such as non-normality of errors, functional form and heteroscedasticity hardly received any attention.

The remainder of this paper is organised as follows: Section 2 briefly reviews the past attempts of macroeconometric modelling in Bangladesh. Section 3 describes the specification and structure of the model along with the illustration of the underlying linkages. Section 4 elaborates the estimation strategies for sector-specific individual equations while the estimation results are presented in Section 5. Finally, some concluding observations are placed in Section 6.

2. A Brief Review of Macroeconomic Models of Bangladesh

The earliest attempt to construct a macro model of Bangladesh can be traced back to the work of Islam (1965). He built a macroeconometric model of Pakistan in which Bangladesh (the then East Pakistan) was treated as a region. The objectives of this early exercise were to collect, organize, and process systematically both the published and unpublished data together, formulate the model, and undertake statistical estimation. In that model there were 50 equations of which 20 were behavioural and 62 variables of which 12 were exogenous.

The model developed by Islam (1965) was highly aggregative and the specifications of the equations were simple. The simple nature of the equations was partly forced on the model by data limitations. Amongst serious omissions, the specification of the equation for agricultural output did not recognize the importance of weather conditions and the manufacturing output was not constrained by imported intermediate and raw materials.[5] Besides, the model contained only the real sector of the economy and thus the link between real and monetary sector was not considered.

'A Macro-econometric Model of Bangladesh' by Rashid (1981) was developed at the University of New Castle. The model was also quite simple containing six sectors, viz., expenditures, production, monetary, government and foreign trade. There are only 8 behavioural equations in the model and 9 definitions and identities. The model, however, represents much realism by capturing both demand and supply constraints on the macroeconomic relationships. Like Islam (1965), Rashid's model is also highly aggregative and cannot explain some special features of the economy. For example, the random shocks in production (mainly in agriculture) created by weather conditions are not recognized. Moreover, the specification of the output equation with labour is of questionable relevance in a labour surplus economy. Further, the price determination process is not stressed in the model.

In another attempt Hossain (1995) developed a macroeconomic model using quarterly data for the period 1974:2 to 1985:4. The main objective of the model was to explain the determination of inflation, economic growth, and the balance of payments. The model consisted of six blocks, viz., the fiscal sector, the monetary sector, the expenditure sector, the production sector, the external trade and payments sector, and wages and prices. In the fiscal sector, different types of nominal taxes are determined which together with exogenous government expenditures determine the budget deficit. The money supply is determined by reserve money and lagged money supply. Reserve money is determined by its various components-both domestic and external. Among the domestic components, Bangladesh Bank's credit to the government is related to government budget deficit and Bangladesh Bank's credit to scheduled banks is related to their total credit and their loan rate. Net foreign assets is determined through an identity in the external sector.

The specification of the private expenditure equation is considered in the expenditure block. The private expenditure is modelled in the spirit of Monetary and New Cambridge schools, which consider private consumption and investment expenditure jointly. Private expenditure is determined by real permanent income, the stock of real money balances and the expected rate of inflation. Production is divided into agriculture, manufacturing, and services. Agricultural production is determined by planted acreage and per acre yield. The acreage response function is related to 4 quarter lagged wholesale agricultural price and yield. Yield per acre is determined by fertilizer and rainfall. The short-run variation in manufacturing production is not constrained by capital stock. It depends on real expenditure, real imported raw materials and capital goods, real bank credit and a dummy variable capturing effect of denationalization. Services sector output is determined by real expenditure and real bank credit. On the external front, exports of tea, fish and leather goods are determined under small country assumption but in the case jute and jute goods Bangladesh is believed to influence the world market price. The imports of major commodities are explained by real expenditure,

[5] However, the price of manufactures was influenced by price of imported intermediates and raw materials.

relevant relative price variable and lagged imports. Government adjusts exchange rate on the basis of movement of relative price of traded and non-traded goods.

The complete model consists of 113 equations – 34 of which are behavioural – and has 46 exogenous variables. The model is made dynamic by including the lagged dependent variable. The causal structure of the model is as follows: an increase in budget deficit, caused by either increase in government expenditure or reduction in government revenue, increases reserve money, which increases money supply.[6] This, in turn, raises both the price level and imports by increasing real expenditure. Increases in expenditure and imports increase output and government revenue, which reduce the budget deficit and hence the stock of money.

Although the model used by Hossain captures the linkages in different sectors of the economy, specifications of some of the equations might be problematic. In the fiscal sector, government expenditure is treated as an exogenous variable. The revenue-expenditure of the government has certain degree of automaticity and irreversibility so that it can be regarded as an endogenous variable. This also indicates the limitation of fiscal policy changes especially in the case of reduction of public expenditure. The modelling of private expenditure by aggregating consumption and investment expenditure has been done on the grounds of stability of the function. The lack of separate private investment function in the model misses the link between private investment and the financial sector. Separate treatment of private investment is needed especially in view of its growing importance in the economy. Further, the crowding out or complimentary effect of public expenditure cannot be assessed if private investment is not included in the model separately. Both manufacturing and services production are determined by demand variables and scarce factors of production. The logic of including both demand variables and factors of production in production function can be questioned.

The model developed by Rahman and Shilpi at the Bangladesh Institute of Development Studies (1996) was an extension of an earlier work undertaken by Rahman (1990) under the South Asian Link Model Project sponsored by the Asian and Pacific Development Centre. The model comprises 42 equations of which 21 are behavioural equations and the rest are

[6] Reserve money can also change because of changes in net foreign assets caused by changes in exports or imports.

identities. The model consists of 5 blocks: viz., expenditure, fiscal, money and finance, trade, and aggregate supply. The model includes 32 exogenous variables and 20 dummy variables.

Real output in the model is determined from an aggregate production function, which includes accumulated capital stock and imported intermediates as the arguments. Capital stock is determined through a dynamic process in which lagged depreciated capital stock plus current investment determine the current capital stock. Investment is divided into public investment and private investment. While public investment is exogenous, private investment is endogeneized in the model. Private investment is made to depend on lagged investment, current and lagged interest rate, lagged capital stock and current output. Imports of intermediates is influenced by real output and unit price of these imports. Imports of other types of goods, exports, exchange rate and capital account balance determine the overall balance of payments in the economy.

Nominal output in Rahman and Shilpi model is determined by real output and price, the latter being determined by output, import price and money supply. Nominal output influences direct and indirect taxes. Real taxes on income and wealth are deducted from real output to derive disposable income, which determines consumption. Budget deficit is determined by the difference between exogenous public expenditure and endogenous revenue. Deficit financing by the government provides the nexus between the monetary and fiscal sector of the economy. Government borrowing from the central bank increases the monetary base and affects the money supply.

Although the Rahman-Shilpi model captures important linkages in the economy, it has certain limitations. The model uses an aggregate production function to explain supply and therefore fails to recognize that the behaviour of production may differ among sectors. The specification of the investment function lacks an appreciation of the availability of credit from banks at exogenously fixed interest rate as a constraint on investment. The estimation of the investment function experimented with different variables including change in bank credit to the private sector, which incidentally came up with a statistically significant coefficient. Further, the crowding out or complementarity relationship between private and public investment is not known. In addition, the specification of the equations in the monetary sector does not seem to be well-built into the model. For example, the equations explaining the demand for demand deposits and time deposits are not integrated in the model. Besides, one can argue that separate specification of the demand for money, the supply of money and GDP deflator are not required. As the interest rate is not determined in the model, the monetary sector is in equilibrium through changes in GDP deflator, in which case the demand for money function can be dropped from the model.

One fundamental problem associated with all the above-mentioned macroeconomic models is that the estimated regression equations are based on the assumption that the time series data used in various equations are stationary. However, there is compelling evidence to suggest the non-stationarity of many macroeconomic time series in which case the use of OLS in estimating the relationship might result in inconsistent and inefficient results and might produce spurious relationship. Therefore, the recent econometric exercises with the time series data have emphasized the time series properties of the variables and have used, in the case of the presence of non-stationary data, cointegration techniques to overcome the problem of spurious relationship. As the integrating properties of the data used in the model have not been investigated in the previous studies, the estimated long-run relationships as reflected in various specifications may be called into questions.

The brief review of the models shows that they represent valuable contribution to modelling the macroeconomic behaviour of the economy. However, each model has certain limitations. The construction of macroeconometric model of an economy is an evolving process and a particular model cannot consider all aspects of the economy. The model presented in this paper is built on the accumulated knowledge of the past models and makes an attempt to provide improved specifications and estimation.

3. Specification of the model

The model consists of five blocks – production, expenditure, foreign trade, government, and monetary and price block. It is recognized that in Bangladesh supply constraints have a major influence on the macroeconomy though in some areas demand side factors may also exert some influence.

3.1. Production Block

In the production block three equations have been specified – one each for agriculture, manufacturing and services. The services sector includes all producing sectors except agriculture and manufacturing and is termed services for brevity. In each sector, production depends on sectoral capital stock. Further, in agriculture, fertilizer and irrigation act as constraint on production. Fertilizer and irrigation are highly correlated and any one of them can be used in the model. Here ratio of irrigated area to the total acreage is used as an explanatory variable. Besides, weather conditions comprising amount and distribution of rainfall, sunshine etc. exert significant influence on agriculture. Here, an index of rainfall is included in the agricultural production function as a proxy for weather conditions.

In the manufacturing sector, availability of imported raw materials and intermediate goods are important factors of production and hence it is included in the manufacturing production function. In a labour surplus economy, labour does not act as constraint to production. Hence labour is not included in the production function. Domestic output of agricultural raw materials is also an important determinant of manufacturing output. The production functions are specified as follows:

3.1.1 Value added in agriculture

$$YAGR = \beta_0 + \beta_1 KAGR + \beta_2 RIR + \beta_3 RAIN \qquad (3.1)$$

3.1.2 Value added in manufacturing

$$YMAN = \beta_4 + \beta_5 KMAN + \beta_6 MRMR + \beta_7 DRM \qquad (3.2)$$

3.1.3 Value added in services

$$YSER = \beta_8 + \beta_9 KSER \qquad (3.3)$$

3.2. Investment Block

Investment demand in the economy is divided into private and public investment demand. Public investment is a policy variable. Private investment is classified by production sectors. Interest rate is not an argument in the investment function. Availability of credit is an important determinant of private investment. Further, public investment, which concentrates mostly on infrastructure exerts an important influence on private investment. It is often suggested that public investment complements private investment instead of crowding out in developing countries such as Bangladesh.

3.2.1 Private investment in agriculture

$$INAGRP = \beta_{10} + \beta_{11}(CRAGR / WPI) + \beta_{12}\, GI \tag{3.4}$$

3.2.2 Private investment in manufacturing

$$INMANP = \beta_{13} + \beta_{14}(CRMAN / WPI) + \beta_{15} GI \tag{3.5}$$

3.2.3 Private investment in services

$$INSERP = \beta_{16} + \beta_{17}(CRSER / WPI) + \beta_{18}\, GI \tag{3.6}$$

3.3 Expenditure Block

The expenditure block consists of only one equation for private consumption. Government consumption is treated in the government block. Private consumption is specified as a function of disposable income.

3.3.1. Private consumption expenditure

$$CON = \beta_{19} + \beta_{20}\, YD \tag{3.7}$$

3.4 Foreign Trade Block

The foreign trade block consists of 5 equations – 3 equations explain the determination of the volume of exports and 2 equations determine the volume of imports. Export of jute and jute manufactures, RMG and knitwear, and other products including frozen food, tea and leather, are specified by corresponding export supply functions. It is assumed that Bangladesh is a price taker in the world market, Consequently the export prices are given by the world market prices. It is further assumed that main constraint on Bangladesh's exports arises from the supply side. Bangladesh can export whatever it is able to supply to the export market. Export supply is determined by export price relative to domestic price level as well as the rate of utilization of capacity. Export supply of RMG and knitwear may be argued to be affected by the availability of quota in the U.S.A. and Canada. Nevertheless, a recent study (Panagariya, *et al.* 2001) has found that the price elasticity of demand for Bangladesh's exports of textiles and apparels in the US market is unusually high and more closely fits to the assumption of small country case.[7]

Behavioural relationship in the import sector has been specified for imports of raw materials and intermediate goods and for imports of other goods. Import of raw materials and intermediate goods is determined by import price relative to domestic price of raw materials and gross domestic product. Similar specification is also used for imports of other goods.

3.4.1 Exports

3.4.1.1 Exports of jute and jute manufactures

$$XJMR = \beta_{21} + \beta_{22}\frac{PXJ * EXR}{WPI} + \beta_{23} JMC \tag{3.8}$$

3.4.1.2 Exports of readymade garments

$$XRMGR = \beta_{24} + \beta_{25}\frac{PXRMG * EXR}{WPI} + \beta_{26} RMGC \tag{3.9}$$

[7] While conventional studies have reported very low price elasticity of demand for Bangladesh's exports, Panagariya *et al.* (2001) has found that such elasticity in the US market to be as high (absolutely) as –26. This high magnitude of price elasticity may be approximated by infinity. Note that as an LDC Bangladesh has quota-free access into the EU market provided that it fulfills the EU rules of origin. Therefore, export to the EU market is a direct function of Bangladesh's capacity to domestically produce (rather than importing) intermediate goods for export-oriented RMG industry. Only recently (since January 2003) Canada has allowed quota-free access to its market.

3.4.1.3 Exports of other goods

$$XOGR = \beta_{27} + \beta_{28} \frac{PXOG * EXR}{WPI} + \beta_{29} XOGC \qquad \textbf{(3.10)}$$

3.4.2 Imports

3.4.2.1 Imports of intermediate goods and raw materials

$$MRMR = \beta_{30} + \beta_{31} \frac{PMRM * EXR}{WPI} + \beta_{32} Y \qquad \textbf{(3.11)}$$

3.4.2.2 Imports of other goods

$$MOGR = \beta_{33} + \beta_{34} \frac{POMG * EXR}{WPI} + \beta_{35} Y \qquad \textbf{(3.12)}$$

3.5 Government Block

3.5.1 Revenue

Government revenue originates from tax and non-tax sources. Because of substantial dependence of tax on imports, taxes are divided into trade related taxes and internal taxes. Trade related taxes include customs duty, VAT and supplementary duty on imports and are related to imports. Internal taxes are specified as a function of nominal GDP. Similarly, non-tax revenue is made a function of nominal GDP. Though internal taxes and non-tax revenues are both functions of nominal GDP, two separate functions are estimated as they are of different nature having different degrees of response to changes in income.

3.5.1.1 Revenue from import based taxes

$$REVM = \beta_{36} + \beta_{37} M \qquad \textbf{(3.13)}$$

3.5.1.2 Revenue from internal taxes

$$REVIN = \beta_{38} + \beta_{39} NY \qquad \textbf{(3.14)}$$

3.5.1.3 Revenue from non-tax revenue

$$REVNT = \beta_{40} + \beta_{41} NY \qquad \textbf{(3.15)}$$

3.5.2 Government Expenditure

Government expenditure is divided into consumption expenditure and investment expenditure. Consumption expenditure includes both consumption expenditure and transfer payments. Expenditures depend only on nominal GDP. As mentioned before, investment expenditure is regarded as a policy variable.

3.5.2.1 Government consumption expenditure

$$GCE = \beta_{42} + \beta_{43} NY \qquad (3.16)$$

3.6. Monetary and Price Block

3.6.1 Supply of Money

The monetary system of Bangladesh consists of the Bangladesh Bank (the central bank) and the scheduled banks. They interact with the public and create money held by the public. The total money supply is specified as a simple function of government budget deficit financed by domestic borrowing and total credit to private sector. Government budget deficit provides an important mechanism for providing monetary base in Bangladesh.

3.6.1.1 Money supply

$$MS = \beta_{44} + \beta_{45} GDF + \beta_{46} CRT \qquad (3.17)$$

3.6.2 Price Level

Interest rate is regulated by the government. There is also rigidity in the interest rate because of the oligopolistic structure of the banking system. The equilibrating mechanism in the monetary sector does not work through the demand for and supply of money determining the rate of interest. Rather, the change in money supply affects the aggregate price level. Accordingly, instead of specifying demand for money function, a price level equation is specified which is derived from the simple money demand function. The price level is determined by money supply and income as measured by GDP.

3.6.2.1 Rate of inflation

$$RGP = \beta_{47} + \beta_{48} RGM + \beta_{49} RGY \qquad (3.18)$$

Before leaving this section it needs to be mentioned here that although the specifications are defined on the level of the variables, actual estimations in section 5 will be undertaken after taking the logarithmic transformation of the variables in various regression models. The most important advantage of log transformation is that the estimated parameters can be directly interpreted as the elasticities with respect to the relevant variables. Moreover, it has been suggested that such transformation often reduces the problem of heteroscedasticity by scaling down the variables. In macroeconometric modeling exercise the use of 'logged' variables has also become very common (e.g., see Mallick, 1999 and Musila, 2002).

Table 3.1: Description of Endogenous Variables

Variables	Description	Unit
CON	Private consumption expenditure in1995-96 price	Million taka
DT	Direct tax	Million taka
GCE	Government consumption expenditure in 1995-96 price	Million taka
GDF	Government budget deficit financed by domestic borrowing	Million taka
M	Total Imports	Million taka
INAGRP	Private investment in agriculture in 1995-96 prices	Million taka
INMANP	Private investment in manufacturing in 1995-96 prices	Million taka
INSERP	Private investment in services in 1995-96 prices	Million taka
KAGR	Capital Stock in agriculture in 1995-96 prices	Million taka
KMAN	Capital Stock in manufacturing activities in 1995-96 prices	Million taka
KSER	Capital stock in services in 1995-96 prices	Million taka
MRMR	Import of intermediate goods and raw materials in 1995-96 prices	Million taka
MS	Money supply calculated as average of stock of M2 on 30 June of current year and preceding financial year	Million taka
NY	Gross domestic product in current market price	Million taka
P	Implicit GDP Deflator (1995-96=1.00)	Index
REVIN	Revenue from domestic activity based taxes	Million taka
REVM	Revenue income from imports calculated as sum of import duty, supplementary duty on imports, and VAT on imports	Million taka
REVNT	Non-tax Revenue	Million taka
WPI	Wholesale price index	Index
XJMR	Real export of raw jute and jute manufactures	Million taka
XOGR	Real export of other goods which includes export of tea, leather, and fish	Million taka
XRMGR	Real export of readymade garments and knitwear	Million taka
Y	Gross Domestic Product in 1995-96 price	Million taka
YAGR	Value added in agriculture in 1995-96 prices	Million taka
YD	Disposable income in 1995-96 prices	Million taka
YMAN	Value added in manufacturing in 1995-96 prices	Million taka
YSER	Value added in services in 1995-96 prices	Million taka

Table 3.2: Description of Exogenous Variables

Variables	Description	Unit
RIRR	Ratio of irrigated area to total acreage	ratio
CRAGR	Private sector credit to agriculture in 1995-96 prices	Million taka
CRMAN	Private sector credit to manufacturing sector in 1995-96 prices	Million taka
CRSER	Private sector credit to services sector in 1995-96 prices	Million taka
CRT	Total credit to private sector	Million taka
EXR	Exchange rate	Taka per US$
GI	Total public investment in 1995-96 prices	Million taka
KAGRRES	Residual in agriculture capital stock identity	Million taka
KMANRES	Residual in manufacturing capital stock identity	Million taka
KSERRES	Residual in services capital stock identity	Million taka
PGI	Price deflator of public investment (1995-96=100)	Index
PMRM	Import price index of imported raw materials calculated as weighted index of dollar prices of raw materials, share of respective commodities being used as weights	Index
DRM	Quantity index of domestic raw materials calculated as weighted index of sugarcane, oilseeds, tobacco, cotton, jute, and wheat with share of value of production in 1995-96 used as weights	Index
PJ	Weighted index of price of raw jute and jute manufactures (1995-96=100)	Index
PMOG	Import price index of other imports	Index
PRMG	Export price of RMG	Unit Price (taka)
POG	Export price index of other goods (1995-96=100)	Index
RAIN	Annual average rainfall	Millimeters
JMC	Capacity utilization rate in the jute sector calculated as ratio of jute output to trend output of jute	Ratio
RMGC	Capacity utilization rate in the RMG sector calculated as ratio of RMG output to trend output of RMG	Ratio
OGC	Capacity utilization rate in the other goods sector calculated as ratio of other goods output to trend output of other goods	Ratio
GBO	Residual in the government budget deficit identity	Million taka
YDRES	Residual in disposal income identity	Million taka

3.7 Linkages in the Model

The model captures different linkages as they exist in the economy. While the specific relationships between the exogenous variables and outcomes may be seen in Figure 3.1, general features of these linkages are noted in the following:

(i) Production affects consumption expenditure, imports and thus balance on trade, government revenue and government consumption expenditure. Finally, it affects the price level in the economy.

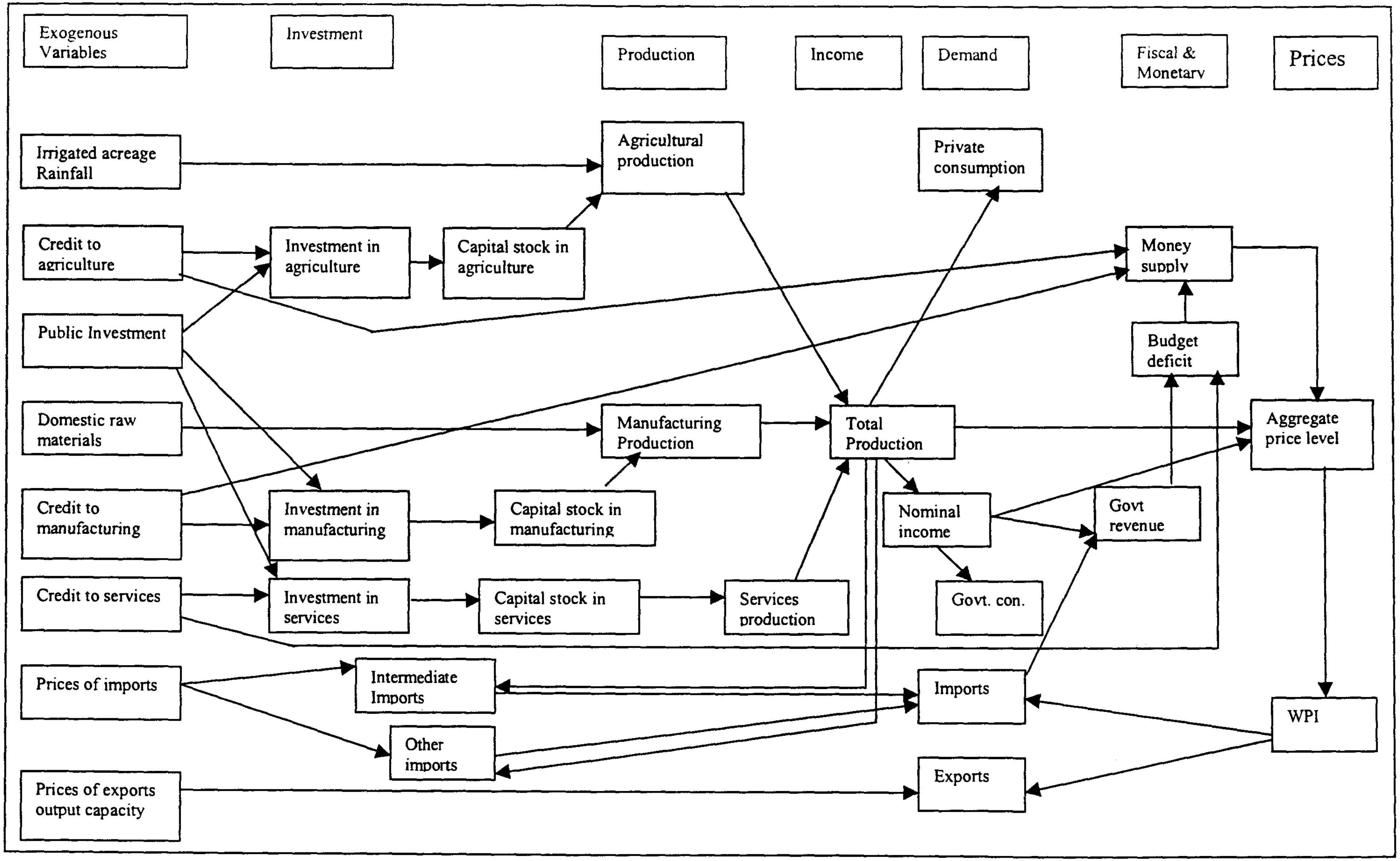
(ii) Banking sector credit to the private sector affects sectoral investment. Investment in turn affects the sectoral capital stock, which affects output.

(iii) Money supply is affected by private sector credit and government deficit. Money supply in turn influences the price level.

(iv) Price level is affected by both real sector and monetary sector variables. Wholesale price level influences exporter's incentive to export and importer's decision to import.

(v) Public investment influences private investment. which moves the economy through the linkages as mentioned earlier.

Figure 3.1: Linkages in the Model



4. Estimation Strategy

4.1. Time series properties of the variables

Recent developments in econometrics have emphasised a lot the characteristics of the time series data. Central to this is the distinction made between the stationary and non-stationary time series in contrast to the traditional practice of assuming all variables in the regression model to be stationary. A time series is said to be stationary if its mean, variance and auto-covariance are independent of time. By now there is compelling evidence that many macroeconomic time series are non-stationary in nature and as a consequence the ordinary least squares (OLS) regressions using these data might produce not only inconsistent and inefficient estimates but also spurious results. In other words, one could obtain a highly significant correlation between variables although in reality there may not exist any such relationship.[8] In order to avoid such problems of estimating non-sense relationship, the integrating properties of the variables should be examined carefully by testing for the existence of unit roots in variables under consideration.

The two most popular tests for unit roots, which we intend to use for the present study, are the Dickey Fuller (DF) and Augmented Dickey Fuller (ADF) tests. The DF test is based on equation (5.1) where Y is the variable under consideration, Δ is the first difference operator, subscript t denotes time period, T is the time trend and e is the error term. The null hypothesis for this test is that $(\psi - 1) = 0$ (i.e., Y_t is non-stationary) against the alternative of $(\psi - 1) < 0$ (i.e., Y_t is stationary). The 't' test on the estimated coefficient of Y_{t-1} provides the DF test for the presence of a unit root. The ADF test, on the other hand, is a modification of the DF test and involves augmenting equation (5.1) by lagged values of the dependent variables.[9] This is done to ensure that the error process in the estimating equation is residually uncorrelated.[10] More precisely, the ADF version of the test is based on the equation (5.2). As in the case of the DF test, the t-ratio on $(\psi - 1)$ provides the ADF test statistic.

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[8] One interesting example of spurious regression is due to Hendry (1980) who found a very strong positive relationship between inflation rate and the accumulated annual rainfall for the United Kingdom.

[9] Note that the DF and ADF tests are usually carried out with and without the time trend term (T) in the regression. If the variable is trended the insertion of the term is required. However, if the variable is not trended DF-ADF regressions can be applied without it.

[10] In the case of the annual data incorporation of the first lag of the dependent variable most often overcomes the problem of residual correlation. Higher order of lags would be necessary for quarterly and other high frequency data.

$$\Delta Y_t = \tau + (\psi - 1)Y_{t-1} + \chi T + e_t \quad (5.1)$$

$$\Delta Y_t = \tau + (\psi - 1)Y_{t-1} + \chi T + \delta \Delta Y_{t-1} + e_t \quad (5.2)$$

In both the equations (5.1) and (5.2) the estimated t-ratios on $(\psi - 1)$ are non-standard requiring the computed test statistics to be compared with the corresponding critical values to infer about the stationarity of the variables.[11] DF and ADF tests can, however, provide contrasting evidence and there appears to be a consensus in the literature that ADF test is preferable to DF. It is quite common to find that macroeconomic time series data are non-stationary on their levels but stationary on their first or higher order differences. Following Engle and Granger (1987) a time series is said to be integrated of order *d* [usually denoted as *~I(d)*] with *d* is the number of times the series needs to be differenced in order to become stationary.

It needs to be mentioned that in small sample the testing procedure for unit roots might be very complicated. Not only that the results emanating from different unit root test regressions can be inconclusive but also the critical values for such tests may prove to be very demanding. Further, it is well known that the low power of the DF and ADF tests is an unavoidable fact as Harris (1995) points out that the most important problem faced when applying the unit root test is their probable poor size and power properties.[12] This is often reflected in the tendency to over-reject the null when it is true and underreject the null when it is false. In a small sample the problem is likely to be even worse. Thus in the case of small sample Hall (1986) suggests the inspection of the autocorrelation function and correlogram as an important tool in determining whether the variables are stationary or not. The autocorrelation function for any variable at any lag *k* is defined by the ratio of covariance at lag *k* divided by variance.[13] When the estimated autocorrlation coefficients at different lags are plotted against *k*, population correlogram is

[11] These critical values were first computed by Dickey and Fuller (1981). If the computed test statistics exceed the critical values, the null hypotheses underlying the DF-ADF tests are rejected. Computed t-ratios and the corresponding critical values are compared on their absolute levels.

[12] Engle and Granger (1987) also highlighted the low power of the DF and ADF tests.

[13] The autocorrelation coefficient like any ordinary correlation coefficient lies between −1 and +1.

obtained.[14] For non-stationary variables correlograms die down slowly giving rise to either a secular declining or a constant trend in the graph of autocorrelation coeffieicnts while in the case of stationary variables they damp down almost instantly and then show random movement. For the present paper, therefore, we will employ the DF-ADF tests, autocorrelation coefficients and correlograms to determine the integrating properties of the variables.

4.2. Cointegration and Error Correction Modelling

4.2.1. The Engle Granger Procedure

Once it is determined that the variables in the model are non-stationary, the only way to infer about the long-run relationship is to employ some kind of cointegration technique. There are several cointegration methodologies in the literature – the simplest one being the Engle-Granger two step procedure. The basic idea behind the Engle-Granger technique is that if two variables say Y_t and X_t are both $\sim I(d)$, a linear combination of these two variables such that $V_t = X_t - \theta Y_t$, in general, will also be $\sim I(d)$. Engle and Granger, however, showed that in an exceptional case if the constant θ yields an outcome where $V_t \sim I(d-a)$ and $a>0$, then X_t and Y_t will be cointegrated. Usually the linear combination represented by the residuals from the OLS regression is tested for stationarity. Thus, if Y_t and X_t are both $\sim I(1)$, they will be cointegrated and have a valid long-run relationship if residuals from the OLS regression of X_t on Y_t is $\sim I(0)$. This is what is known as the first step of Engle-Granger procedure.

One important contribution of Engle and Granger (1987) was to find that if variables were cointegrated, there would have existed an error-correction model (ECM) of that cointegrating relationship. The ECM will then capture the short-run dynamics of the long-run behaviour, which is known as the second step of Engle-Granger procedure. The ECM is constructed by regressing the dependent variable in stationary form, onto its own lagged values and the current and lagged values of the stationary forms of the dependent variables, and the lagged error term

[14] Note that in practice we only have a realisation of a stochastic process and therefore can only compute sample autocorrelation function, which is defined as: $\frac{\sum (Y_t - \bar{Y})(Y_{t+k} - \bar{Y})}{\sum (Y_t - \bar{Y})^2}$

from the cointegrating relationship. If we assume that both Y_t and X_t are ~ $I(1)$ such that ΔY_t and ΔX_t are ~ $I(0)$, the ECM can be represented as:

$$\Delta Y_t = \pi_0 + \sum_{i=0}^{m} \pi_{1i} \Delta X_t + \sum_{i=1}^{n} \pi_{2i} \Delta Y_t + \pi_3 \hat{\upsilon}_{t-1} + \varepsilon_t \tag{5.3}$$

Equation (5.3) gives a very general representation of the ECM. Since all variables in (5.3) are ~ $I(0)$, the problem of spurious regression is overcome. It is worth noting that the ECM is not a mere regression of the stationary variable rather it includes $\hat{\upsilon}_{t-1}$, the deviation from the steady-state long-run path, which basically contains the long-run information. Thus the ECM captures the short-run relationship taking into consideration the long-run information. A valid representation of the ECM will require $0 > \pi_3 \geq -1$. The usual practice with the error correction modelling is to follow the "general to specific" methodology by constructing a general model in the beginning and subsequently reduce it to a parsimonious form after dropping all the insignificant variables step-by-step.

4.2.2. The Phillips-Hansen Fully Modified OLS

In estimating the equations as specified in Section 4 we can employ the Engle-Granger cointegration procedure to test for a valid long-run relationship. However, although this procedure can test for cointegration, it yields standard errors that do not provide the basis for valid inferences. In equations with more than two explanatory variables this can be problematic in the sense that even if the variables are found to be cointegrated we cannot be certain whether any particular explanatory variable is significant or not.[15] We propose to handle this problem by using the Phillips-Hansen Fully Modified OLS (PHFMOLS) technique (Phillips and Hansen, 1990). The Phillips-Hansen method is an optimal single-equation technique, which is asymptotically equivalent to maximum likelihood procedure. It makes a semi-parametric correction to the OLS estimator to eliminate dependency of the nuisance parameters and

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[15] That is, for example, in a three variable, say Y, X and Z, regression model cointegration does not necessarily suggest statistically significant influence of both the explanatory variables, X and Z. It might be that only X is significant but not Z and vice-versa. Since the computed standard errors in the first step of the Engle-Granger procedure is not valid, correct statistical inference from the estimated model is not possible.

provides standard errors that follow standard normal distribution asymptotically and thus are valid for drawing inferences. Due this particular advantage the use of PHFMOLS has become quite popular in international trade and macroeconomic modeling.[16] The PHFMOLS procedure can be described by the following.[17]

Consider the data generating mechanism for Y_t following the cointegration system:

$$Y_{1t} = \alpha_0 + \alpha_1 + \beta' Y_{2t} + u_{1t} = \lambda R_t + u_{1t} \tag{5.4}$$

$$\Delta Y_{2t} = u_{2t} \tag{5.5}$$

$$u_t = \begin{pmatrix} u_1 \\ u_2 \end{pmatrix} = \psi(L)\varepsilon_t, \qquad E(\varepsilon_t \varepsilon_t') = PP' \tag{5.6}$$

where, Y_{1t} and Y_{2t} are scalar and $m \; x \; t$ vector of $I(1)$ stochastic processes, $\lambda' = (\alpha_0 + \alpha_1 + \beta')$ and $R_t = (Y_{1t}, Y_{2t})$. We define:

$$\Omega = \psi(1)P, \qquad \Sigma = \Omega\Omega' = \begin{pmatrix} \Sigma_{11} & \Sigma_{12}' \\ \Sigma_{21} & \Sigma_{22} \end{pmatrix} \tag{5.7}$$

Σ is the long-run covariance matrix of u_t. As mentioned earlier, the PHFMOLS estimator is an optimal single equation method based on the use of OLS on equation (5.4) with semi-parametric corrections for serial correlation and potential endogeneity of the right hand side variables.

Consider the OLS estimator of the cointegrating equation (5.4) by $\hat{\lambda} = (R_t' R_t)^{-1} R_t' Y_{1t}$, where R_t and Y_{1t} are respectively $Tx(m+2)$ and $Tx1$ matrices of observations on R_t and Y_{1t}. Due to serial correlation in u_{1t} and endogeneity of Y_{1t} $\hat{\lambda}$, in general, is consistent but biased. The FM procedure modifies the OLS estimator $\hat{\lambda}$ to correct for serial correlation and endogeneity bias. The FM estimator is given by:

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[16] Amongst others, Athukorala and Riedel (1995) and (1996), Muscatelli, et al. (1992), Senhadji (1998) and Senhadji and Montenegro (1998) have used the Phillips-Hansen procedure to modeling trade for various countries, while Mallick (1999) has applied the procedure to macroeconometric modelling for India.

[17] This is based on the illustrations in Senhadji (1998).

$$\hat{\lambda}^{*} = \begin{pmatrix} \hat{\alpha}^{*}_{0} \\ \hat{\alpha}^{*}_{1} \\ \hat{\beta}^{*} \end{pmatrix} = \begin{pmatrix} T & \sum_{t=1}^{T} Y'_{2t} & \sum_{t=1}^{T} t \\ \sum_{t=1}^{T} t & \sum_{t=1}^{T} tY'_{2t} & \sum_{t=1}^{T} t^{2} \\ \sum_{t=1}^{T} Y_{2t} & \sum_{t=1}^{T} Y_{2t}Y'_{2t} & \sum_{t=1}^{T} Y_{2t}t \end{pmatrix}^{-1} \begin{pmatrix} \sum_{t=1}^{T} \hat{Y}_{1t}^{+} \\ \sum_{t=1}^{T} tY_{1t}^{+} \\ \sum_{t=1}^{T} Y_{2t}Y_{1t}^{+} - T\hat{V}_{T}^{+} \end{pmatrix} \tag{5.8}$$

$$\hat{Y}_{1t}^{+} = Y_{1t} - \hat{\Sigma}_{21}{}' \hat{\Sigma}_{22}^{-1} \Delta Y_{2t} \tag{5.9}$$

$$\hat{\Sigma} = \begin{pmatrix} \hat{\Sigma}_{11} & \hat{\Sigma}_{21}{}' \\ \hat{\Sigma}_{21} & \hat{\Sigma}_{22} \end{pmatrix} = \hat{\Gamma}_{0} + \sum_{v}^{q} \left(1 - \frac{v}{q+1}\right)\left(\hat{\Gamma}^{v} + (\hat{\Gamma}^{v} + (\hat{\Gamma}^{v})')\right) \tag{5.10}$$

$$\hat{\Gamma}_{v} = T^{-1} \sum_{t=v+1}^{T} \begin{pmatrix} \hat{u}_{1t}\hat{u}_{1t-v} & \hat{u}_{1t}\hat{u}_{2t-v}{}' \\ \hat{u}_{1t-v}\hat{u}_{2t} & \hat{u}_{2t}\hat{u}_{2t-v} \end{pmatrix} = \begin{pmatrix} \hat{\Gamma}_{11}^{v} & \hat{\Gamma}_{12}^{v} \\ \hat{\Gamma}_{21}^{v} & \hat{\Gamma}_{22}^{v} \end{pmatrix} \tag{5.11}$$

$$\hat{V}_{T}^{+} = \sum (1 - \frac{v}{q+1})[(\hat{\Gamma}_{12}^{v})' + (\hat{\Gamma}_{22}^{v})'] \tag{5.12}$$

Where q is the bandwidth parameter in the Bartlett window used in the estimation of the long-run covariance matrix. The difference between the OLS and FM estimators is highlighted in the last vector of (5.8) where Y_{1t} is replaced by $\hat{Y}_{1t}^{+}$ (which corrects for the potential endogeneity of Y_{2t}) and the factor TV_{T}^{+} (which corrects for the potential autocorrelation of the error term). The FM estimator $\hat{\lambda}^{*}$ has the same asymptotic behaviour as the full information system maximum likelihood estimators.

4.2.3. Existence of a Long-run Relationship

4.2.3.1. Testing for Cointegration

From the above our estimation strategy can be summarized as follows. First, the time series properties of the variables will be analysed and in the case of equations containing non-stationary variables PHFMOLS method will be used, which would provide standard errors for valid inferences. The estimation by PHFMOLS itself does not guarantee cointegration needing one to check for residual stationarity. In the literature the standard practice of testing for cointegration has been the use of ADF test, which is given in equation (5.13). Note that in contrast to the regular ADF regressions, the test for residual does not include any intercept term.[18]

$$\Delta\hat{v}_t = \rho\hat{v}_{t-1} + \kappa\Delta\hat{v}_{t-1} + \tau \qquad (5.13)$$

The null hypothesis for the test is that $\rho=1$ (non-cointegration) against the alternative of $\rho<1$ (cointegration). Like the regular ADF test statistics the estimated standard errors in (5.13) are non-standard and hence they will have to be compared with the appropriate critical values as estimated by Engle and Granger (1987) and Mackinnon (1991).[19]

Despite its widespread use, the low power of the ADF test is considered to be a serious shortcoming for cointegration test. Engle and Granger showed that when $\rho=0.9$ the ADF test for cointegration has about 28 per cent chance of not rejecting the null of no cointegration even when it is false.[20] In small sample, testing for cointegration is more troublesome as apart from the low power critical values for such tests become more demanding. One effective way of tackling this problem will be to follow Hall (1986) and examine the autocorrelation coefficients and the resultant correlograms of the estimated error term from the static long-run equations.

4.2.3.2 Testing Cointegration for Variables With Different Orders of Integration: The Pesaran et al. Test

It is very well possible to have an equation with variables of different integrated orders – for example, a mixture of *I(1)* and *I(0)* variables. Then the question is whether the *~I(0)* regressors play a role in determining the *I(1)* variable. In a study Holden and Perman (1994) have

[18] This is because by definition the mean of the residual is zero.

[19] These days many econometric software routinely compute such critical values.

considered a model with two *I(1)* and one *I(0)* variables. The authors used the Johansen rank cointegration procedure to determine a valid long-run relationship between the two *I(1)* variables and then included the *I(0)* variable only in the short-run error-correction model. The procedure, thus, assumes that the *I(0)* variable does not have any role to play in the long-run disregarding the economic theory behind it. On the other hand, Pesaran *et al.* (2001) observe that "the strict precondition for the same order of integration of the variables in a model involves a certain degree of pre-testing, thus introducing a further degree of uncertainty into the analysis of a long run relationship". They have strongly argued that the fact that the variables in the estimating equation have different orders of integration does not necessarily mean that they are unlikely to have any long-run impact.[21] Pesaran, *et al.* (2001) have also devised a strategy, which tests the existence of a long-run relationship when the variables are a mixture of *I(0)* and *I(1)*. This procedure is based on an OLS estimation of unrestricted error correction model, a general specification of which with respect to our model can be written as:

$$\Delta \ln X_t = \alpha + \gamma \ln X_{t-1} + \xi Z_{t-1} + \sum_{i=1}^{p} \pi_i \Delta \ln X_{t-1} + \sum_{i=0}^{g} \delta_i \Delta Z_{t-i} + \varepsilon_i \qquad (5.14)$$

Estimation of (11) in itself is not interesting since the existence of a long-run relationship can only be tested by examining the joint null hypothesis that $\gamma = \xi = 0$ with the help of either a *Wald* or an *F* test. The presence of a long run relationship requires the rejection of this null. However, as the asymptotic distribution of these statistics is non-standard, Pesaran *et al.* provide the necessary critical upper (F_U) and lower (F_L) bound for the *F* test.[22] The F_U are derived under the assumption that all variables are *I(1)* and the F_L considers all of them to be *I(0)*. If the computed F statistic (*F*), which is obtained by restricting that $\gamma = \xi = 0$, is greater than the critical upper value, i.e. $F > F_U$, we reject the null and conclude that there is a valid long-run relationship among the variables. If $F < F_L$ then no long-run relationship exists, and finally, if $F_L < F < F_U$ the test is inconclusive. Pesaran *et al.* clearly point out that "[I]f the computed *Wald* or *F*-statistic falls outside the critical value bounds a conclusive inference can be drawn without needing to know the integration/cointegration status of the underlying regressors."[23]

[20] Razzaque and Ahmed (2000) illustrate a case when the ADF test on the residuals falls into the trap of its low power.
[21] See, Pesaran *et al.* (1999), particularly page 19, first paragraph.
[22] Pesaran *et al.* give both the critical values for Wald and F Statistics. In this paper we will only consider the F-statistics.
[23] Pesaran *et al.* (2001), p.1.

4.2.4. Reason for Employing Single Equation Estimation Technique

A problem with the Engle-Granger and PHFMOLS is that they ignore the possibility of multiple cointegrating vectors. This problem can be tackled by Johansen's (1988) Full Information Maximum Likelihood (FIML) procedure. However, there are two important problems associated with this approach. First of all, the results from the Johansen procedure can be very sensitive to the choice of lag-length (Hall, 1991; Banerjee *et al.* 1993). Although there are statistical tests for choosing the appropriate lag-lengths, in a small sample such tests may not be feasible. Moreover, severe problem of collinearity among the regressors may also arise when a considerable size of VAR is used (Athukorala and Riedel, 1996). As in the present study we will be dealing with a small sample size (annual observations for 20 years), the Johansen procedure may not be an appropriate one. Therefore, we will have to rely on the single equation procedures, such as those of Engle-Granger and PHFMOLS.

4.2.5. Short-run Dynamics With Partial Adjustment Modelling Technique

There might be situations when the two-step Engle-Granger procedure might be problematic. This situation emerges when the first step simple estimation involves serially correlated error, which even cannot be rectified through the process of PHFMOLS. The problem would require incorporating the lag dependent variable on the right hand side of the equation.[24] Under such circumstances the use of partial adjustment modelling (PAM) technique can be conveniently made providing for short-run and long-run coefficients based on a single regression equation. Before the advent of cointegration technique, the use of PAM was very popular and most macroeconometric models had used them regularly.[25] However, this is important to remember that the use of OLS in estimating PAM is problematic for the same reasons as explained before. Consequently, in the present study the partial adjustment models will be estimated by PHFMOLS.

[24] Note that the Engle-Granger procedure usually does not include the lag dependent variable among the set of explanatory variables.
[25] For example, see Rahman and Shilpi (1996) and Rashid (1981).

5. Estimation Results

5.1. Production Block

5.1.1. Order of Integration of Variables

We begin our estimation with the models in production block as given in equations (3.1)-(3.3). As a first step toward the estimation, all variables in these three equations are tested to determine whether they can be represented as stationary or non-stationary processes by employing the unit root tests and examining the correlograms and autocorrelation functions. Table 5.1 gives the results of DF and ADF tests on level and first difference of all 10 variables in equations (3.1)-(3.3), while Figure 5.1 presents their graphical plots as well as correlograms. In most cases the results of DF-ADF tests appear to be inconclusive requiring us to determine the integrating properties of the variables carefully.

Table 5.1: Unit Root Test For Variables in the Production Block

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
lnYAGR	-0.37	-0.34	-2.44	-2.90
ΔlnYAGR	-3.88	-3.21	-3.74	-3.07
lnKAGR	-3.62	-3.33	-3.43	-3.22
ΔlnKAGR	-2.11	-1.69	-1.70	-1.06
lnRIRR	-1.25	-1.17	-1.35	-1.72
ΔlnRIRR	-3.42	-3.52	-3.46	-3.72
lnRAIN	-5.42	-2.77	-5.80	-3.13
ΔlnRAIN	-9.18	-6.60	-8.85	-6.43
lnYMAN	0.16	-0.07	-1.77	-1.98
ΔlnYMAN	-3.72	-3.33	-3.53	-3.01
lnKMAN	0.58	0.91	-2.17	-1.87
ΔlnKMAN	-4.89	-1.67	-5.20	-1.69
lnMRMR	-1.10	-0.77	-3.04	-2.58
ΔlnMRMR	-5.12	-3.63	-4.92	-3.88
lnDRM	-2.48	-2.62	-2.88	-3.07
ΔlnDRM	-4.04	-3.65	-3.91	-3.92
lnYSER	4.09	2.61	0.26	0.21
ΔlnYSER	-2.05	-1.12	-3.63	-2.40
lnKSER	5.59	2.57	-0.99	-0.96
Δ lnKSER	-2.69	-1.94	-4.35	-2.13

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69. The prefix *ln* represents natural logarithmic transformation of the variables while Δ denotes their first difference.

Figure 5.1: Plot of variables on their levels and first differences and Correlograms

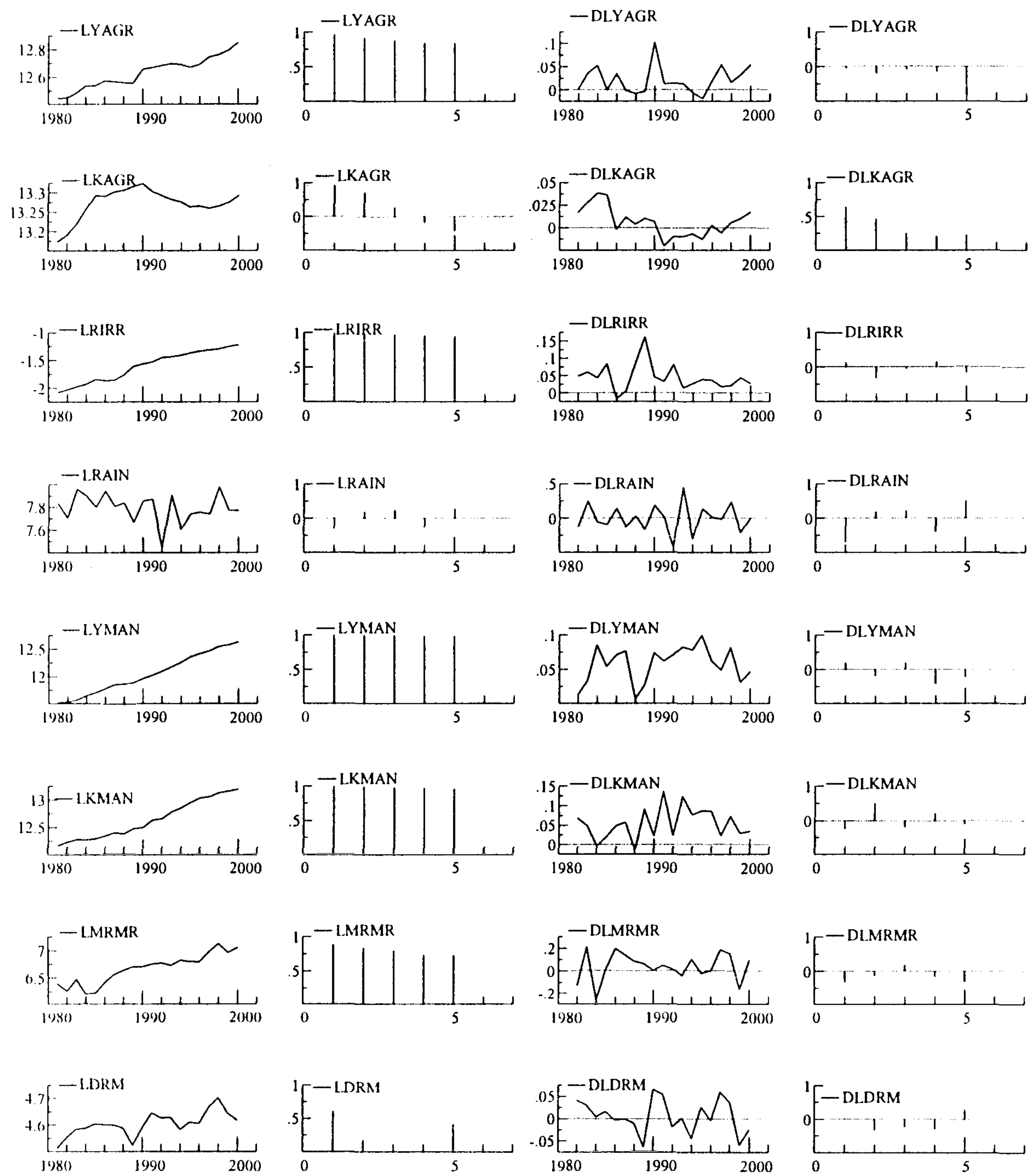
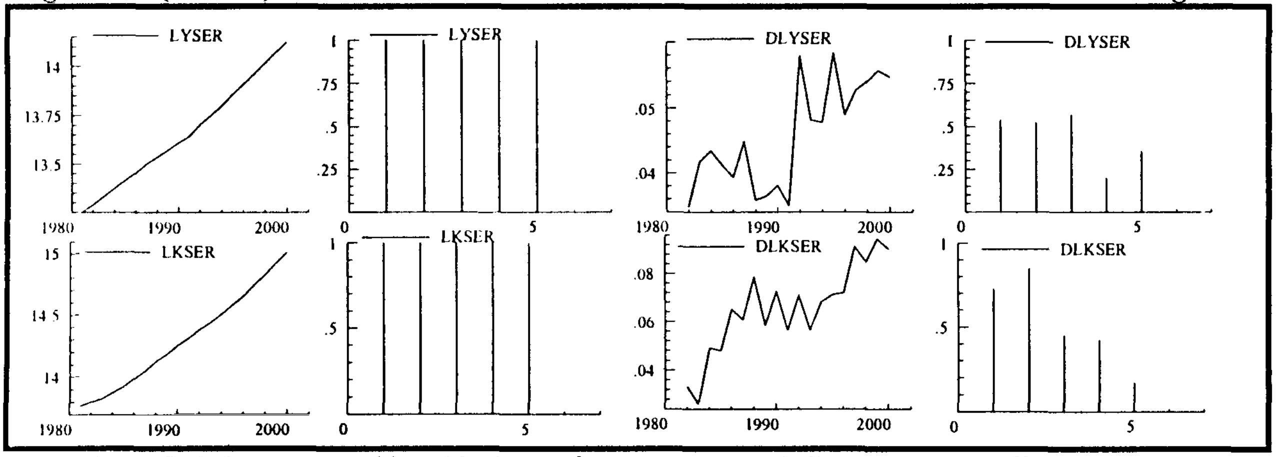


Figure 5.1 (Contd.): Plot of variables on their levels and first differences and Correlograms



Note: The prefix L denotes natural logarithmic transformation, while D indicates their first difference.

First, in three cases the unit root test results are unambiguous. For lnRIRR, lnMRMR, and lnDRM all the computed DF and ADF statistics are found to be lower than their critical values while all computed test statistics for ΔlnRIRR, ΔlnMRMR and ΔlnDRM are smaller than the critical values. These results imply that the null hypothesis of non-stationary cannot be rejected on the levels of the variables but the same hypothesis can be rejected on their first differences. That is, lnRIRR, lnMRMR, and lnDRM are *~I(1)*. The correlograms of lnRIRR and lnMRMR hardly show any tendency of damping down while those of ΔlnRIRR and ΔlnMRMR die down instantly and then show random movement just like any stationary variable thereby supporting the conclusion of the DF-ADF tests. On the other hand, the correlograms of lnDRM although damps down but remains at a considerably higher level at the first lag and unlike a stationary variable all autocorrelation coefficients are found to be positive. The correlogram of ΔlnDRM, however, behaves more like a stationary variable.

In the case of the lnYAGR, all DF and ADF tests suggest its non-stationarity but on ΔlnYAGR the ADF test with the trend term contradicts the results of other DF-ADF tests. However, as the first difference of lnYAGR is not trended (see Figure 5.1), considering the ADF test with trend term it can be inferred that ΔlnYAGR is *~I(0)*. The correlograms of lnYAGR and ΔlnYAGR also seem to behave like an *~I(1)* and an *~I(0)* variables respectively. Turning to the lnRAIN, apart from the two DF tests on the levels, all tests indicate non-stationarity of the level variable and the stationarity of the first difference of the corresponding variable. Since the ADF test is preferable to DF tests, lnRAIN can be considered as an *~I(1)* variable. The test results for lnYMAN are inconclusive due to the ADF test with the trend term on ΔlnYMAN, which is at variance with the same test without the trend term. The graphical plot of ΔlnYMAN is clearly not-trended and hence it should be treated as a stationary variable in which case lnYMAN is *~I(1)*.

It appears that three variables possess the most problematic time series properties. These are: lnKAGR, lnKMAN, lnYSER, and lnKSER. The ADF tests cannot reject the non-stationary hypothesis either for lnKMAN, lnYSER, and lnKSER or for their corresponding first differenced variables. In the case of lnKAGR although the ADF test without the trend term points toward the stationarity of the level variable, the same test with the trend provides contrasting evidence. Since lnKAGR is trended, it would be more appropriate to consider the result of the ADF regression with the trend term. Besides all the DF-ADF tests on ΔlnKAGR overwhelmingly suggest its non-stationarity. Among these four problematic variables, correlograms for lnKMAN differ strikingly from those of others. In fact, the correlograms of lnKMAN and ΔlnKMAN are just like any other ordinary *~I(1)* and *~I(0)* variables respectively. On the other hand, correlograms for ΔlnKAGR, ΔlnYSER and ΔlnKSER in no way show random movement and, therefore, appear not to represent stationary variables. Second difference transformation of lnKAGR, lnYSER, and lnKSER also failed to transform the variables into stationary series.[26]

To sum up, for the agricultural value added model lnYAGR, lnRIRR, and lnRAIN are *~I(1)* while the integrating order of lnKAGR is not clear. In the manufacturing value added model all variables namely lnYMAN, lnKMAN, lnMRMR, lnDRM are *~I(1)*, while for the service sectoral value added model the integrating properties of the two variables, lnYSER and lnKSER are not clear.

5.1.2. Estimates for agricultural value added model

The OLS estimate of equation (3.1) is given in the top row of Table 5.2, where it is found that the capital elasticity of agricultural value added is 0.26, while the corresponding elasticities for

[26] These variables might have been caught into the trap of the low power of DF-ADF tests. They might also be subject of structural breaks, in which case usual unit root test results are not valid. Small sample size of ours do not permit rigorous structural break testing procedures.

RIRR and RAIN are respectively 0.38 and 0.067. Due to the unusually high magnitude of residual for 1989 the model is estimated with a dummy variable, which assumes a value of 0 for 1989 and 1 for all other years.[27] Since in the presence of non-stationary variables standard errors do not provide for valid inferences, they are not reported. The value of adjusted R^2 and the Durbin-Watson statistics are, however, reported to show that the estimated relationship might not be a spurious one as the value of R^2 is greater than the Durbin-Watson statistic.[28]

Table 5.2: Estimated Long-run Agricultural Value Added Model

OLS estimate

lnYAGR = 9.20 + 0.26 lnKAGR + 0.38 lnRIRR + 0.067 lnRAIN + 0.08 D89

Adjusted R^2 = 0.94, D.W. = 1.15

PHFMOLS estimate

lnYAGR =	6.84^{***}	+ 0.43^{**} lnKAGR	+ 0.39^{***} lnRIRR	+ 0.07^{*} lnRAIN	+ 0.08^{***} D89
(s.e.)	(2.18)	(0.16)	(0.019)	(0.04)	(0.022)
t-ratio	3.14	2.65	19.93	1.86	3.37

Note: Statistical significance at the one, five and 10 per cent levels are indicated by ***, **, and * respectively.

In order to make statistically valid inferences the same model is estimated by the PHFMOLS method – results of which are given in the bottom row of Table 5.2. It is now observed that all variables in the model are significant at least at the ten per cent level. A one unit increase in capital stock is associated with 0.43 unit rise in agricultural output, which is considerably greater than the similar coefficient estimated by the simple OLS regression reported above. The effects of irrigated area and weather (proxied by rain) change a little with the use of PHFMOLS.

Examination of the variables in the model of agricultural value added could not confirm first order integration of lnKAGR like the three other variables. However, if residuals from the long-run relationship turn out to be stationary, a valid long-run relationship can be ascertained.[29] The ADF test on the residuals obtained from the PHFMOLS estimates of the long-run relationship failed to reject the null hypothesis of non-stationarity. However, in light of the low power of such tests and demanding nature of critical values in short sample further examinations were carried

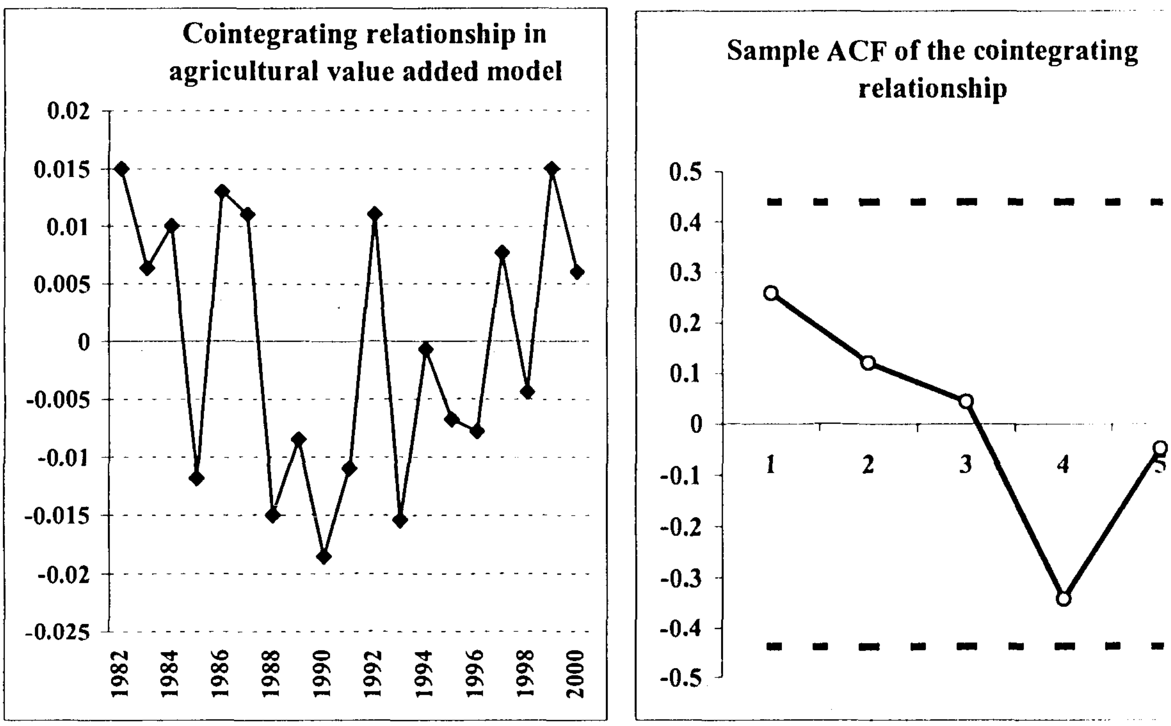
[27] It may be mentioned here that a devastating flood caused large damage to agriculture resulting in less output in 1989.

[28] Nevertheless, we cannot be certain about the statistical significance of individual variables.

[29] If lnKAGR is *~I(2)* (or of higher order), a linear combination of the variables is most likely to be *~I(2)* (or of higher order) in which case even *I(1)* residual will be sufficient to support cointegration result. On the other hand, if the residual is found to be *~I(0)*, it can be inferred that lnKAGR is either *~I(1)* or *~I(0)*.

out to infer about the long-run relationship. First, the graphical plot of the residuals (RESAGR), as given in Figure 5.1, seems to demonstrate a pattern similar to stationary variables. If a series is stationary, sample autocorrelation coefficients are approximately normally distributed with zero mean and standard error $1/\sqrt{n}$, where n is the sample size. Since we have a sample of 20 observations, it will imply a standard error of 0.2236. Now, following the properties of the standard normal distribution, the 95 per cent confidence interval for any of the sample autocorrelation coefficient will be $\pm\ 1.96(0.2236) = \pm\ 0.4382$. If all observations fall inside this interval, as in figure 5.2, the null hypothesis of true autocorrelation coefficient being equal to zero cannot be rejected. Besides, one can test the joint hypothesis that all the autocorrelation coefficients are simultaneously equal to zero by employing the Box-Pierce and Ljung-Box statistics.[30] Both these statistics could not reject null hypothesis thereby providing support for the stationarity of RESAGR.

Figure 5.2: Cointegrating relationship in Agricultural Value Added Model and Sample ACF



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[30] The Box-Pierce Statistic (known as Q statistic) is given by $Q = n\sum_{k=1}^{m} \hat{p}_k^{\,2}$, where n is the sample size, m is the lag length and $\hat{p}$ is the sample autocorrelation coefficient. On the other hand, the Ljung-Box statistic (LB) is derived as: $LB = n(n+2)\sum_{k=1}^{m}\left(\frac{\hat{p}_k^{\,2}}{n-k}\right)$.

The estimated long-run relationship allows modeling the corresponding short-run dynamic adjustments using the error-correction mechanism. Following the Engle and Granger procedure the error-correction model incorporates the lagged residual ($RESAGR_{t-1}$) from the estimated long-run equation, which captures the time required to converge to the steady state relationship from any short-run deviations. With the error-correction models the common practice is to adopt the 'general to specific' modelling strategy of specifying a very general model in the beginning by including the first difference of the variables along with their first or higher order lags and subsequently deleting the insignificant variables to arrive at the most parsimonious representation.

Table 5.3 gives the resultant parsimonious short-run model.[31] It is observed that in the short-run capital and acreage under irrigation and the 1989 dummy are highly significant while as a result of its insignificance weather had been dropped from the most parsimonious model. Both the coefficients on lnKAGR and lnRIRR appear to be considerably bigger than the corresponding long-run model. The error-correction term, $RESAGR_{t-1}$, is correctly signed and significant at the five per cent level indicating a valid representation of the long-run model. The coefficient suggests that it takes about two years to correct all short-run disequilibrium errors. 80 per cent variation in the growth of the real agricultural output can be explained by right-hand side explanatory variables. For diagnostics Godfrey's (1978) LM test for serial correlation, Ramsey's (1969) RESET test for functional form, White's (1980) test for heteroscedasticity and Jarque-Bera (1987) test for normality of errors are performed. The computed test statistics for serial correlation, functional form and heteroscedasticity follow a chi-square distribution with 1 degree of freedom, while the normality test statistic has a chi-square distribution with 2 degrees of freedom. Since the 95 per cent critical values for $\chi^2(1)$ and $\chi^2(2)$ are 3.84 and 5.99 respectively, on the basis of the computed diagnostic statistics we cannot reject any of the null hypotheses of no problem of serial correlation, no wrong functional form problem, normality of residuals and homoscedastic distribution of errors.

[31] Due to small sample size only the first lag of the first differenced variables were tried in the general mode. None of these variables were found to be statistically significant justifying their deletion.

Table 5.3:Short-run Error Correction Model for Value Added in Agriculture

$\Delta\ln YAGR =$	-0.04^{***}	$+ 0.812^{***}\ \Delta\ln KAGR$	$+ 0.511^{***}\ \Delta\ln RIRR_{t-1}$	$+ 0.04^{***}\ D89$	$- 0.453^{**}\ RESAGR_{t-1}$
(s.e.)	(0.01)	(0.23)	(0.08)	(0.01)	(0.16)
t-ratio	-4.5	3.48	6.07	4.63	-2.70

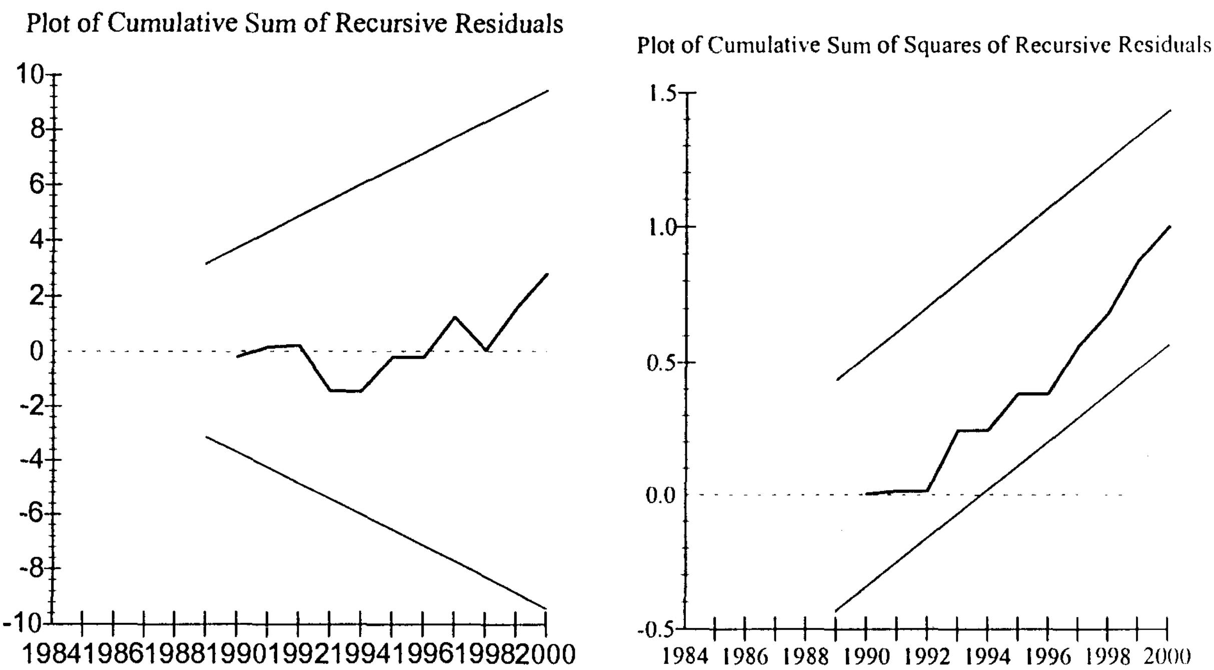
Diagnostic Tests

Adjusted $R^2 = 0.80$

Serial Correlation: $\chi^2(1) = 0.96$	Functional Form: $\chi^2(1) = 0.56$
Normality: $\chi^2(2) = 0.12$	Heteroscedasticity: $\chi^2(1) = 0.07$

Note: ***, and ** are for statistical significance at the one and five percent levels, respectively. The serial correlation test is based on Godfrey's (1978) LM test for serial correlation; Functional Form on Ramsey's (1969) RESET test; Heteroscedasticity on White's (1980) test; and Normality of residuals on Jarque-Bera (1987) test. The computed test statistics for serial correlation, functional form and heteroscedasticity follow a chi-square distribution with one degree of freedom while normality test statistic follows a chi-square distribution with 2 degrees of freedom.

Figure 5.3: Plot of Recursive Residuals



Note: The straight lines represent critical bounds at the five per cent level.

Figure 5.3 gives the plots of cumulative sum of recursive residuals and the cumulative sum of squares of recursive residuals along with their critical bounds at the five per cent level. As both the cumulative recursive residuals fall well within the confidence interval, we can be certain that the estimated short-run model does not have any problem of structural break.

5.1.3. Estimates for manufacturing value added model

We recall that value added in manufacturing sector is explained by the capital stock in the sector, imported intermediate and raw materials and domestically produced raw materials – all of which were found to be $\sim I(1)$. Therefore, a valid long-run relationship would result in an $I(0)$ cointegrating relationship. The PHFMOLS estimate of the long-run relationship yielded the following:

lnYMAN =	-1.69 +	0.90^{***} ln KMAN +	0.15^{*} lnMRMR +	0.29 lnDRM
(s.e.)	(1.18)	(0.06)	(0.079)	(0.30)
t-ratio	-1.43	13.72	1.90	0.96

Thus, the capital stock, imported intermediate goods and domestic raw materials elasticites are respectively 0.90, 0.15 and 0.29. While KMAN and MRMR come out to be significant at the one and 10 per cent level respectively, DRM failed to achieve statistical significance.[32] There is some strong evidence for cointegration in the estimated relationship, as the ADF test of residual stationarity was estimated at 4.10 against its 95 per cent critical value of 3.74. Figure 5.4 provides the graphical plots of the cointegrating relationship in the manufacturing value added model and its sample autocorrelation functions. Since all values of sample ACF fall within the critical bands, the residuals should be considered as a stationary series.

Table 5.4 is the short-run error correction model corresponding to the long-run manufacturing value added relationship as estimated above. In the short-run the capital stock and the domestic raw materials are found to be influencing the manufacturing value added significantly. Due to insignificance of imported raw materials (MRMR), we deleted it from the parsimonious model. The short-run capital stock elasticity is 0.31 – much lower than its long-run counterpart, while the domestic raw material elasticity is 0.24. The coefficient on the error correction term, $RESMAN_{t-1}$, is correctly signed and significant although the magnitude of the coefficient is quite low suggesting that it takes about almost 3 years to correct all disequilibrium errors. The short-run model explains 47 per cent variation in the growth of manufacturing value added and the diagnostic tests do not indicate any misspecification problem.

[32] Deletion of lnDRM would not have changed the coefficients and statistical significance of other variables much.

Figure 5.4: Cointegrating vector from the manufacturing value added model and sample ACF

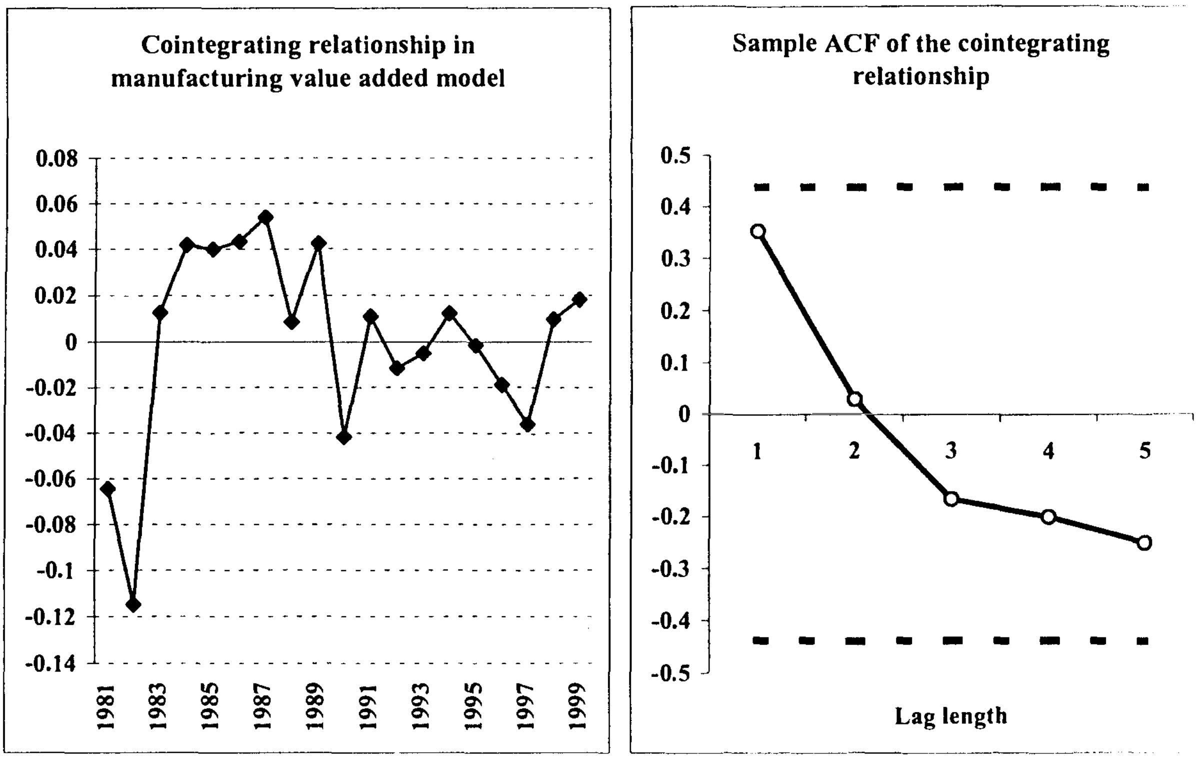


Table 5.4 : Short-run Error Correction Model for Value Added in Manufacturing

$\Delta \ln YMAN =$	-0.045^{***}	$+ 0.31^{**} \Delta \ln KMAN$	$+ 0.24^{**} \Delta \ln DRM$	$- 0.35^{**} RESMAN_{t-1}$
(s.e.)	(0.007)	(0.11)	(0.11)	(0.11)
t-ratio	6.30	2.83	2.07	-3.04

Diagnostic Tests

Adjusted $R^2 = 0.47$

Serial Correlation: $\chi^2(1) = 0.18$	Functional Form: $\chi^2(1) = 2.17$
Normality: $\chi^2(2) = 0.64$	Heteroscedasticity: $\chi^2(1) = 0.59$

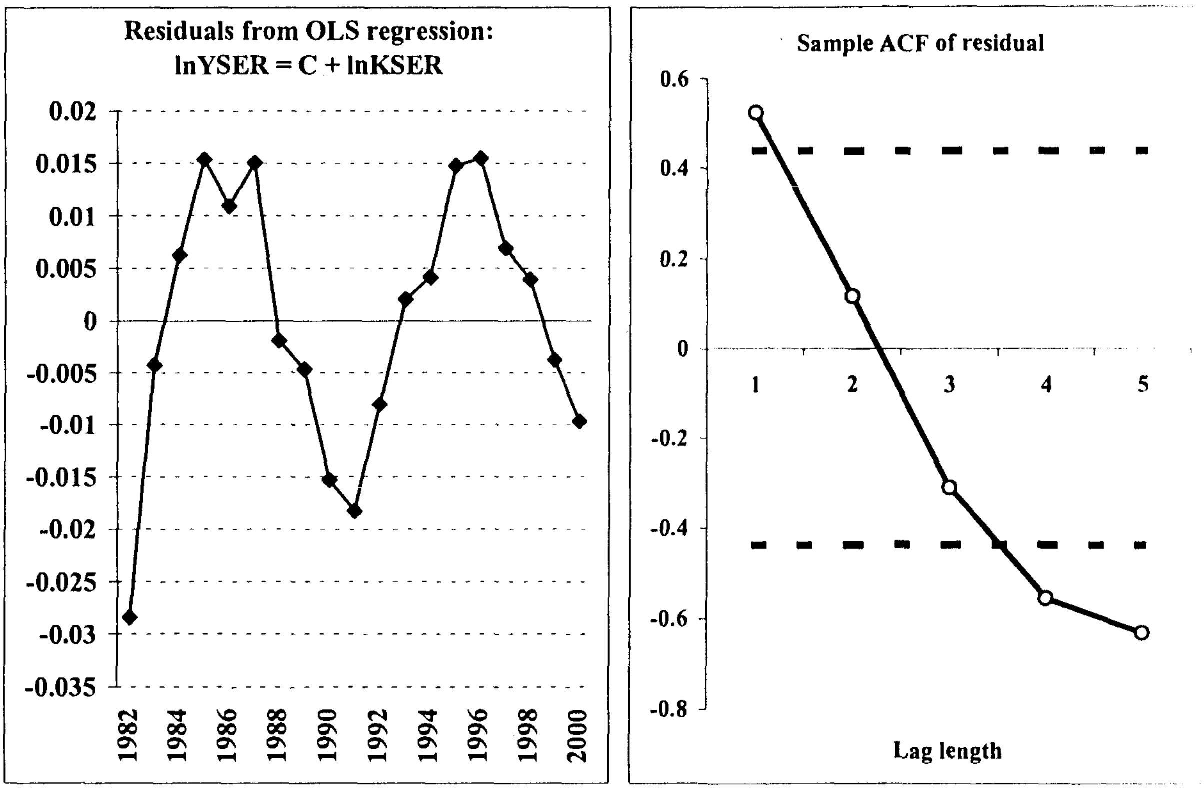
5.1.4. Estimates for value added in services model

The equation for services sector value added model comprises only two variables, lnYSER and lnKSER, both of which appeared to have very complicated data generating process resulting in indetermination of their integrating properties. Ignoring this problem a simple OLS regression results in the following equation (standard errors or t-ratios are not reported as they do not provide for valid inferences):

$$\ln YSER = 4.14 + 0.665 \ln KSER$$
$$R^2 = 0.998 \qquad D.W. = 0.565$$

Therefore, in line with our expectation the coefficient on lnKSER is positive. The fit of the model is also very impressive but the value of R^2 greater than the DW statistic is worrying and is a symptom of spurious regression. Figure 5.5 plots the residual estimated from the above relationship and its sample ACF, which clearly exhibits the problem of the above estimated long-run relationship as the residual appears to be non-stationary. The left panel in Figure 5.5 shows strong serial correlation problem, which is also reflected in the right panel as the first, fourth and fifth lag lengths in the sample autocorrelation function have values greater than their critical bands. Unfortunately, a mere use of the PHFMOLS does not help much as it also resulted in a similar pattern of unexplained variation in the data.

Figure 5.5: Residual from OLS Regression of lnYSER on lnKSER and sample ACF



Our experiments revealed that the incorporation of the lagged dependent variable removed the problem and therefore we decided to use a partial adjustment modelling framework to estimate the services sector value added model. In the partial adjustment model along with using the lagged dependent variable, higher order lags of the explanatory variable were tried and the PHFMOLS estimates of the parsimonious specification is:[33]

$\ln YSER$ =	-0.28 –	0.303 $\ln KSER_{t-1}$ +	0.353 $\ln KSER_{t-2}$ +	0.97 $\ln YSER_{t-1}$
(s.e.)	(0.43)	(0.10)	(0.11)	(0.105)
t-ratio	-0.67	2.33	3.19	9.20

Figure 5.6: Long-run relationship and sample ACF

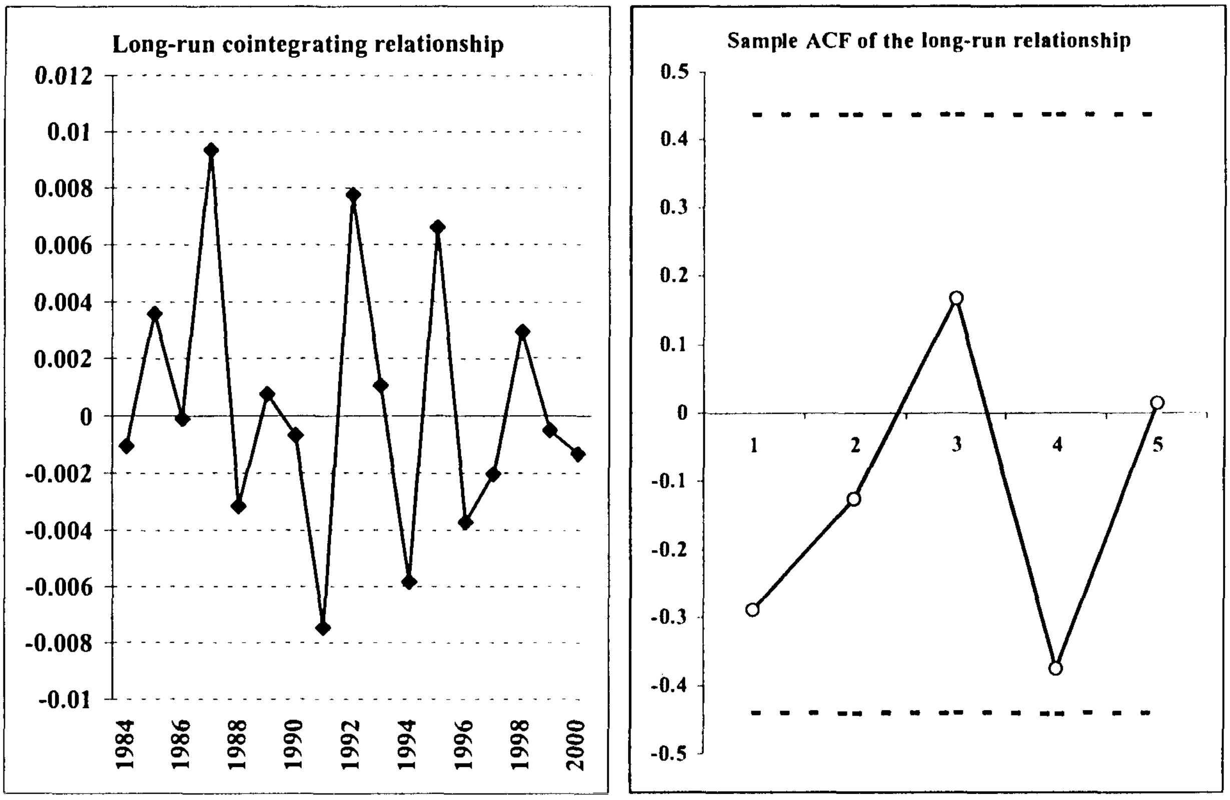


Figure 5.6 gives the resultant residual series from the partial adjustment model along with its sample autocorrelation coefficients at different lag orders. The residual now behaves more like a stationary variable. The coefficients on capital stock variables in the partial adjustment model correspond to short-run parameters and the net short run impact is 0.052. The long-run capital stock elasticity is estimated as: 0.052/(1-0.97) = 1.73.

[33] Note that using higher order lags in the short-run model is consistent with the "general to specific" methodology of model building.

5.2. Investment Block

5.2.1. Integrating Properties of Variables

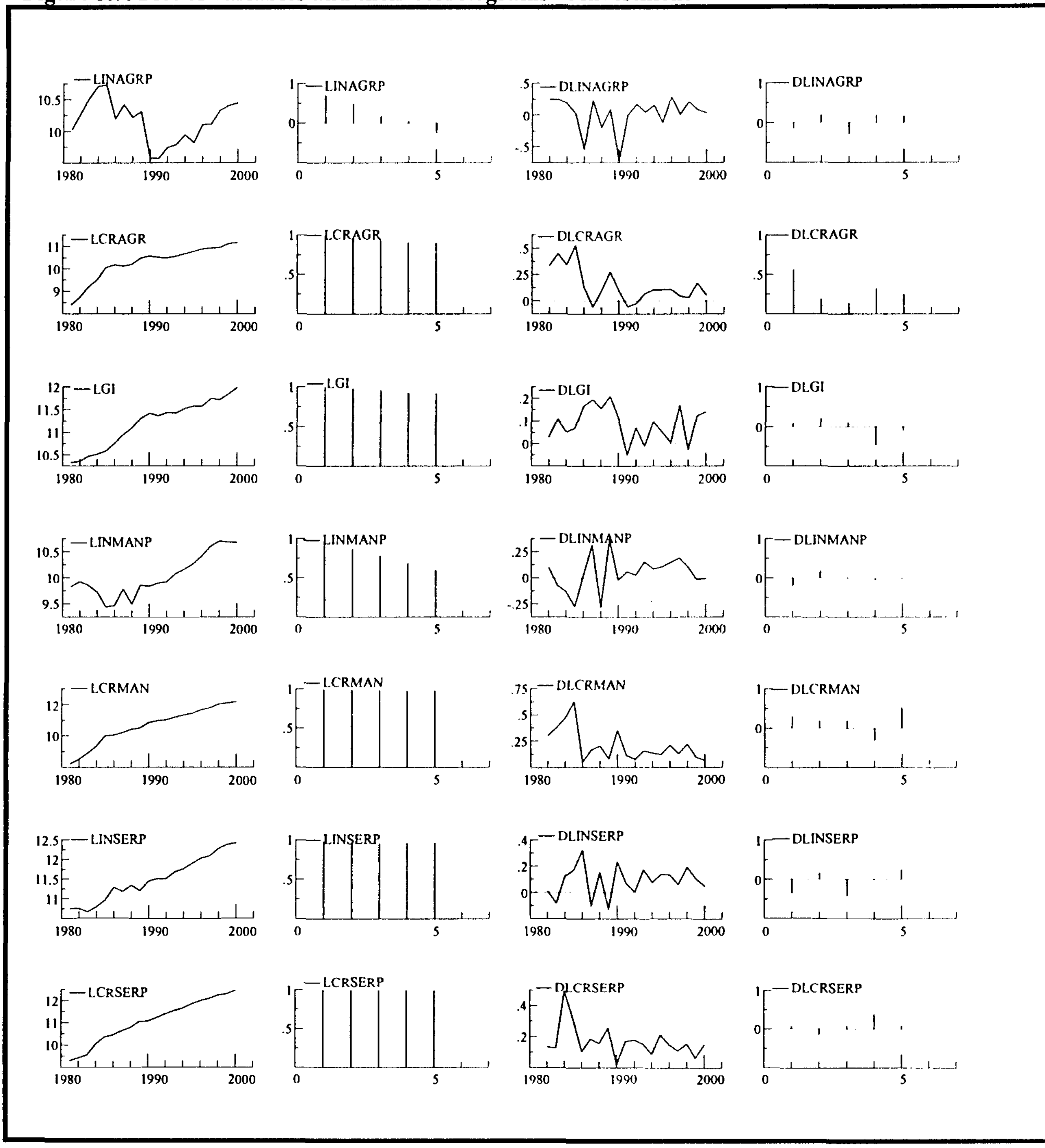
Table 5.5: Unit root test for variables in the investment block

Variables	Without trend		With trend	
	DF	ADF	DF	ADF
lnINAGRP	-1.54	-1.38	-1.36	-1.19
ΔlnINAGRP	-4.78	-2.62	-4.96	-2.77
lnCRAGR	-4.47	-3.02	-4.18	-3.85
ΔlnCRAGR	-2.78	-2.69	-2.79	-2.93
LnGI	-1.17	-1.11	-1.43	-1.56
ΔlnGI	-3.49	-2.01	-3.48	-1.99
lnINMANP	-0.13	0.26	-2.93	-2.51
ΔlnINMANP	-4.83	-2.55	-5.35	-3.08
lnCRMAN	-3.50	-3.03	-3.11	-3.01
ΔlnCRMAN	-2.99	-2.02	-3.71	-2.80
lnINSERP	-0.08	3.39	-3.52	-2.78
ΔlnINSERP	-6.72	-3.74	-6.42	-3.43
lnCRSERP	-2.40	-2.38	-2.27	-2.13
ΔlnCRSERP	-3.66	-3.99	-5.06	-5.13

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

There are three equations in the investment block – one each for agriculture, manufacturing and services sector. Private Investment in agriculture is being explained by private sector credit to agriculture (CRAGR), and public investment (GI). The DF and ADF tests clearly show that lnINAGRP is non-stationary on its level. On the other hand, since the first difference of lnINAGRP is not trended (see Figure 5.7), based on the evidence of the ADF test without the trend term, ΔlnINAGRP should be regarded as *~I(0)*. Therefore, lnINAGRP is ~ *I(1)*. Computed unit root test statistics for lnCRAGR seem to suggest stationarity of the variable on its level, which, however, is inconsistent with the expectation that first difference of a stationary variable should also be *~I(0)* as the tests fail to reject unit root for ΔlnCRAGR. Correlograms of ΔlnCRAGR do not reflect a pattern of stationary variable either. In the case of lnGI no ADF test can reject the non-stationary hypothesis both for lnGI and ΔlnGI. Moreover, the unit root test on the second difference of lnGI also failed to reject the non-stationary hypothesis. Correlograms of lnGI and ΔlnGI exhibit patterns of non-stationary and stationary variables respectively and based on this evidence lnGI will be considered as an *~I(1)* variable. Examination of autocorrelation coefficients using Box-Pierce and Ljung-Box statistics also supported this assessment.[34]

Figure 5.7: Plot of variables and their correlograms in investment block



[34] This involves first checking whether individual autocorrelation coefficients are significant. Then the Box-Pierce and Ljung-Box tests are used to determine whether all autocorrelation coefficients are simultaneously equal to zero. Since only 20 years' annual observations are in use, a maximum of 5 lag lengths were used.

Two variables in the equation explaining investment in manufacturing sector, lnINMANP and lnCRMANP are also problematic but based on the correlograms (of their levels and first differences) they can be treated as ~*I(1)* series.[35] Examinations of autocorrelation coefficients also robustly suggested first order integrating properties for these variables. Finally, the two last variables in the services sector investment equation, lnINSERP and lnCRSERP are found to be non-stationary as the ADF test with the trend term cannot reject the unit root hypothesis on their levels but on their first differences. The results of the ADF tests are also supported by the correlograms of the variables and the examinations of autocorrelation coefficients.

5.2.2. Estimates for Private Investment in Agriculture

Estimation of the private sector investment in agriculture, as specified in equation 3.4, resulted in insignificance of both the explanatory variables in the model. Besides, the residuals exhibited a pattern, which clearly indicated model misspecification. The graphical plot of lnINAGRP in Figure 5.7 shows a structural break suggesting that the behaviour of the private investment in the 1980s and 1990s might be different. Since the early 1980s structural adjustment programmes resulted in the implementation of a number of reforms concerning the agriculture sector and thus the experiences of the 1980s and 1990s may be compared as reform and post-reform periods. To capture this, the model has been re-estimated after including both the intercept shift and slope dummies for the 1990s, which is reported in the top row of Table 5.5. Since the slope dummy CRAREG turned out to be insignificant, the regression equation in the bottom row drops it. At this, all other variables become statistically significant at the conventional level. The significance of REGIME and REG2 imply that in the 1990s there has been a shift in the intercept along with a change in the slope coefficient of lnGI. This leads to two separate equations for lnINAGRP:

For 1980s: lnINAGRP = 11.76 + 0.36 lnCRAGR – 0.47 lnGI
For 1990s: lnINAGRP = -9.02 + 0.36 lnCRAGR + 1.31 lnGI

Thus, it is observed that the elasticity of private sector credit to private investment in agriculture is 0.36 for both periods of 1980s and 1990s. The impact of government investment is, however, strikingly different. In the 1980s lnGI is found to be negatively associated with private

[35] lnGI is the other variable in the manufacturing sector investment equation, integrating property of which has been discussed above.

investment while the in the later decade a highly positive relationship is discernible. In the 1980s, agriculture was heavily subsidized along with public investment of high magnitude. It might be that excessive investment by the government resulted in crowding-out effect. In the post-reform period, removal of government subsidies and other reforms might have made public investment complementary to the private sector.

Table 5.5: Long-run estimates of private investment in agriculture

Long-run relationship: PHFMOLS estimate

$$\ln INAGRP = 11.86^{***} + 0.37^{**}\ln CRAGR - 0.49^{**}\ln GI - 17.41^{***}REGIME - 0.88\ CRAREG + 2.36^{***}\ REG2$$

(s.e)	(1.41)	(0.18)	(0.21)	(4.32)	(0.86)	(0.56)
t-ratio	8.39	2.02	-2.34	4.03	-1.02	4.13

Adjusted $R^2 = 0.87$

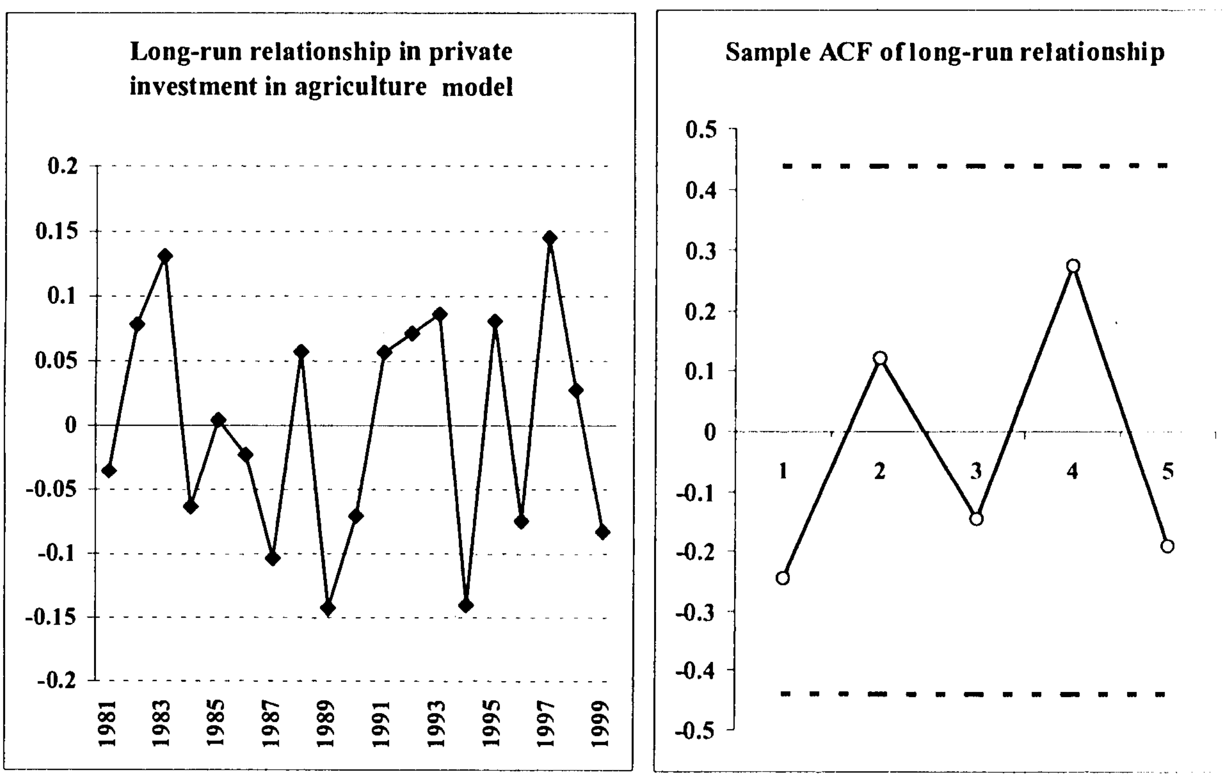
$$\ln INAGRP = 11.76^{***} + 0.36^{***}\ \ln CRAGR - 0.47^{***}\ \ln GI - 20.78^{***}\ REGIME + 1.77^{***}\ REG2$$

(s.e)	(1.44)	(0.18)	(0.21)	(2.69)	(0.23)
t-ratio	8.14	1.95	-2.22	-7.71	7.44

Adjusted $R^2 = 0.85$

Note: REGIME is a dummy variable with 0 for 1981-89 and 1 for 1990-2000. REG2 is the slope dummy for lnGI (defined as REGIME × lnGI) while CRAREG is the slope dummy for lnCRAGR (REGIME × lnCRAGR). Statistical significance at one, five and ten per cent levels are denoted by respectively ***, **, and *.

Figure 5.8: Long-run relationship in private investment in agriculture model and sample ACF



Earlier the unit root tests found that lnINAGRP and lnGI could be considered as *~I(1)* variables while the integrating order of lnCRAGR could not be determined. Therefore, it is important to check the stationarity [i.e., *~I(0)*] of residuals. The ADF test statistic in this regard was computed at 5.24 against the 95 per cent critical value of 3.74 thereby rejecting the null hypothesis of non-stationarity; in other words, the null hypothesis of non-cointegration is rejected. Figure 5.8 provides the long-run cointegrating relationship along with its sample ACF, which exhibits patterns similar to those of stationary variables.

Table 5.6: Short-run error-correction model for private investment in agriculture

ΔlnINAGRP=	-1.50^{***}	$-0.77^{*}\Delta lnGI_{t-1}$	$+0.42^{*}\Delta lnCRAGR_{t-1}$	$+0.86^{***}D86$	$+0.74^{***}D90$	$-0.78^{**}RESINAGR_{t-1}$
(s.e)	(0.23)	(0.42)	(0.218)	(0.155)	(0.137)	(0.355)
t-ratio	-6.45	-1.81	1.92	5.52	5.44	-2.19
Diagnostic Test						
Adjusted R^2=0.81						
Serial Correlation: $\chi^2(1)$= 4.24^{**}				Functional Form: $\chi^2(1)$=0.69		
Normality: $\chi^2(2)$= 0.08				Heteroscedastisity: $\chi^2(1)$=1.46		

Table 5.6 gives the short-run error-correction model for private investment in agriculture. In the short-run model REGIME and other slope dummies were not significant and hence had been deleted from the parsimonious model in Table 5.6.[36] The short-run effect of public investment is found to be negative while the elasticity of private sector credit is positive. The error-correction term is correctly signed and significant at the five per cent level. The coefficient on it suggests that all disequilibrium errors are corrected within less than one and a half year. The diagnostics point to a serial correlation problem but if the first order lag of the dependent variable was included in the model, the problem could have been overcome without causing any significant changes to any other estimated coefficients. The short-run model also turned out to be stable.[37]

5.2.3. Estimates for Investment in Manufacturing Sector

Table 5.7 provides long-run estimates of private investment in manufacturing sector. Like agriculture initial experiments with the manufacturing sector also revealed structural break in the model. Therefore, the model is estimated considering the possibility of differential intercept and

[36] The short-run model, however, includes a dummy variable for 1990, which was required to tackle the problem of non-normality of residuals.

[37] The stability of the short-run model was tested with the help of cumulative sum of recursive residuals and the cumulative sum of squares of recursive residuals.

slope coefficients for the 1980s and 1990s (see the estimated equation in Table 5.7). All variables in the equation are statistically significant at the one per cent level. It is obvious that for the 1980s the credit given to manufacturing sector is negatively related to private investment, while the effect of public investment is positive. Interestingly, however, the relationships are reversed in the later decade; it can be estimated that in the 1990s a one per cent rise in government resource displaces private investment by 0.31 per cent while there is a one to one correspondence between credit to private sector and private investment (i.e., the elasticity is estimated to be 1.04). Therefore, in the post reform period it is the supply of credit that plays the most important role in determining the private investment in the manufacturing activity.

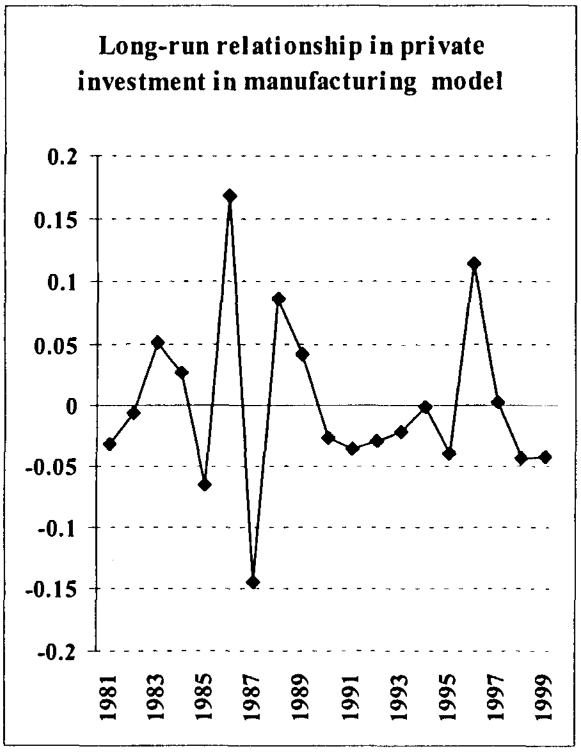
Table 5.7: Long-run estimates of private investment in manufacturing

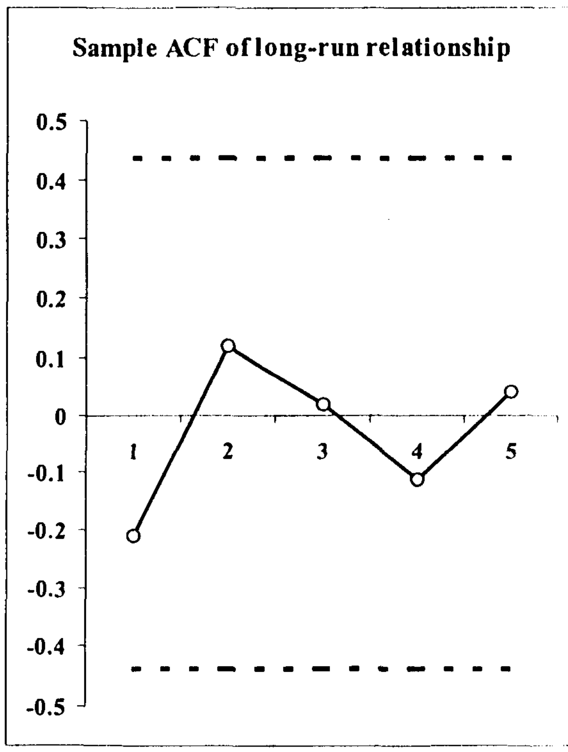
Long-run relationship: PHFMOLS estimate

lnINMANP =	8.300^{***}	-0.68^{***}lnCRMAN	$+0.79^{***}$lnGI	-6.37^{***} REGIME	$+1.83^{***}$ REGMAN	-1.21^{***} REG2
(s.e)	(0.78)	(0.08)	(0.12)	(1.94)	(0.19)	(0.33)
t-ratio	10.53	-7.97	6.04	-3.28	9.56	-3.65

Note: REGIME is a dummy variable with 0 for 1981-89 and 1 for 1990-2000. REG2 is the slope dummy for lnGI (defined as REGIME × lnGI) while REGMAN is the slope dummy for lnCRMAN (REGIME × lnCRMAN). Statistical significance at one, five and ten percent levels are denoted respectively by ***, **, and *.

Figure 5.9: Long-run relationship in the private investment in manufacturing and sample ACF





All the three variables in the model, viz., lnINMANP, lnCRMAN, and lnGI appeared to be ~I(1) and consequently a valid long-run relationship of among these variables should be represented by a stationary series. The residual obtained from the estimated equation in the bottom row of Table 5.7 was tested for stationarity and the resultant ADF test statistic was computed at -4.11 as against its 95 per cent critical value of -3.74. Since the computed ADF test statistic exceeds the critical value the null hypothesis of non-stationarity of the residual and hence non-cointegration is rejected. The graphical plot of the long-run cointegrating relationship and its sample ACF, as given in Figure 5.9, also behave like stationary series.

Table 5.8: Short-run Model for Investment in Manufacturing Sector

ΔlnINMANP =	0.138^{***}	$-\ 0.46^{**}$ ΔlnCRMAN	$-\ 1.30^{**}$ $RESMAN_{t-1}$
(s.e)	(0.045)	(0.18)	(0.40)
t-ratio	3.01	-2.48	3.24
Diagnostic Test			
Adjusted R^2=0.56			
Serial Correlation: $\chi^2(1)$= 0.038		Functional Form: $\chi^2(1)$=0.022	
Normality: $\chi^2(2)$= 0.89		Heteroscedastisity: $\chi^2(1)$=0.005	

Note: Statistical significance at the one and five percent levels are denoted by respectively *** and **.

The short-run model for private investment in manufacturing is given in Table 5.8. Initial experiments failed to establish statistical significance of the differential intercept and slope dummies and thus they were deleted from the most parsimonious representation in Table 5.8. In the short-run, therefore, only the credit to manufacturing sector significantly influences private investment. The error-correction term, $RESMAN_{t-1}$, though correctly signed bears a value higher than 1 (absolutely) making the system explosive. However, imposition of the restriction that $RESMAN_{t-1}$ = -1 results in a Wald test statistic of 0.58. Since the Wald test statistics follow a chi-square distribution with degrees of freedom being equal to the number of restriction (in this case 1) we cannot reject the restriction – in which case all disequilibroum errors will be corrected within just one year. Diagnostic tests associated with the short-run model do not detect any problem.

5.2.4. Estimates for Private Investment in Services Sector

The model on private investment in the services sector was also estimated considering a structural break for the 1990s.[38] The top row in Table 5.9 shows that in terms of the intercept and the slope coefficient associated with lnGI, the estimated equation for 1990s is different from that of 1980s. From the bottom row of Table 5.9 it is obvious that credit to services sector elasticity is 0.79 for both periods. However, while public investment seems to have crowded-out private investment in services in the 1980s (the elasticity being −0.15), in the post-reform period of the 1990s lnGI has been complimentary to the lnINSERP. It is observed that in the 1990s a-one per cent rise in public investment is associated with about 0.47 per cent increase in private investment in services. The residual representing the long-run relationship between lnINSERP, lnCRSERP and lnGI appears to be stationary (see Figure 5.10) and the test for cointegration resulted in an ADF test statistic of 3.91 vis-à-vis the critical value of 3.74 thereby providing evidence to a valid long-run relationship.

Table 5.9: Long-run estimates of private investment in services

Long-run relationship: PHFMOLS estimate

lnINSERP =	3.96***	+ 0.94*** lnCRSERP	- 0.30 lnGI	– 7.67*** REGIME	- 0.22 REGSER	+ 0.91** REGI
(s.e)	(0.69)	(0.17)	(0.18)	(1.60)	(0.22)	(0.30)
t-ratio	5.74	5.61	-1.66	-4.78	-1.03	2.98

lnINSERP =	3.89***	+ 0.79*** lnCRSERP	– 0.15* lnGI	– 6.97*** REGIME	+ 0.62*** REG2
(s.e)	(0.65)	(0.10)	(0.09)	(1.41)	(0.12)
t-ratio	5.92	7.83	-1.72	-4.93	4.99

Note: REGIME is a dummy variable with 0 for 1981-89 and 1 for 1990-2000. REG2 is the slope dummy for lnGI (defined as REGIME × lnGI) while REGSER is the slope dummy for lnCRSERP (REGIME × lnCRSERP). Statistical significance at one, five and ten percent levels are denoted by ***, **, and * respectively.

[38] This is because initial experiments showed that estimates for the 1980s had been significantly different from those of the 1990s.

Figure 5.10: Long-run relationship in private investment in services and sample ACF

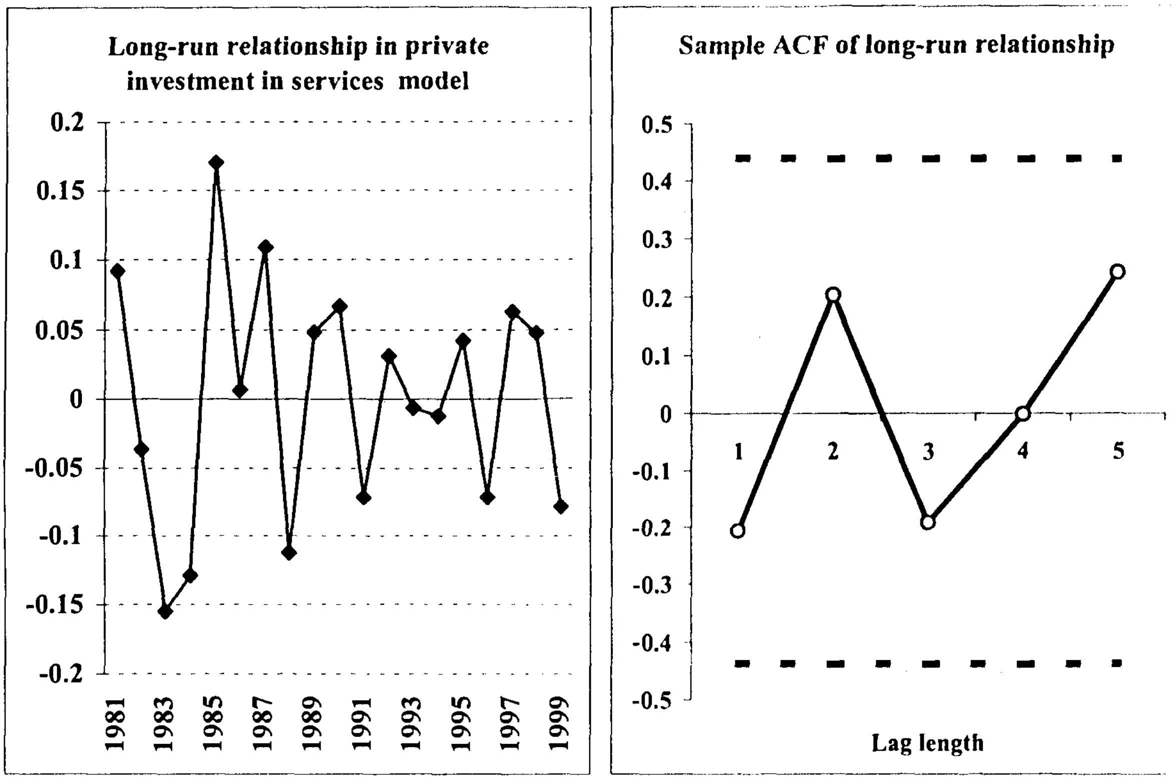


Table 5.10: Short-run model for private investment in the services sector

ΔlnINSERP =	0.136^{***}	$- 0.22^{*}$ ΔlnCRSERP	$- 1.22^{***}$ $RESSER_{t-1}$
(s.e)	(0.02)	(0.11)	(0.136)
t-ratio	6.06	-1.98	-8.988

Diagnostic Test
Adjusted R^2=0.83

Serial Correlation: $\chi^2(1)$= 3.80	Functional Form: $\chi^2(1)$=2.40
Normality: $\chi^2(2)$= 1.17	Heteroscedastisity: $\chi^2(1)$=0.41

Note: Statistical significance at one, five and ten percent levels are denoted by respectively ***, **, and *.

In the short-run error-correction model only lnCRSERP is found to be significantly influencing the private investment but its significance appears to be somewhat at a lower level of 10 per cent. The error-correction term, $RESSER_{t-1}$, is correctly signed and significant. Although the coefficient on the error-correction terms is greater than 1 (absolutely), its 95 per cent confidence interval retains the value of 1. Therefore, the hypothesis that any deviation from the long-run relationship cannot be dissipated can be rejected.

5.3. Private Consumption Function

5.3.1. Unit root test of variables

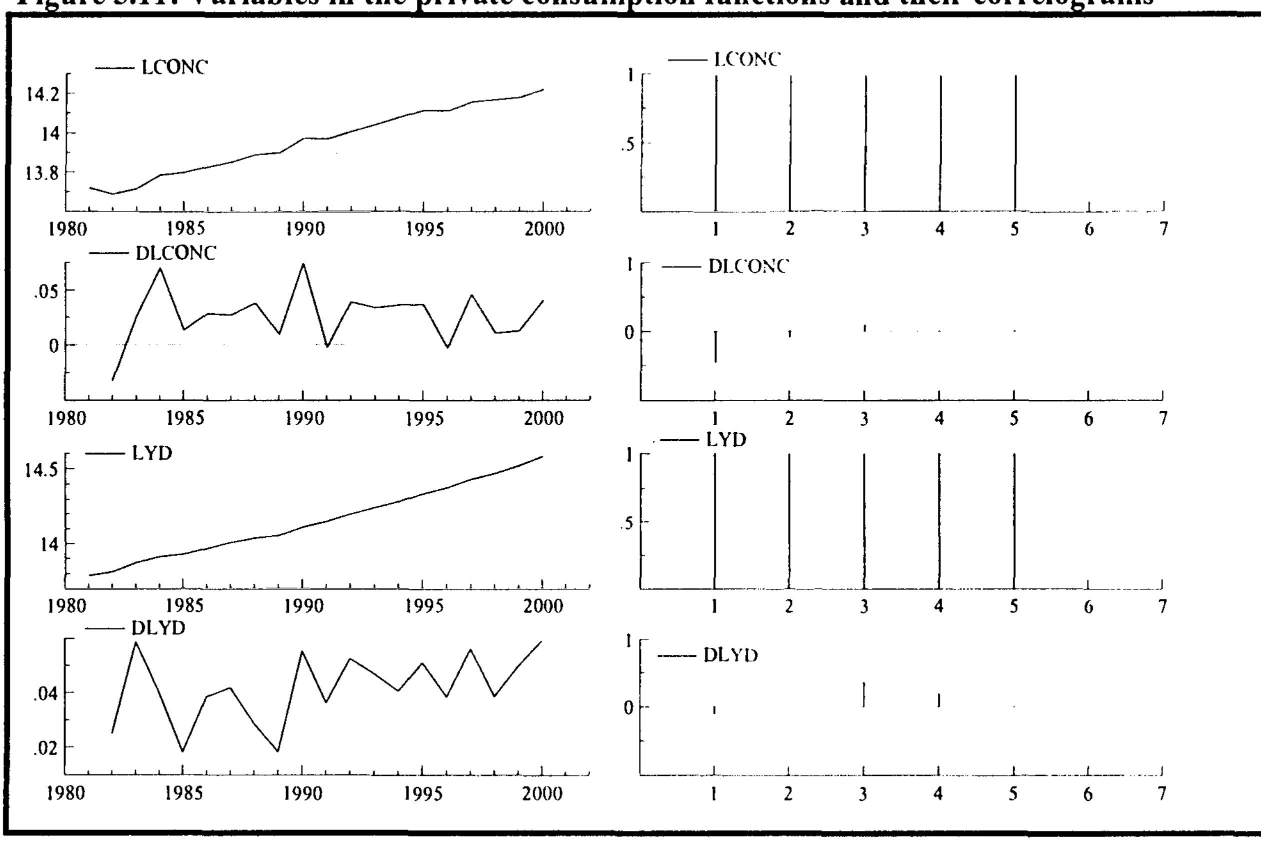
Private consumption expenditure is the most important component of aggregate demand and has been specified as a function of disposable income alone. The results of the DF-ADF regressions unambiguously show that lnCONC is *~I(1)*, as the variable on its level is non-stationary while its first difference is stationary. Every DF-ADF test also indicates the non-stationarity of disposable income (lnYD) on its level. Since the graphical plot of ΔlnYD is slightly trended (see Figure 5.11), based on the evidence of the ADF test with the trend term it should be treated as an *~I(0)* variable. Therefore, lnYD is *~I(1)*.

Table 5.11: Results of DF-ADF Tests

	Without Trend		With Trend	
Variables	DF	ADF	DF	ADF
LnCONC	-0.98	-0.84	-3.51	-2.51
ΔlnCONC	-7.53	-6.57	-8.34	-7.69
LnYD	1.57	1.96	-0.63	-0.19
ΔlnYD	-3.92	-2.46	-5.40	-4.39

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

Figure 5.11: Variables in the private consumption functions and their correlograms



5.3.2. Estimating the long-run relationship

A simple OLS estimate of the consumption function, as reported in the top row of Table 5.12, appeared to be problematic as the R^2 value was higher than the DW statistic. The residuals were found to be strongly serially correlated and thus were not stationary.[39] The PHFMOLS method did not cure the problem requiring us to reformulate the empirical specification of the consumption function.

Table 5.12: Simple linear and Partial Adjustment Specification of the Consumption function

Simple linear specification: OLS estimate

$\ln CONC = -4.02 + 0.70 \ln YD$

Adjusted $R^2 = 0.97$ D-W statistic = 0.59

Partial Adjustment Mechanism: PHFMOLS estimate

$\ln CONC =$	1.89^{***}	$+ 0.26^{**} \ln YD$	$+ 0.60^{***} \ln CONC_{t-1}$	$- 0.036^{**}$ D90
(s.e.)	(0.53)	(0.08)	(0.12)	(0.0139)
t-statistic	3.52	2.97	4.83	-2.59
Adjusted $R^2 = 0.99$				

Note: *** and ** imply statistical significance at the one and five per cent levels respectively. D90 is a dummy variable with 0 for 1990 and 1 for otherwise.

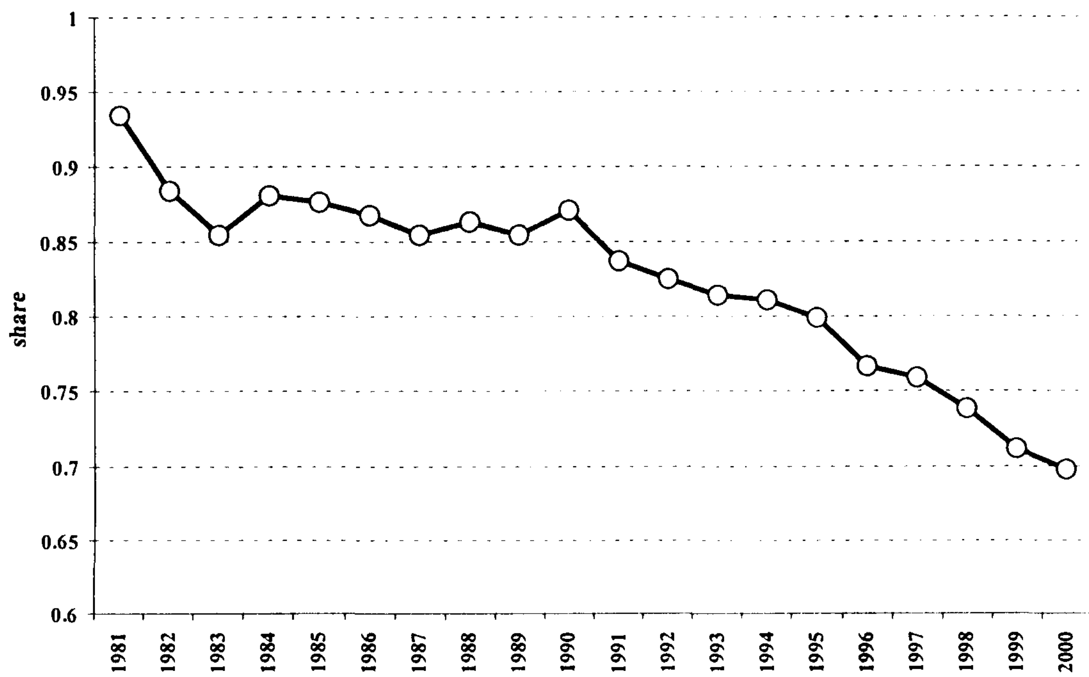
One alternative estimation strategy is to employ the partial adjustment framework, which resulted in the estimates as reported in the bottom row of Table 5.12. Due to non-stationarity of the data the model is estimated using the PHFMOLS procedure. In the regression the coefficient on the lagged dependent variable is correctly signed and highly significant thus demonstrating a valid partial adjustment mechanism. The short-run income elasticity is 0.26, which is extra-ordinarily low, while its long-run counterpart is estimated to be 0.65. It is obvious that the PAM is estimated after inserting a dummy variable for 1990 as the year proved to be atypical in initial runs. The fit of the model is very good and the examination of the autocorrelation coefficients of the resultant residual provided the evidence for its stationarity.

While the partial adjustment modelling framework worked quite well, the information gathered from the data pointed toward a structural break in the estimated relationship. This is being reflected in Figure 5.12 where it is found that compared to the 1980s there has been a drastic

[39] The reported DW statistic itself shows the serial correlation problem.

decline in the share of private consumption expenditure in total disposable income. After an initial fall in 1982, the average propensity to consume in the 1980s stabilized just above 0.85, but in the following decade it had gone down dramatically to less than 0.7. Because of this large fall in relative share in private consumption, the ratio of national saving to GDP has increased from 5.69 per cent in 1989 to 17.54 per cent in 2000.[40] As Figure 5.12 shows, the period of fast decline in private consumption has coincided with the policy reforms of the 1990s. For a low-income country like Bangladesh it is not very easy to explain such a drastic fall in the average propensity of consumption. It might be that national income has risen to a level to support a considerable amount of national saving.[41] Trade liberalization measures might also have provided incentive to divert a portion of consumption expenditure to saving and investment, although there is no solid micro evidence to substantiate the hypothesis. Whatever may be reasons, without inferring anything about the causal relationship one can state that the post-reform period is associated with a considerable fall in private consumption-ratio.

Figure 5.12: Share of private consumption expenditure in total disposable income



Source: Authors' own computation from the National Accounts data (BBS 2000 and 2001)

[40] National saving is the GDP less private consumption expenditure less government consumption.

[41] But still almost half of the population remains poor failing to fulfill bare necessities. Therefore, one big question is whether the rise in saving took place as a result of rise in inequality.

Table 5.13: Private consumption function with structural breaks and the short-run error-correction model

Long-run private consumption function: PHFMOLS estimates

lnCONC =	1.92^{***}	$+ 0.85^{***}$ lnYD	$+ 4.19^{***}$ REGIME	$- 0.295^{***}$ REGYD	$+ 0.026^{**}$ D83	$- 0.0271^{***}$ D95
(s.e)	(0.506)	(0.036)	(0.555)	(0.039)	(0.008)	(0.0079)
t-ratio	3.79	23.35	7.54	-7.46	3.00	-3.42

Short-run error-correction model: OLS estimates

ΔlnCONC =	0.06^{**}	$+ 0.42^{+}$ΔlnYD	$- 0.04^{***}$ D84	$- 0.04^{**}$ D90	$+ 0.03^{**}$ D96	$- 0.68^{*}$ $RPHCON_{t-1}$
(s.e)	(0.02)	(0.25)	(0.012)	(0.012)	(0.012)	(0.38)
t-ratio	2.56	1.68	-3.49	-3.27	2.51	-1.78

Diagnostic Test

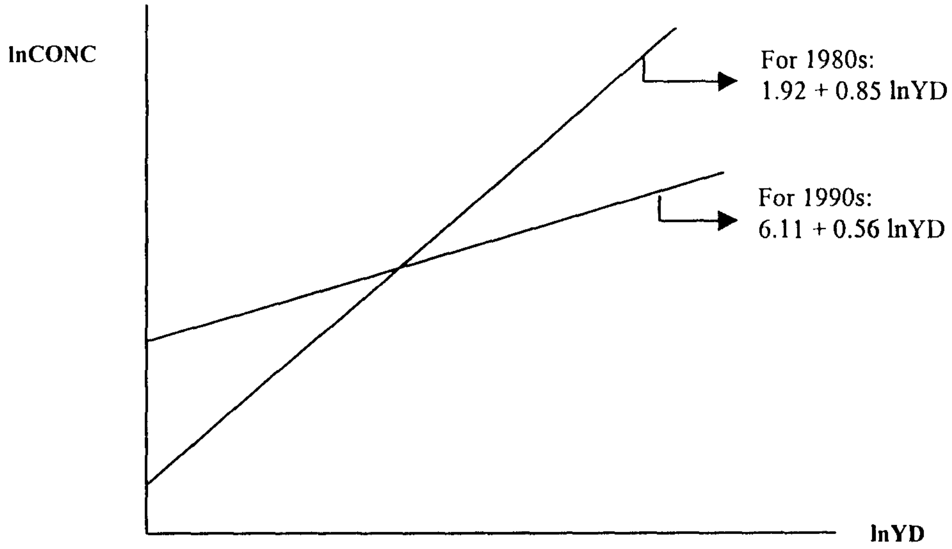
Adjusted R^2=0.70

Serial Correlation: $\chi^2(1)$= 2.05	Functional Form: $\chi^2(1)$=1.03
Normality: $\chi^2(2)$= 0.38	Heteroscedasticity: $\chi^2(1)$=1.45

Note: REGIME is a dummy variable with 0 for 1981-89 and 1 for 1990-2000. REGYD is the slope dummy for lnYD (defined as REGIME × lnYD). ***, ** and * imply statistical significance at the one and five and ten per cent levels respectively. $^{+}$ denotes significance at the 12 per cent level.

The top row in Table 5.13 therefore gives the long-run estimates of the private consumption expenditure function for Bangladesh considering a structural break for the 1990s. The model is estimated with the PHFMOLS procedure giving valid standard errors for statistical inferences. In the estimated regression along with the intercept and lnYD, both the differential intercept (REGIME) and the slope dummy REGYD are found to be significant at the one per cent level. This means that the consumption function for the 1990s will have a different intercept and different income elasticity from that of the 1980s. The estimated coefficient on lnYD, 0.85, in the regression denotes the consumption elasticity for the 1980s while the corresponding elasticity for the 1990s is estimated to be 0.56 (i.e., 0.85–0.295). Similarly, the intercepts for the 1980s and 1990s are respectively 1.92 and 6.11. Figure 5.13 thus exhibits the graphical plots of the consumption functions for two different periods. It needs to be mentioned here that due to non-normality of errors, the long-run relationship has been estimated with two dummy variables for 1983 and 1995, which are also significant at the 1 per cent level.

Figure 5.13: Private consumption function for the 1980s and 1990s

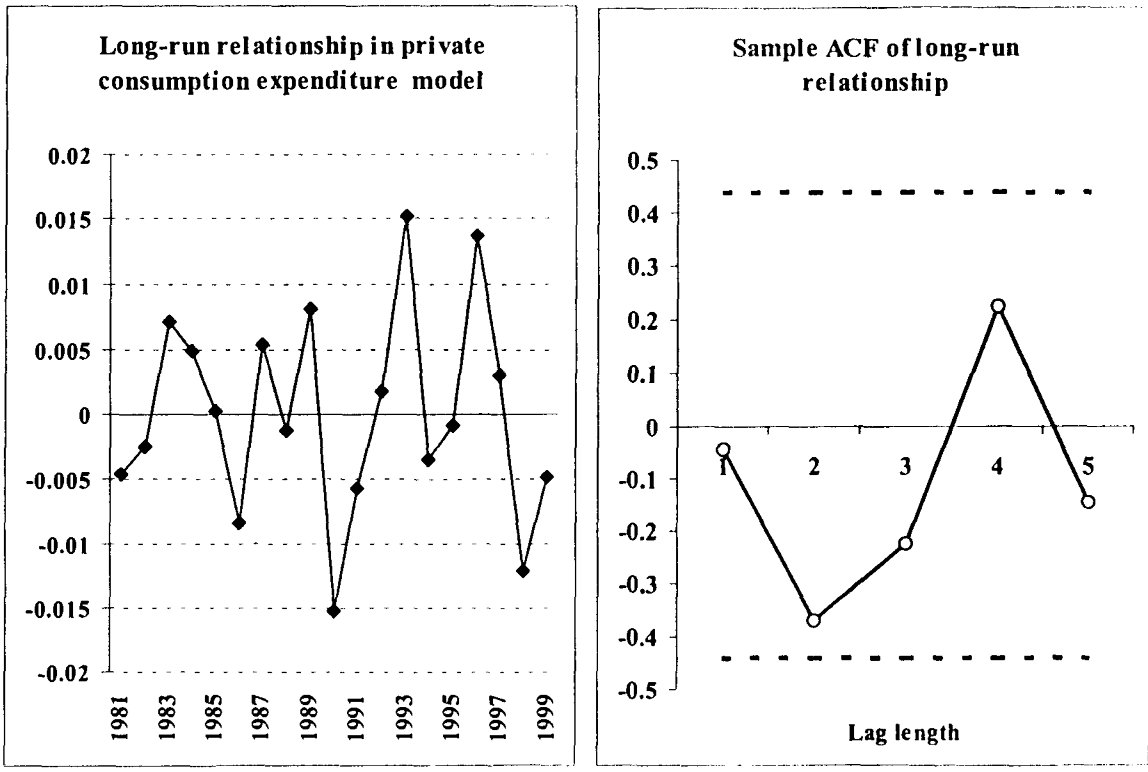


With regard to the coefficient on lnYD, it should be noted that two previous studies on Bangladesh also reported small magnitudes (along with other large estimates) of MPC. In various specifications Rahman and Shilpi (1990) estimated the long-run income elasticity of consumption to be in the range 0.49-0.99. However, rather than using a specific equation, they experimented with a number of models and then selected one of the elasticity estimates that best suited to be 'reasonable'.[42] On the other hand, using a number of theoretical specifications Kabir's (1997) estimates of long-run MPC fell in the range 0.42-0.89. For both the studies data were available for the 1970s and 1980s only and, therefore, the possibility of having a different consumption function for the post-reform period could not be explored. Given the fact that the 1990s had witnessed a dramatic rise in domestic saving, our estimate of low income elasticity is not unreasonable. In Figure 5.12, the APC in 2000 is 0.68 and since throughout the 1990s APC is falling, MPC must be even lower. Thus even without any econometric estimate, it can be deduced that in the late 1990s MPC has been quite low.[43]

[42] In their consumption functions, Rahman and Shilpi (1990) experimented with such variables as broad money and remittance. The use of broad money in the consumption function is unusual while it is not clear why the effect of remittance will not be captured in disposable income.

[43] Unless, there are reasons to believe that the data are subject to measurement error.

Figure 5.14: Long-run cointegrating relationship and its sample ACF



The estimated long-run private consumption expenditure, as presented in the first row of Table 5.13, provided robust evidence for cointegration between lnCONC and lnYD as the ADF test statistic for testing the non-stationarity of residual was estimated at –4.23 against its 99 per cent critical value of 3.90. The graphical plots of residual and its autocorrelation coefficients, as given in Figure 5.14, also look like those of a stationary variable.

The corresponding short-run error-correction model is presented in the bottom row of Table 5.13. In the short-run model the intercept and slope dummies were not significant while the estimated income elasticity, 0.42, achieved statistical significance only at the 15 per cent level. Three dummies – one each for 1984, 1990 and 1996 – all are significant at the conventional statistical significance levels were required to control for unusual movements in residuals. The coefficient on the error-correction term ($RPHCON_{t-1}$) is estimated to be –0.68, which is significant only at the twelve per cent level. Diagnostic tests do not report any problem with the estimated model.

5.4. Foreign Trade Block

5.4.1. Exports

5.4.1.1. Unit Root Tests of the Variables

In the exports block there are three equations explaining exports of jute and jute manufactures, ready-made garments, and exports of all other items. First, in the case of jute and jute goods the relevant variables are lnXJMR, lnPJ and lnJMC of which, both on the basis of DF-ADF tests and correlograms (as given in Figure 5.15), we have unambiguous evidence that lnPJ is ~ *I(0)*. Based on the graphical plot and correlograms, lnXJMR is to be considered as non-stationary while its first difference appears to be stationary. Thus despite inconclusive evidence of the DF-ADF tests, lnXJMR can be treated as an *~I(1)* variable. For lnJMC no trend is observed on its level and consequently taking into account of the results of DF-ADF regressions without the trend term the null hypothesis of non-stationarity can be rejected. That the first difference of a stationary variable is also stationary is maintained by the unit root tests on ΔlnJMC. Figure 5.15 shows that both the correlograms of lnJMC and ΔlnJMC behave in a similar way and are in conformity with a stationary variable.

Table 5.14: Unit root test of the variables in the Exports Block

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
lnXJMR	-1.30	-1.13	-2.29	-2.23
ΔlnXJMR	-4.55	-2.93	-4.41	-2.86
lnPJ	-2.05	-1.99	-2.28	-2.44
ΔlnPJ	-4.22	-4.15	-4.13	-4.14
lnJMC	-3.22	-3.52	-3.13	-3.33
ΔlnJMC	-4.59	-4.07	-4.54	-3.99
lnXRMGR	-4.79	-3.91	-2.36	-2.43
ΔlnXRMGR	-2.43	-1.47	-3.61	-2.56
lnPRMG	-0.63	-1.98	-1.37	-2.43
ΔlnPRMG	-1.46	-2.28	-2.64	-3.77
lnRMGC	-2.65	-1.53	-4.68	-3.10
ΔlnRMGC	-7.36	-4.81	-7.15	-4.67
lnXOGR	-1.57	-1.52	-3.20	-3.86
Δ lnXOGR	-4.11	-4.21	-3.88	-4.05
lnPOG	-0.97	-0.87	-2.59	-2.95
Δ lnPOG	-4.73	-4.31	-4.85	-4.59
lnXOGC	-2.86	-3.58	-2.80	-3.45
ΔlnXOGC	-4.09	-4.35	-3.89	-4.12

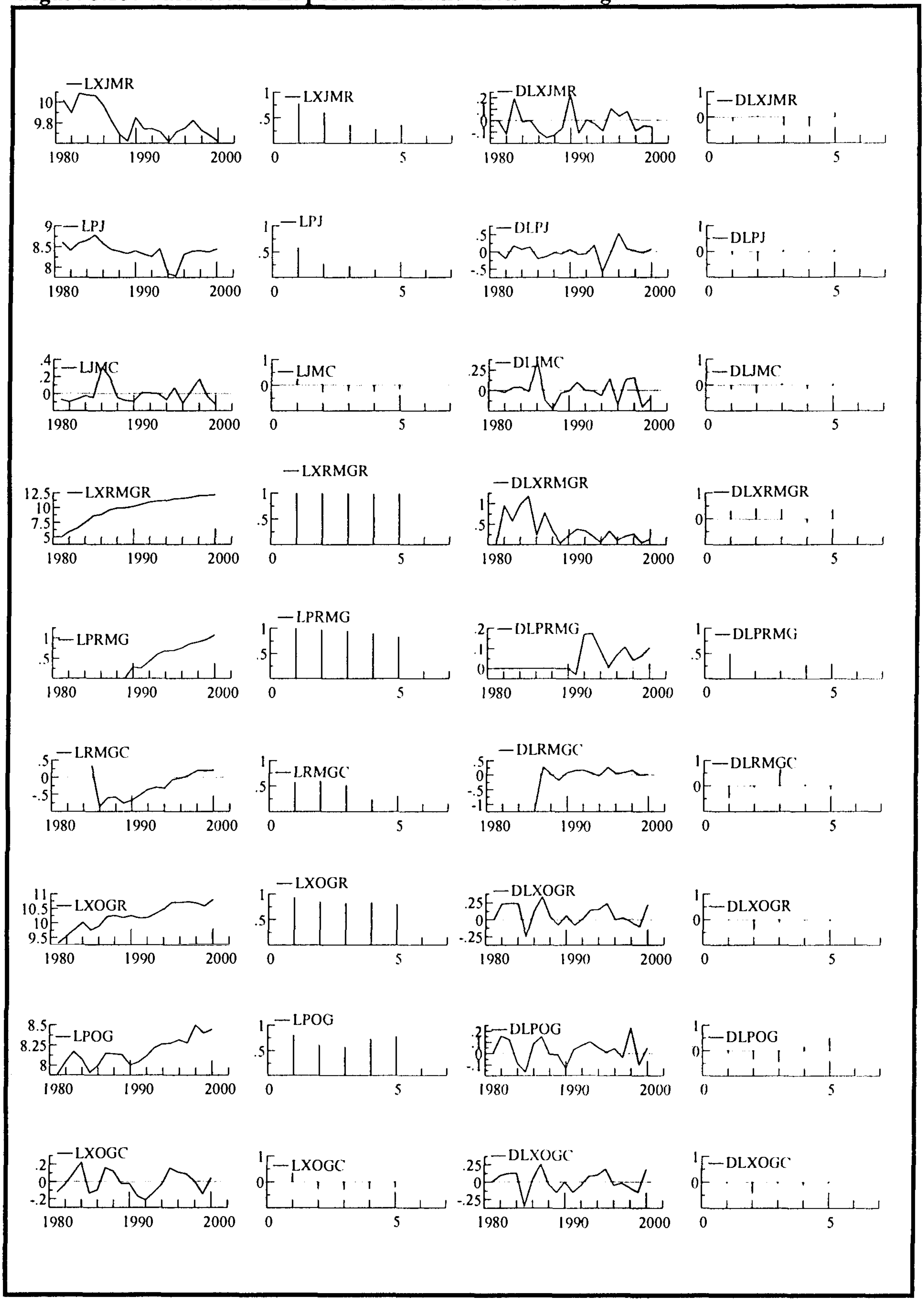
Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

In the ready-made garments (RMG) equation, three relevant level variables are lnXRMGR, lnPRMG and lnRMGC. lnXRMGR is strongly trended and the relevant unit root test cannot reject the null hypothesis of non-stationarity. The DF-ADF tests cannot also reject the non-stationarity for the first difference of lnXRMGR. In addition, the correlograms of ΔlnXRMGR are unlikely to be of a stationary variables. Unit root tests were carried out on the second difference of lnXRMGR but no unambiguous results could be obtained. All DF-ADF tests on lnPRMG demonstrate non-stationarity of the variable on its level but the test-results are not straightforward on its first difference, ΔlnPRMG. There are only 10 years' data on lnPRMG with which any meaningful test for unit root is impossible to carry out. Therefore, like most other variables we will assume that lnPRMG ~I(1). The capacity utilization variable in the production of RMG, lnRMGC, is strongly trended and based on the ADF test with the trend term it is a non-stationary variable.[44] The correlograms of lnRMGC also portrays a pattern not characteristic of a stationary variable. The first difference of lnRMGC, ΔlnRMGC, is clearly found to be an *~I(0)* variable, on the basis of DF-ADF tests and correlograms. As a consequence, lnRMGC can be treated as an *~I(1)* variable.

Turning to all other exports, the ADF test with the trend suggests stationarity of lnXOGR although the graphical plot and correlograms clearly suggest otherwise. The tests and correlograms, however, overwhelmingly indicate stationarity of the first difference variable, ΔlnXOGR. Therefore, it would be most appropriate to consider lnXOGR as an *~I(1)* series. All the DF-ADF tests cannot reject the unit root hypothesis for lnPOG, while the same tests on its first difference can. Hence lnPOG is *~I(1)*. The graphs and correlograms of lnPOG and ΔlnPOG also support this conclusion. Finally, lnXOGC may be regarded as an *~I(0)* variable. The graphical plot shows that the variable is not trended and the ADF test without the trend term rejects the null hypothesis of non-stationarity. The stationarity of the first difference of lnXOGC, ΔlnXOGC, is borne out by every test and correlograms.

[44] The DF test with the trend term, however, rejects the unit root for RMGC.

Figure 5.15: Variables in Exports Block and their correlograms



5.4.1.2. Estimating the Supply Function for Jute and Jute Goods

Bangladesh has been the largest exporter of raw jute and during the past three decades its share in world raw jute market has increased from 56 to about 95 percent (Razzaque, 2002). However, as jute has witnessed a stiff competition from synthetic products that act as a substitute for jute goods, between the early-1970s and late-1990s the world annual imports of raw jute plummeted by more than 52 per cent from about 900000 to about 430000 metric tons. Similarly, during the same time although Bangladesh's share in jute goods has shown a slightly increasing trend from around 40 per cent to about 50 percent, compared to the early-1980s world exports of jute goods in the late 1990s have fallen by about 50 per cent.[45] Therefore, despite an increasing share in the world market export quantities exported from Bangladesh have been mostly stagnant in the 1990s. The rapid decline in world demand has also resulted in drastic fall in prices of jute and jute goods. In the face of the depressed demand and falling prices Bangladesh has been able to increase its share in world market by virtue of being the most efficient jute producer and thereby driving away other relatively inefficient competitors.

Initial experiments with the long-run export supply function of jute and jute goods revealed the problem of structural break in the estimated equation. This problem is somewhat reflected in the graphical plot of lnXJMR, which declines massively between 1983 and 1989 but in the following decade such a trend is arrested by fluctuating export earnings of jute and jute items. The top row of Table 5.15, therefore, attempts to capture the differential effects between 1980s and 1990s by introducing the intercept and slope dummies. In initial experiments lnJMC and the slope dummy associated with turned out to be not significant and hence in the top row of Table 5.15 they have been dropped. It is now observed that the intercept and price slope dummies are highly significant providing support for the structural breaks. The price elasticity of export supply for the 1980s is 1.09 while the corresponding figure for the 1990s is estimated to be –0.15. The estimated relationship between lnXJMR and lnPJ was found to be a cointegrating one as the ADF test for residual stationarity yielded a test statistic of -3.91 against the 99 per cent critical value of -3.90. The graph of the residuals from the long-run relationship and its correlograms, as given in Figure 5.16, also support a cointegrating relationship.

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[45] These figures are as computed in Razzaque (2002).

Table 5.15: Export Supply Function for Jute and Jute Goods

Long-run supply function: PHFMOLS

lnXJMR =	0.576	+ 1.096*** lnPJ	+ 8.79*** REGIME	-1.25*** REGPJ	
(s.e)	(1.49)	(0.17)	(1.69)	(0.25)	
t-ratio	0.386	6.25	5.19	-5.26	
Adjusted R^2=0.72					

Short-run supply function: OLS estimates

ΔlnXJMR =	-0.0058	+ 1.18*** ΔlnPJ	+ 0.21* ΔlnJMC	- 1.02*** REGDPG	- 0.84** $RPHJ_{t-1}$
(s.e)	(0.017)	(0.26)	(0.120)	(0.27)	(0.298)
t-ratio	-0.33	4.45	1.72	-3.71	-2.83

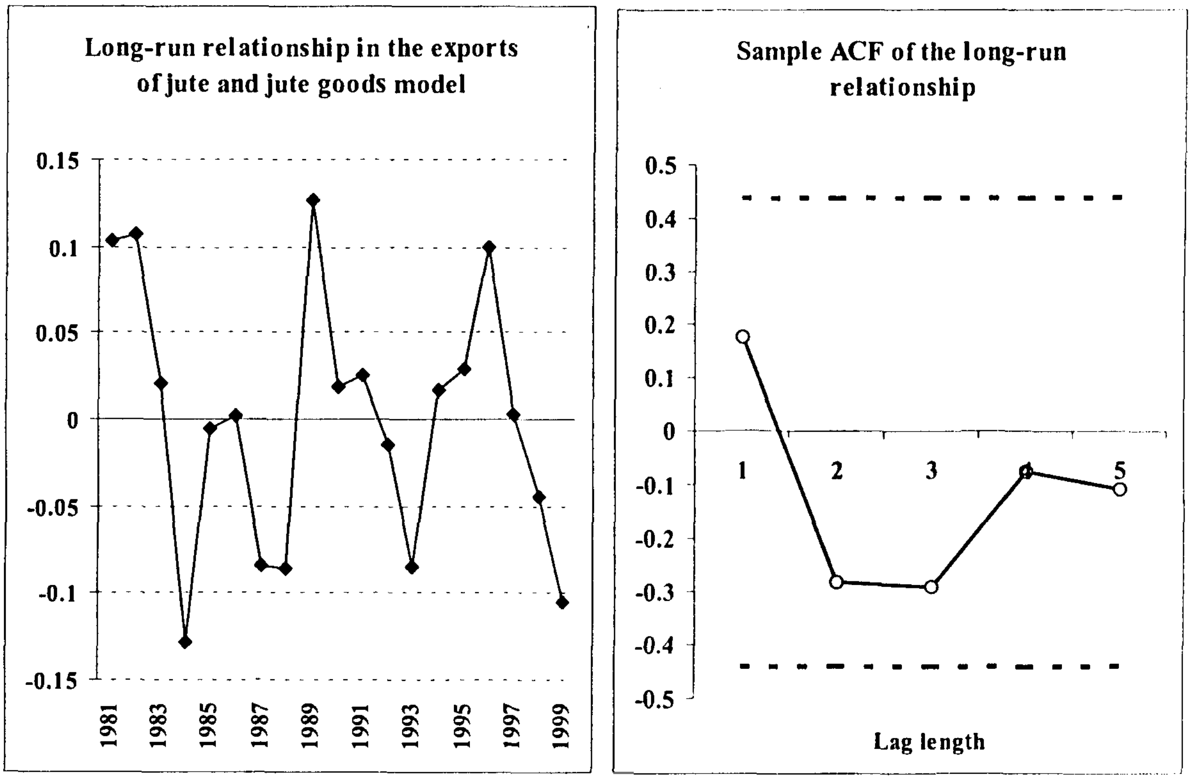
Diagnostic Test

Adjusted R^2=0.53

Serial Correlation: $\chi^2(1)$= 1.11	Functional Form: $\chi^2(1)$=0.20
Normality: $\chi^2(2)$= 0.69	Heteroscedastisity: $\chi^2(1)$=0.26

Note: REGIME is a dummy variable with 0 for 198-89 and 1 for 1990-2000. REGPJ is the slope dummy for lnPJ (defined as REGIME × lnPJ). REGDPJ is the slope dummy for ΔlnPJ (defined as REGIME × ΔlnPJ). Statistical significance at one, five and ten per cent levels are denoted by respectively ***, **, and *.

Figure 5.16: Cointegrating relationship in the exports of jute and jute goods and its sample ACF



In the short-run model, as presented in the bottom row of Table 5.15, the slope dummy associated with the price elasticity remains significant at the one per cent level. The short-run elasticity for the 1980s is 1.18 while for 1990s the figure is estimated to be 0.16. In the short-run model the capacity utilization variable also appears to be significant at the ten per cent level. The error-correction term, $RPHJ_{t-1}$, shows a rapid adjustment to the long-run path as 84 per cent of short-run disequilibrium errors are corrected within just one year. Despite a low value of R^2, the short-run model does not have any problem of serial correlation, functional form, non-normality of errors and heteroscedasticity.

5.4.1.3. Estimating the Supply Function for Ready-made Garments

Table 5.16 provides the estimated supply functions of RMG. It must be mentioned here that due to unavailability of the on price of RMG, the sample size of this regression is only 11 covering 1990-2000. The estimated equation in the top row of Table 5.16 shows that all variables in the long-run equation are statistically significant at the one per cent level. Other things remaining the same, a one per cent increase in the price of RMG, increases its supply by 0.85 per cent. The capacity utilization elasticity is also positive and is obtained as 1.27.

Table 5.16: Estimated Export Supply Function of RMG

		Long-run relationship: PHFMOLS estimates		
lnXRMGR =	10.97^{***}	$+ 0.85^{***}$ lnPRMG	$+ 1.27^{***}$ lnRMGC	
(s.e)	(0.06)	(0.07)	(0.0.07)	
t-ratio	13.64	-9.70	31.41	
		Short-run error-correction model: OLS estimates		
ΔlnXRMGR =	0.82^{***}	$+1.06^{**}$ ΔlnPRMG	$+ 0.48^{***}$ ΔlnRMGC	$- 0.75\ ecm_{t-1}$
(s.e)	(0.22)	(0.48)	(0.10)	(0.20)
t-ratio	3.90	2.25	4.8	-3.0

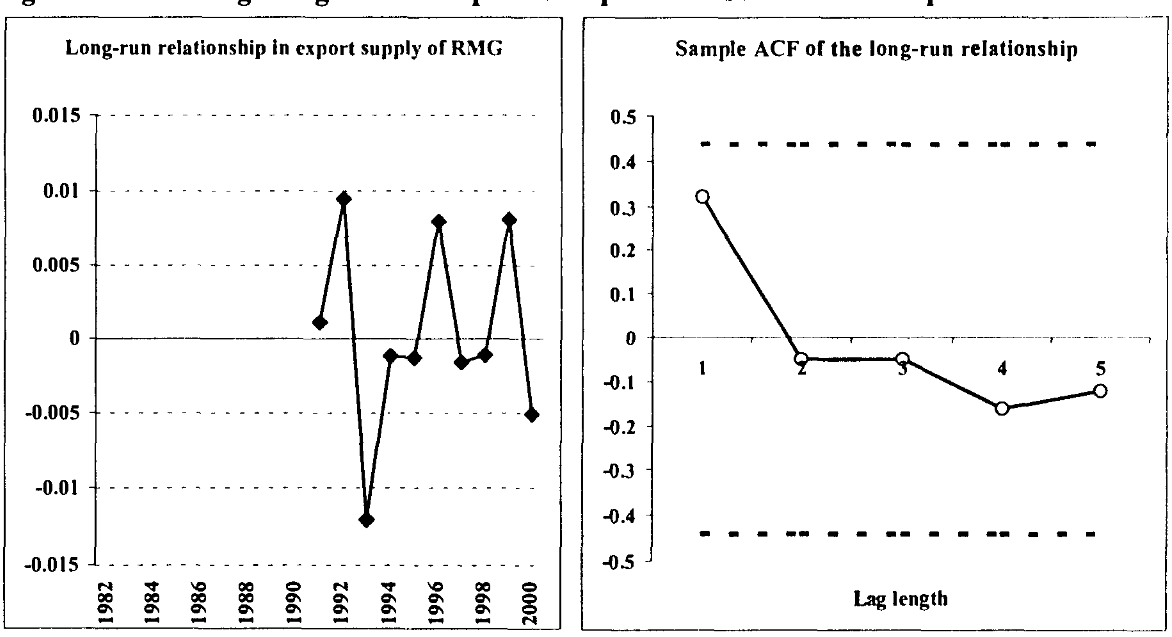
Diagnostic Test

Adjusted R^2=0.75

Serial Correlation: $\chi^2(1)$= 2.15	Functional Form: $\chi^2(1)$=1.22
Normality: $\chi^2(2)$= 0.25	Heteroscedastisity: $\chi^2(1)$=0.34

Note: Statistical significance at the one and five per cent levels are indicated by ***, and ** respectively.

Figure 5.17: Cointegrating relationship in the exports of RMG and its sample ACF



The DF-ADF tests to determine the stationarity of the residuals from the long-run equations is unlikely to be meaningful because small sample size. The plot of the residual and autocorrelation functions, as given in Figure 5.17, however, depict a pattern similar to that of a stationary variable and we will assume that it represent a valid cointegrating relationship.

Assuming that the estimated long-run equation is a cointegrating relationship, the short-run error-correction model is reported in the bottom row of Table 5.16. In the short-run the price elasticity of supply is computed to be 1.04, which is significant at the 5 per cent level. The sign of the coefficient on capacity utilization, like its long-run counterpart, is also positive; the capacity utilization elasticity being 0.48. The coefficient on the error-correction term is also significant, which, in fact, further supports a valid cointegrating relationship of the model presented in the top row of the Table. The speed of adjustment is quite rapid as 75 percent of the disequilibrium errors is adjusted within the first year. The diagnostic tests do not suggest any problem of serial correlations, non-normality of errors, functional form problem and heteroscedasticity.

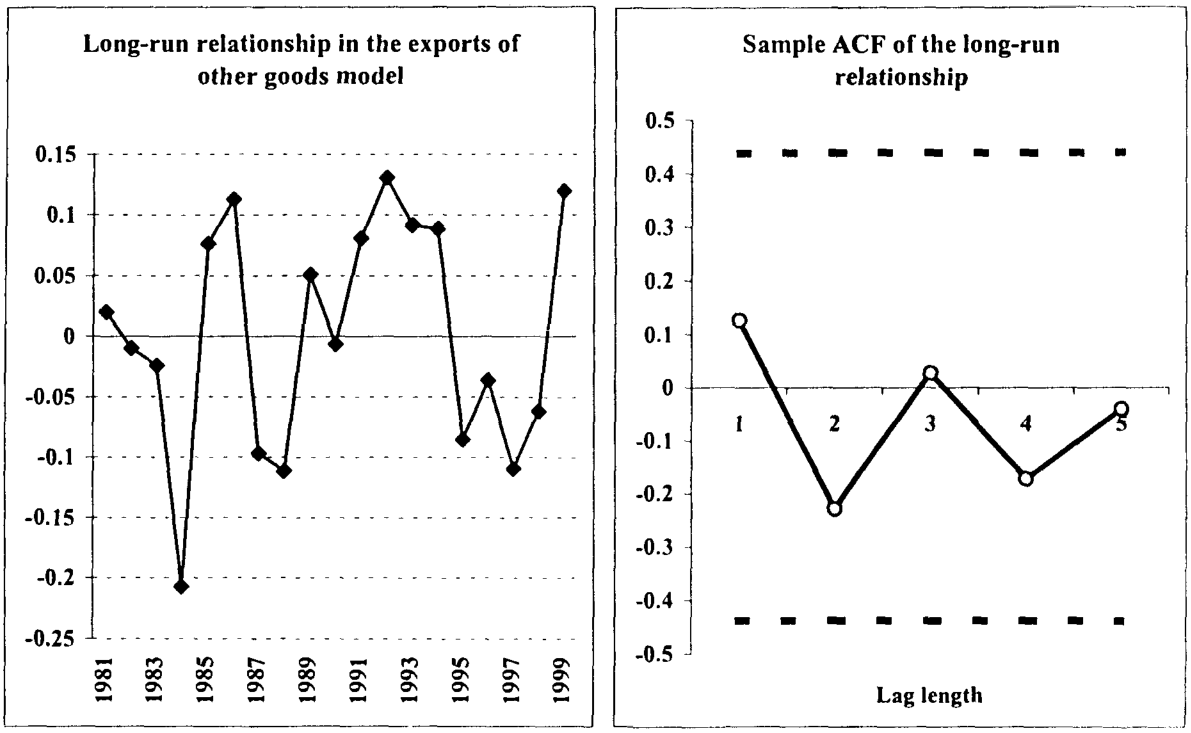
5.4.1.4. Estimating the Supply Function for Exports of Other Goods

Modelling supply of other exports via error-correction techniques was not satisfactory as the residuals from the first step long-run relationship exhibited strong autocorrelation. Initial experiments showed that the problem could be alleviated by including the lagged dependent variable in the regression equation. Therefore, we decided to use the partial adjustment framework in estimating the short-run and long-run parameters of the supply function of exports of other goods. The PHFMOLS estimate of the model is as follows:

$\ln XOGR =$	$-\ 0.29\ +$	$0.279^{**} \ln POG$	$+\ 0.61^{***} \ln XOGC$	$+\ 0.81^{***} \ln XOGR_{t-1}$
(s.e)	(0.78)	(0.139)	(0.12)	(0.06)
t-ratio	-0.37	2.02	5.09	13.18

It is found that the price, capacity and the lagged dependent variables turn out to be statistically significant. The coefficient on $\ln XOGR_{t-1}$ is positive and hence represents a correct specification for partial adjustment model. The short-run price elasticity of supply is 0.28 but the long-run elasticity is estimated to be 1.47. Similarly, the short- and long-run capacity elasticity are respectively 0.61 and 3.21.

Figure 5.18: Cointegrating relationship in the exports of other goods its sample ACF



Due to the presence of a mixture of *I(1)* and *I(0)* variables (lnXOGR and lnPOG are ~I(1), while lnXOGC is an *~I(0)* series), the Pesaran *et al.* test was carried out to test for cointegration. The F-statistic for this purpose was computed at 40.94 against its critical value of 3.74. Therefore, the null hypothesis of non-cointegration is rejected very strongly. For cointegration the combination of *I(1)* and *I(0)* variable must result in an *~I(0)* variable and the graphical plot of the residuals from the estimated partial adjustment model, as given in Figure 5.18, indeed looks like a stationary variable, which is further supported by the sample autocorrelation coefficients.

5.4.2. Imports

5.4.2.1. Unit Root Tests for the variables

On the imports side we specified two equations – one for the imports of intermediate goods and raw materials and the other for all other imports. The equation for imports of intermediate goods (lnMRMR) is specified as a function of price (lnPMR) and GDP (lnY) and similarly the import of other goods' (lnOMR) is explained by price (lnPO) and GDP (lnY). Table 5.17 gives the results of the DF-ADF unit root tests for these variables, while Figure 5.19 plot the variables on their levels, first differences, and their correlograms.

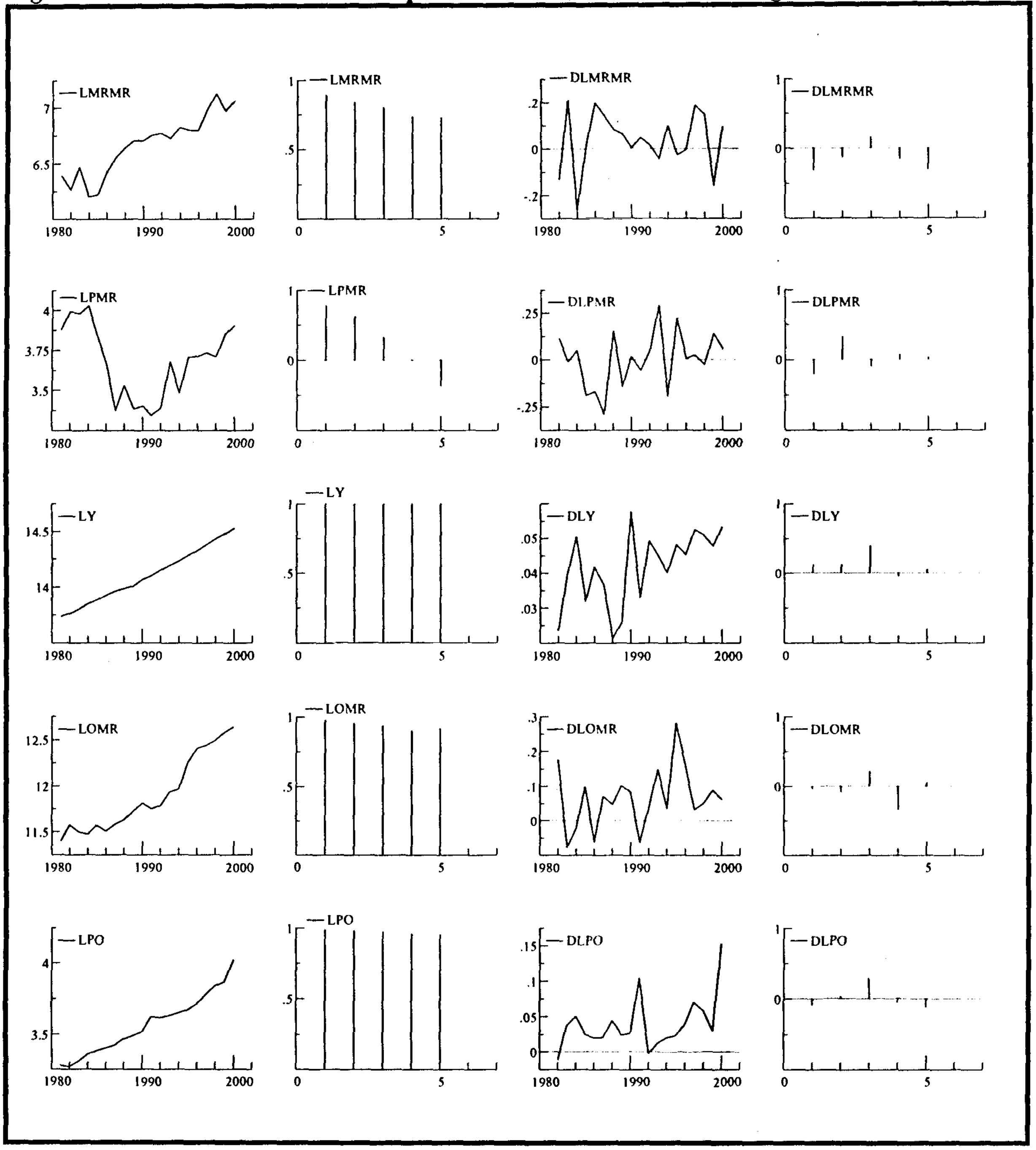
Table 5.17: DF-ADF tests on variables in the imports bloc

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
LnMRMR	-1.10	-0.77	-3.04	-2.58
ΔlnMRMR	-5.12	-3.63	-4.92	-3.58
LnPMR	-1.65	-1.40	-1.33	-0.80
ΔlnPMR	-4.78	-2.08	-5.98	-2.99
LnY	2.18	2.24	-0.30	-0.04
ΔlnY	-3.43	-2.28	-4.41	-3.46
LnOMR	1.00	1.12	-2.41	-2.19
ΔlnOMR	-3.85	-2.86	-4.06	-3.28
LnPO	1.52	1.73	-0.46	-0.01
ΔlnPO	-2.99	-1.99	-3.35	-2.54

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

It is found that the non-stationarity of lnMRMR is being rejected by all DF-ADF tests, but the evidence on stationarity of ΔlnMRMR is ambiguous due to contrasting results of the ADF tests with and without the trend term. However, since the first difference of lnMRMR is not trended (see Figure 5.19), based on the ADF test without the trend term ΔlnMRMR can be considered as an *~I(0)* variable. The graphical plot and DF-ADF tests on lnPMR confirm its non-stationarity, but none of the ADF tests could reject the null of unit root for its first differenced varaible, ΔlnPMR. Examination of graphical plot, correlograms and sample autocorrelation coefficients, however, seem to suggest stationarity of ΔlnPMR in which case lnPMR is to be an *~I(1)* series. For GDP a similar problem arises, as the tests cannot reject unit root hypothesis for ΔlnY. The graphical plot shows that the first difference of lnY is trended, but still the ADF test with the trend term falls short of its 95 percent critical value. The correlograms, however, appear to be supportive of stationary variable and therefore ΔlnY might be treated as a trend stationary process (TSP). It is to be mentioned here that being the growth rate of GDP, non-stationarity of ΔlnY is very unlikely as well as unrealistic.

Figure 5.19: Plot of variables in import functions and their correlograms



Integrating orders of the two remaining variables in Table 5.17, LOMR and LPO, are also inconclusive. Although all the DF and ADF regressions suggest non-stationarity of the level variables, no ADF test can reject the unit root hypothesis for their first difference counterparts. In contrast, correlograms of ΔlnOMR and ΔlnPO clearly depict patterns to be associated with stationary variables. It was found that individual autocorrelation coefficients for both these first differenced variables were not statistically significant and the hypotheses that such autocorrelations were simultaneously equal to zero were upheld by the Box-Pierce and the Ljung-Box statistics.[46] Therefore, it might not be inappropriate to conclude that LOMR and LPO are *~I(1)* series.

5.4.2.2. Estimating the equation for imports of intermediate goods and raw materials

Since non-stationary variables characterize the estimating equation, the PHFMOLS procedure is employed to estimate the static long-run relationship results of which are given in the top row of Table 5.18. The equation is actually estimated after inserting a dummy variable for 1983 (D83) as the year proved to be an atypical one. All variables including D83 are found to be significant at the 1 per cent error probability level. Imported intermediate goods and raw materials are found to be price inelastic, as a one percent rise in price would reduce the demand by only about 0.4 per cent. This reflects very low backward integration into Bangladesh's economy and excessive dependence on imports.[47]

Table 5.18: Long- and short-run estimates of the equation explaining imports of intermediate goods and raw materials

Long-run equation: PHFMOLS estimate

$$\ln MRMR = -8.06^{***} - 0.37^{***}\ln PMR + 1.16^{***}\ln Y - 0.29^{***}D83$$

(s.e)	(0.87)	(0.06)	(0.06)	(0.06)
t-ratio	-9.19	-5.74	18.78	-4.28

Short-run error-correction model: OLS estimates

$$\Delta\ln MRMR = 0.02 - 0.294^{**}\Delta\ln PMR - 0.27^{***}D83 + 0.29^{***}D84 - 0.85^{***}RPHMR_{t-1}$$

(s.e)	(0.097)	(0.099)	(0.075)	(0.065)	(0.27)
t-ratio	0.24	-2.95	-3.64	4.50	-3.15

Diagnostic Tests

Adjusted R^2=0.74

Serial Correlation: $\chi^2(1)$= 0.03	Functional Form: $\chi^2(1)$=1.04
Normality: $\chi^2(2)$= 0.49	Heteroscedastisity: $\chi^2(1)$=1.32

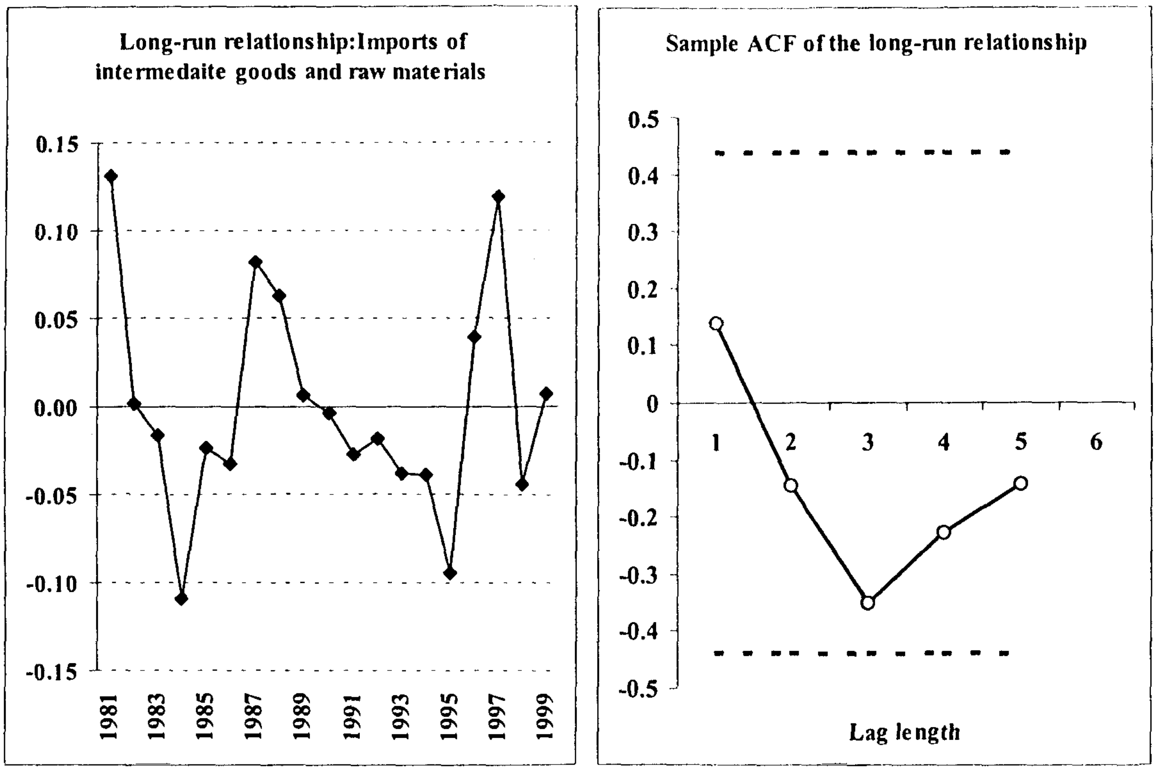
Note: Statistical significance at the one and five per cent levels are indicated by ***, and ** respectively. D83 is the dummy variable representing 0 for 1983 and 1 otherwise. Similarly, D84 represents 0 for 1984 and 1 otherwise.

The evidence of cointegration in the long-run equation is rather weak. The ADF test statistic on the residual is estimated at –3.15 against the 95 per cent critical value of -3.74. The graphical plot of the long-run relationship also does not demonstrate its stationarity convincingly although the autocorrelation coefficients upto the 5th lag length fall within the 95 per cent error bands (see Figure 5.20).

[46] These results are available from authors on request.

[47] It is well known that the biggest export sector, ready made garments, heavily relies on imported intermediate goods and raw materials as the country lacks adequate backward linkage activities to support the final export items.

Figure 5.20: Cointegrating relationship in the imports of intermediate goods and raw materials model and its sample ACF



Despite the somewhat weak evidence of cointegrating relationship, the short-run is modelled following the error-correction mechanism. In the short-run no significant income elasticity is observed but the price inelastic demand for intermediate goods and raw materials is retained. In addition to D83, another dummy for 1984, D84, was required to ensure normality of errors in the short-run model. The error-correction term, $RPHMR_{t-1}$, has a coefficient, which is less than 1 (absolutely), negative and statistically highly significant. All this represents a sound representation of the short-run model and can be used to bolster the evidence of a genuine long-run relationship in the first step static equation.

5.4.2.3. Estimating the equation for imports of other goods

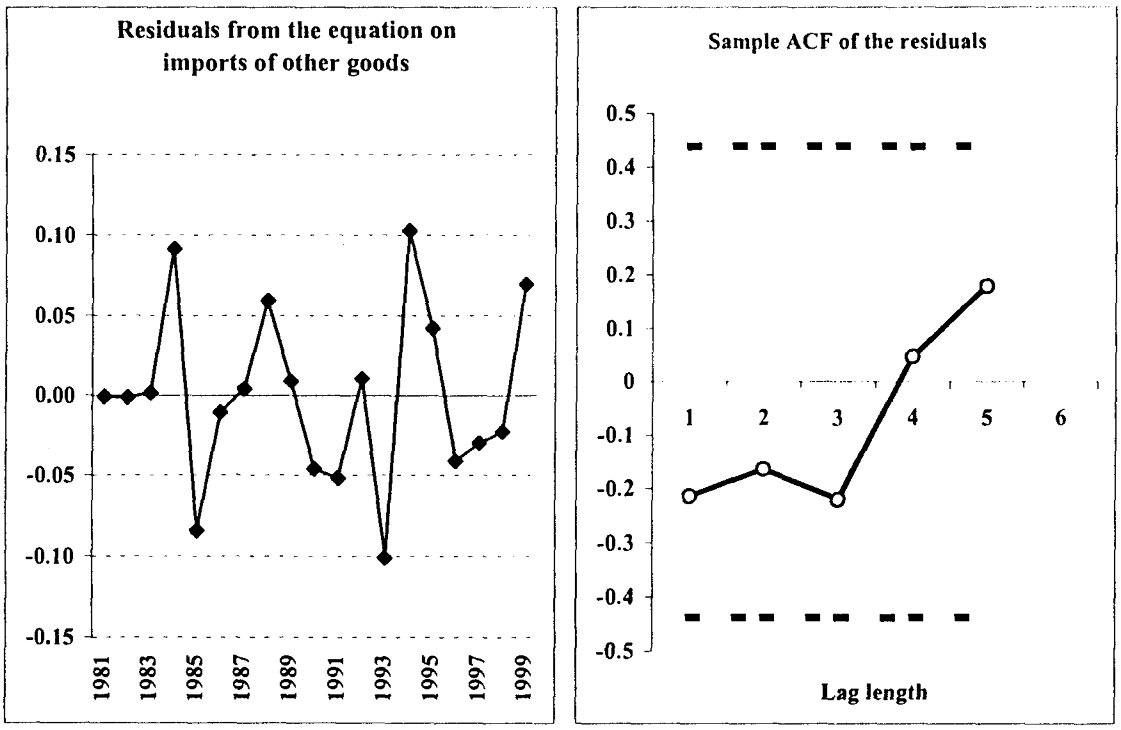
Estimation of single equation static long-run relationship in the case of imports of other goods resulted in strong serially correlated residuals. But the use of the partial adjustment technique removed the problem yielding the following results:[48]

lnOMR =	-15.82^{***}	$-\ 1.18^{***}$ lnPO	$+\ 1.69^{***}$ lnY	$+\ 0.676^{***}$ $\ln OMR_{t-1}$
(s.e)	(3.43)	(0.385)	(0.357)	(0.11)
t-ratio	-4.60	-3.06	4.74	5.89

[48] Due to non-stationarity of the data, this model, as usual, is estimated using the PHFMOLS procedure.

In the above regression result the positive sign on the lagged dependent variable along with its high level of statistical significance provides support to the correct specification of the partial adjustment model. The price and income elasticities also turn out to be significant at the one per cent level. In the short-run a one per cent increase in price reduces the demand for imports by about 1.2 per cent while the same increase in income raises the demand by 1.7 per cent. The long-run price and income elasticities of demand for other imports are estimated to be respectively –3.64 and 5.22. Thus the demand for other imported goods appear to be highly price and income elastic in the long-run.

Figure 5.21: Residuals from the partial adjustment model of imports of other goods and its sample ACF



Since the estimation of partial adjustment model involves integrated variables, the residuals from the estimated equation are checked for stationarity. The ADF test statistic in this regard is estimated to be –5.57 against the 99 per cent error probability level critical value of –4.64. Therefore, there is strong evidence that the partial adjustment model estimates a valid relationship. The plot of the residuals and the sample ACF as presented in Figure 5.21 seem to support this result very strongly.

5.5. Government Revenue Block

5.5.1. Government Revenue

5.5.1.1. Unit Root Test of Variables

Revenue from imports (lnREVM), revenue from domestic sources (lnREVIN) and non-tax revenues (lnREVNT) are the three equations included in the government revenue block. Table 5.19 provides the results of the unit root test of the variables in various revenue specifications where the DF-ADF test results are found to be inconclusive for every variable. This requires us to determine the order of integration of individual variables based on their graphical plots, correlograms, and autocorrelation coefficients. It is obvious from Figure 5.22 that all variables on their levels are non-stationary. On the other hand, while the graphs of first difference of lnREVM and lnIMP clearly look like stationary variables, based on graphical plots and correlograms ΔlnREVIN and ΔlnREVNT might also be considered as *I(0)* variables. However, the graph of ΔlnNY does not seem to be stationary making it a very complicated data generating process. In fact, even the second difference of lnNY also reflected a non-stationary pattern.

Table 5.19: DF-ADF tests on variables in the government revenue bloc

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
LnREVM	-1.08	-1.07	-1.25	-1.54
ΔlnREVM	-3.67	-2.68	-3.86	-2.62
LnIMP	0.68	0.78	-2.58	-2.40
ΔlnIMP	-4.40	-3.44	-4.45	-3.58
LnREVIN	-1.49	-1.46	-0.99	-0.85
ΔlnREVIN	-3.20	-2.11	-3.73	-2.66
LnNY	-3.91	-2.65	-2.33	-2.31
ΔlnNY	-2.11	-1.74	-3.26	-3.20
LnREVNT	-0.008	0.05	-2.45	-2.39
ΔlnREVNT	-4.56	-3.04	-4.36	-2.80
LnNY	-3.91	-2.65	-2.33	-2.31
ΔlnNY	-2.11	-1.74	-3.26	-3.20

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

5.5.1.2. Empirical Estimation of Revenue from Imports

In Table 5.20 the estimated long- and short-run models on revenue from imports are reported. In the long-run there is a one to one correspondence between imports and revenues from it, while in the short-run the magnitude of elasticity is found to be considerably lower at 0.52. The long- and short-run versions of the model are estimated using two dummies for 1992 and 2000, which are statistically significant in both regressions.

Figure 5.22: Plot of variables on their levels and first differences

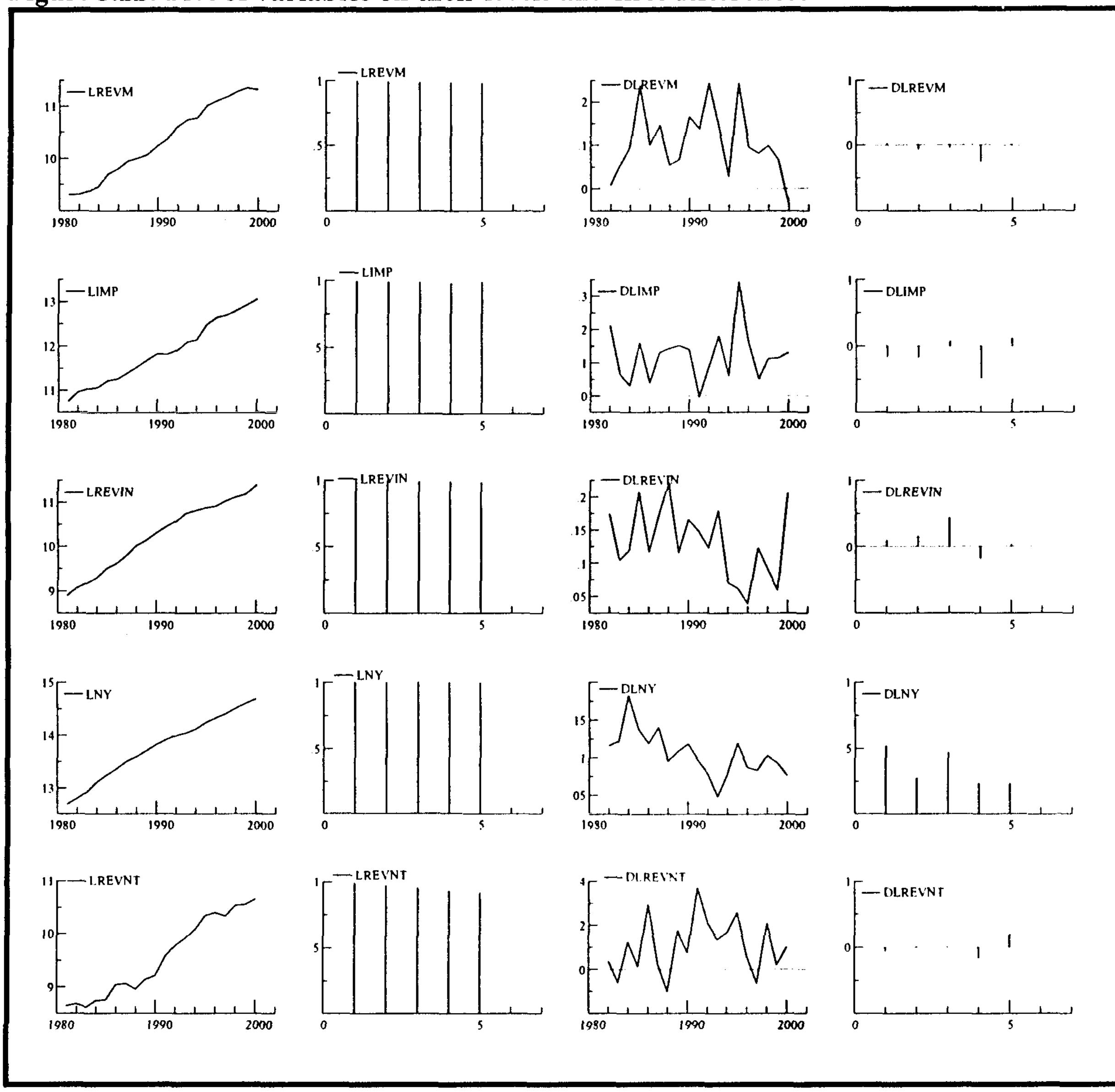


Table 5.20: Estimates of the equation explaining revenues from imports

<u>Long-run equation: PHFMOLS estimates</u>

$\ln REVM$ =	-2.0232^{***}	$+\ 1.0470^{***}\ \ln IMP$	$-\ 0.32105^{***}D92$	$+\ 0.26281^{**}D00$
(s.e)	(0.435)	(0.03)	(0.08)	(0.09)
t-ratio	-4.18	32.06	-3.68	2.74

<u>Short-run error-correction model</u>

$\Delta \ln REVM$ =	0.035	$+\ 0.52^{***}\Delta \ln IMP$	$-\ 0.16^{***}D92$	$+\ 0.18^{***}D00$	$-\ 0.28^{**}RPHREVM_{t-1}$
(s.e)	(0.07)	(0.17)	(0.05)	(0.05)	(0.14)
t-ratio	0.47	3.05	-3.02	3.36	-2.00

Diagnostic Test

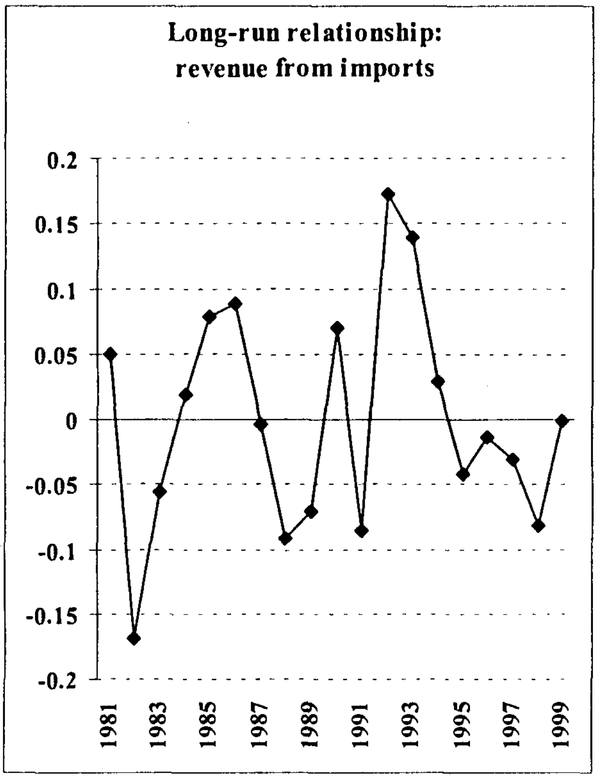
Adjusted R^2=0.57

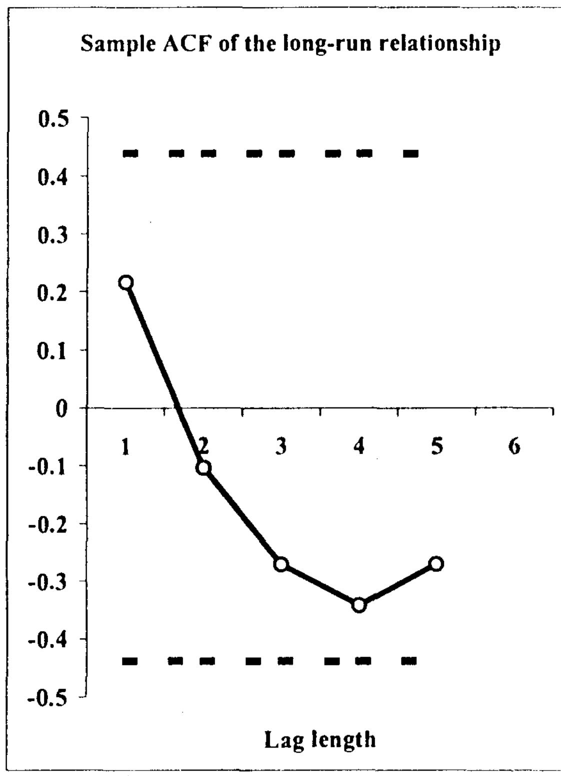
Serial Correlation: $\chi^2(1)$= 0.41	Functional Form: $\chi^2(1)$=2.39
Normality: $\chi^2(2)$= 0.73	Heteroscedasticity: $\chi^2(1)$=0.014

Note: Statistical significance at the one and five per cent levels are indicated by ***, and ** respectively. D92 is the dummy variable representing 0 for 1992 and 1 otherwise. Similarly, D00 represents 0 for 2000 and 1 otherwise.

The error-correction term ($RPHREVM_{t-1}$) in the short-run model is correctly signed and significant at the 5 per cent level but shows a relatively long period of adjustment. According to the estimated equation it takes about 4 years to return to steady state relationship from any short-run deviations.[49] Figure 5.23 exhibits the long-run relationship showing movement similar to a stationary variable and the sample autocorrelation coefficients of the relationship fall within the 95 per cent error-bands. However, the ADF test for residual stationarity fails, as the computed ADF statistic is -2.94 against its 5 per cent critical value of -3.34.

Figure 5.23: Long-run relationship in revenue from imports and its sample ACF





5.5.1.3. Empirical Estimation of Revenue from Internal Taxes

Table 5.21 gives the estimated equation for revenue from internal taxes. It is found that the long-run nominal income elasticity of internal taxes is 1.23 while the short-run estimate stood at 0.57. Both the long- and short-run models were estimated inserting a dummy variable for 1993, which turned out to be significant in both regressions. The error-correction term, $RPHREVIN_{t-1}$, is in line with the correct specification of the model and indicates an adjustment period of two years. The evidence for cointegration is, however, weak. It is very difficult to infer anything about the stationarity of the long-run relationship from its graphical plot, as given in Figure 5.24. Although the estimated autocorrelation coefficients of the long-

[49] Similar long period of adjustment was also associated with partial adjustment modelling technique.

run relationship lie within the 95 per cent error confidence level, rather than providing a random movement it shows a gradually declining trend, which is worrying. It might be that the simple specification used to model the revenue from internal taxes is not capturing important variation in the dependent variable. However, in this paper we will proceed further based on the simplified estimated relationship in Table 5.21.

Table 5.21: Estimates of the equation explaining revenues from internal taxes

Long-run equation: PHFMOLS estimates

LREVIN =	-6.60^{***}	$+ 1.23^{***}$ LNY	$- 0.19^{***}$D93
(s.e)	(0.31)	(0.022)	(0.05)
t-ratio	-21.21	56.56	-3.58

Short-run error-correction model

ΔLREVIN =	0.18^{***}	$+ 0.57^{**}$ ΔLNY	$- 0.099^{***}$D93	$- 0.53^{***}$RPHREVIN $_{t-1}$
(s.e)	(0.05)	(0.25)	(0.039)	(0.16)
t-ratio	3.62	2.15	-2.53	-3.28

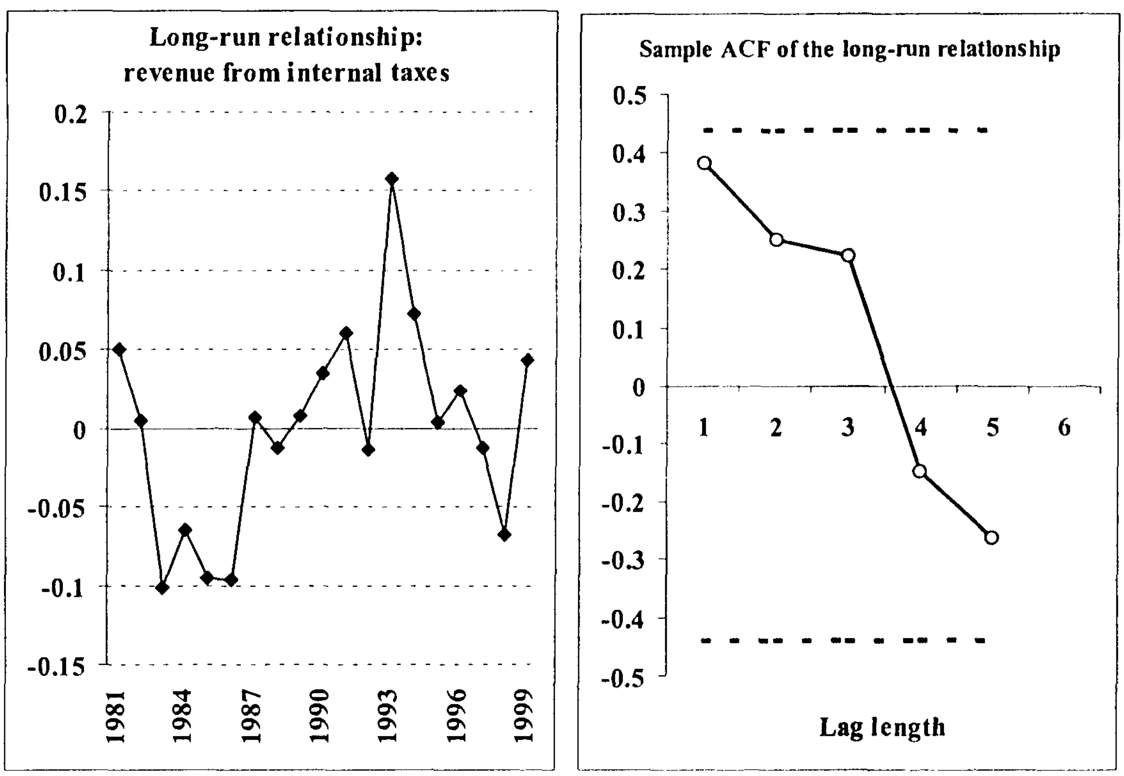
Diagnostic Test

Adjusted R^2=0.54

Serial Correlation: $\chi^2(1)$= 1.82	Functional Form: $\chi^2(1)$=1.25
Normality: $\chi^2(2)$= 0.99	Heteroscedastisity: $\chi^2(1)$=0.06

Note: Statistical significance at the one and five per cent levels are indicated by ***, and ** respectively. D93 is the dummy variable representing 0 for 1993 and 1 otherwise.

Figure 5.24: Long-run relationship in revenue from internal taxes and its sample ACF



5.5.1.4. Empirical Estimation of Revenue from Non-tax Sources

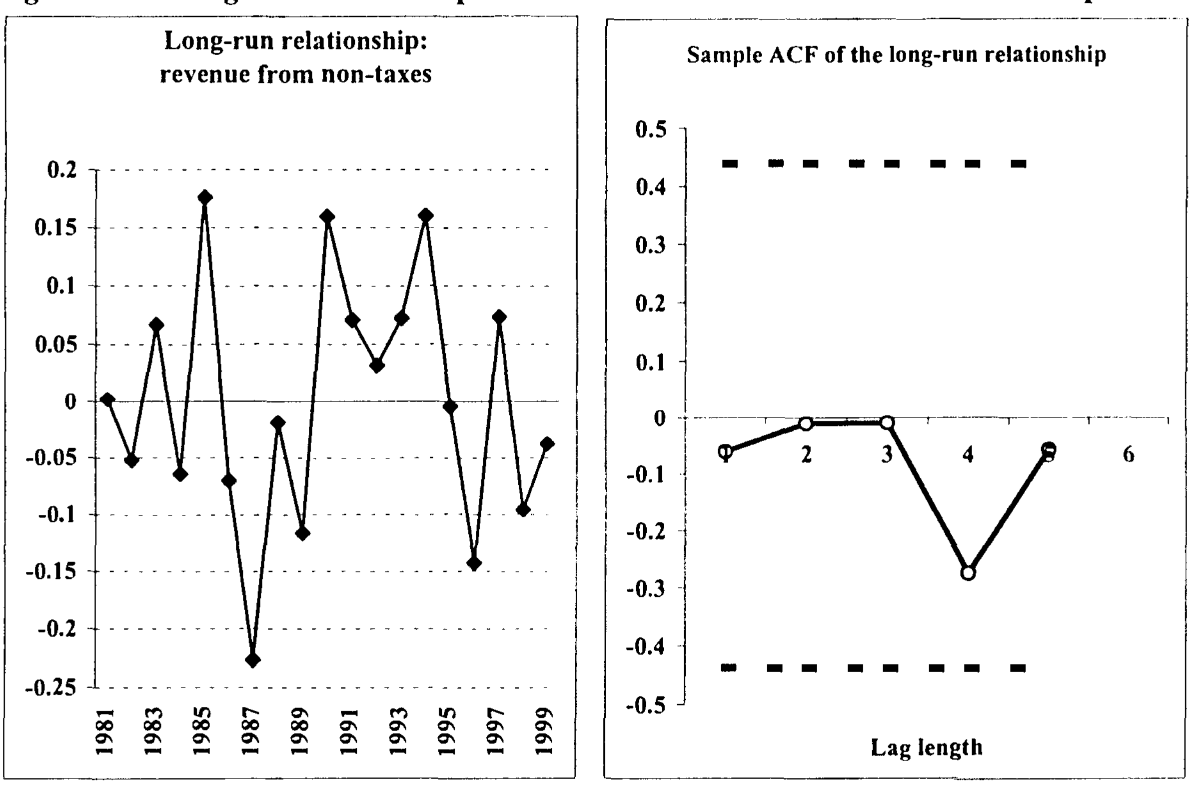
The Long-run steady state relationship for revenue from non-tax sources resulted in autoregressive errors and consequently the equation is estimated using the partial adjustment modelling technique and the PHFMOLS estimation procedure.

$\ln REVNT$ =	-2.6262^{**} +	$0.379^{**}\ln NY$ +	$0.73^{***}\ln REVNT_{t-1}$
(s.e.)	(1.26)	(0.17)	(0.128)
t-ratio	-2.07	2.20	5.70

Adjusted R^2=0.974

The estimated equation shows that in the short-run a one per cent increase in nominal income raises non-tax revenue by about 0.38 per cent. Given that the speed of adjustment is estimated to be 0.27, the long-run nominal income elasticity of REVNT is 1.40. The residual series from the above partial adjustment model seems to be stationary, which is shown in Figure 5.25. The computed ADF test statistic on the residual is estimated at –4.10 against the 95 per cent critical value of -3.74 thereby rejecting the null hypothesis of non-stationarity or non-cointegration.

Figure 5.25: Long-run relationship in revenue from non-tax sources and its sample ACF



5.5.2. Government Consumption Expenditure

5.5.2.1. Unit root test of the variables

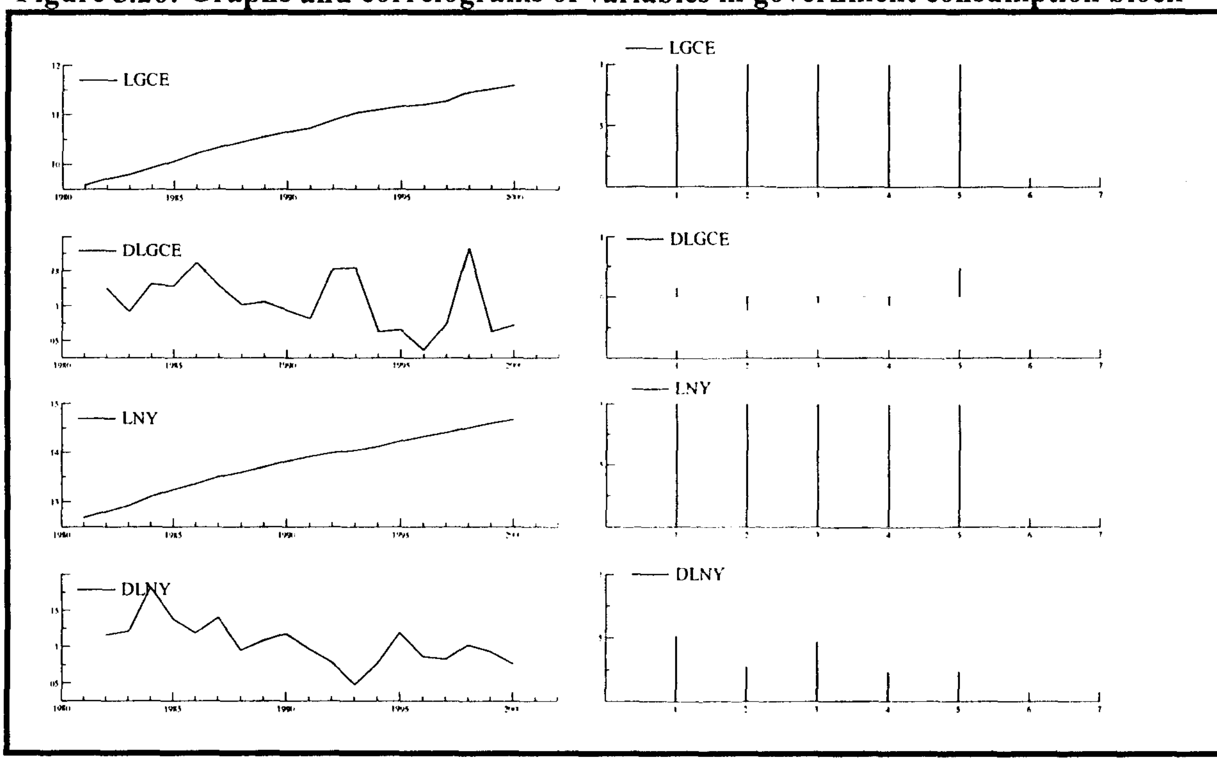
Government consumption expenditure (GCE) is modelled with a simple specification where it is a linear function of nominal national income (NY) alone. Table 5.22 gives the unit root test of lnGCE and lnNY. In the case of lnGCE, all the computed DF-ADF test statistics fall short of the corresponding critical values thus making the variable non-stationary on its level. Since the first difference of lnGCE is trended, on the basis of the ADF test with the trend ΔlnGCE can be treated as a stationary variable. That is, lnGCE ~ *I(1)*. The correlograms of lnGCE and ΔlnGCE, as presented in Figure 5.26, also seem to support this conclusion. On the other hand, the complex data generating process of lnNY was mentioned in the preceding section. Given that lnGCE is *~I(1)*, a valid cointegrating relationship is only possible if only lnNY is also *~I(1)*. Therefore, having estimated the model we should carefully examine the long-run relationship.

Table 5.22: Results of DF-ADF Test

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
lnGCE	-1.65	-1.49	-1.33	-1.46
ΔlnGCE	-3.24	-3.00	-3.75	-3.80
lnNY	-3.91	-2.65	-2.33	-2.31
ΔlnNY	-2.11	-1.74	-3.26	-3.20

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

Figure 5.26: Graphs and correlograms of variables in government consumption block



5.5.2.2. Estimating the Government Consumption Function

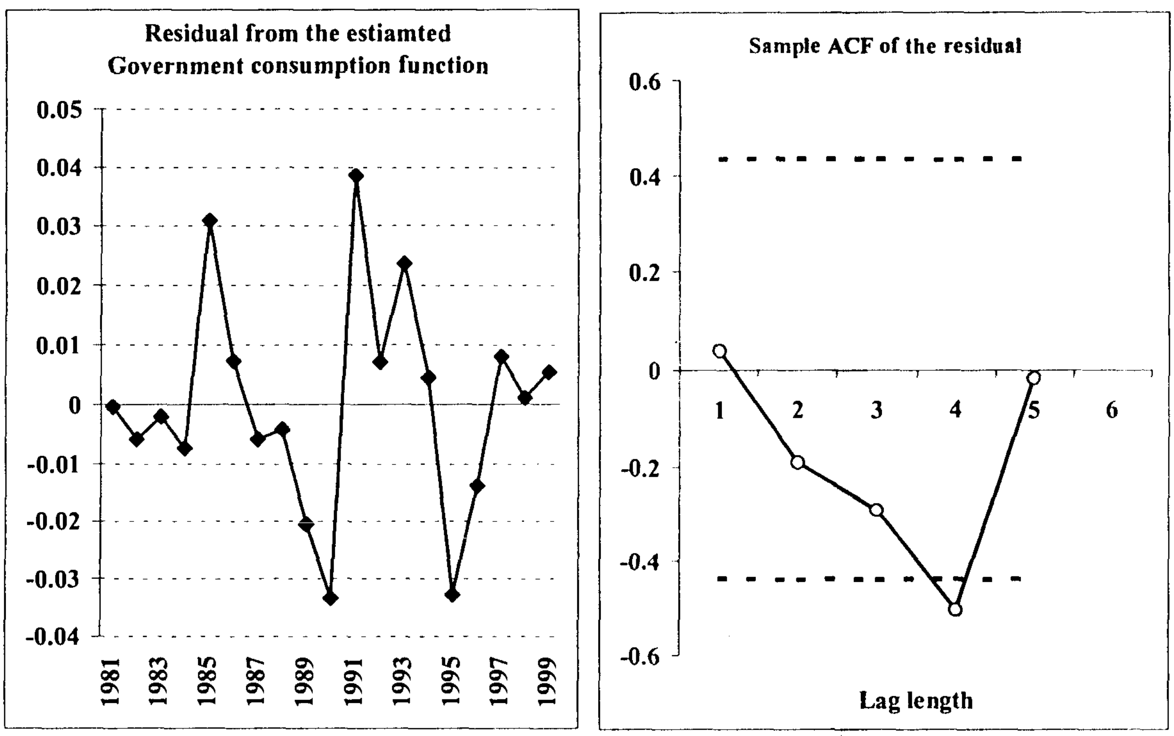
Estimation of the Engle-Granger first step static equation was confronted with a severe residual autocorrelation problem. Experiments suggested the inclusion of the lagged dependent variable in the model to ease the problem. Therefore, the following gives the estimated partial adjustment model of the government consumption expenditure.

$lnGCE = -0.988^{**}$	$+ 0.40^{***}$ lnNY	$+ 0.595^{***}$ $lnGCE_{t-1}$	$- 0.076^{***}$ D93	$- 0.075^{***}$D98	
(s.e)	(0.38)	(0.09)	(0.08)	(0.016)	(0.018)
t-ratio	-2.54	4.42	7.19	-4.56	-4.21

Adjusted $R^2 = 0.998$

It is obvious that the regression equation has been estimated using two dummies – one for 1993 and the other for 1998 – as the large unexplained variation associated with these two years resulted in the non-normality of residuals. All variables in the estimated equation are highly significant. The coefficient on the lagged dependent variable is also correctly signed justifying the use of partial adjustment mechanism. The short-run nominal income elasticity of government consumption expenditure is 0.40 and its long-run counterpart is estimated to be 0.99.

Figure 5.27: Long-run relationship in government consumption function and its sample ACF



Unfortunately, even estimation by using the partial adjustment model could not confirm a statistically valid relationship as there was hardly any evidence for residual stationarity. The graph of the residual (Figure 5.27) does not convincingly demonstrate a stationary pattern and the sample autocorrelation coefficient at the fourth lag length falls outside the 95 per cent error bands thereby rejecting the possibility of stationarity.

5.6. Monetary and Price Block

5.6.1. Money Supply

5.6.1.1. Integrating orders of variables in the equation

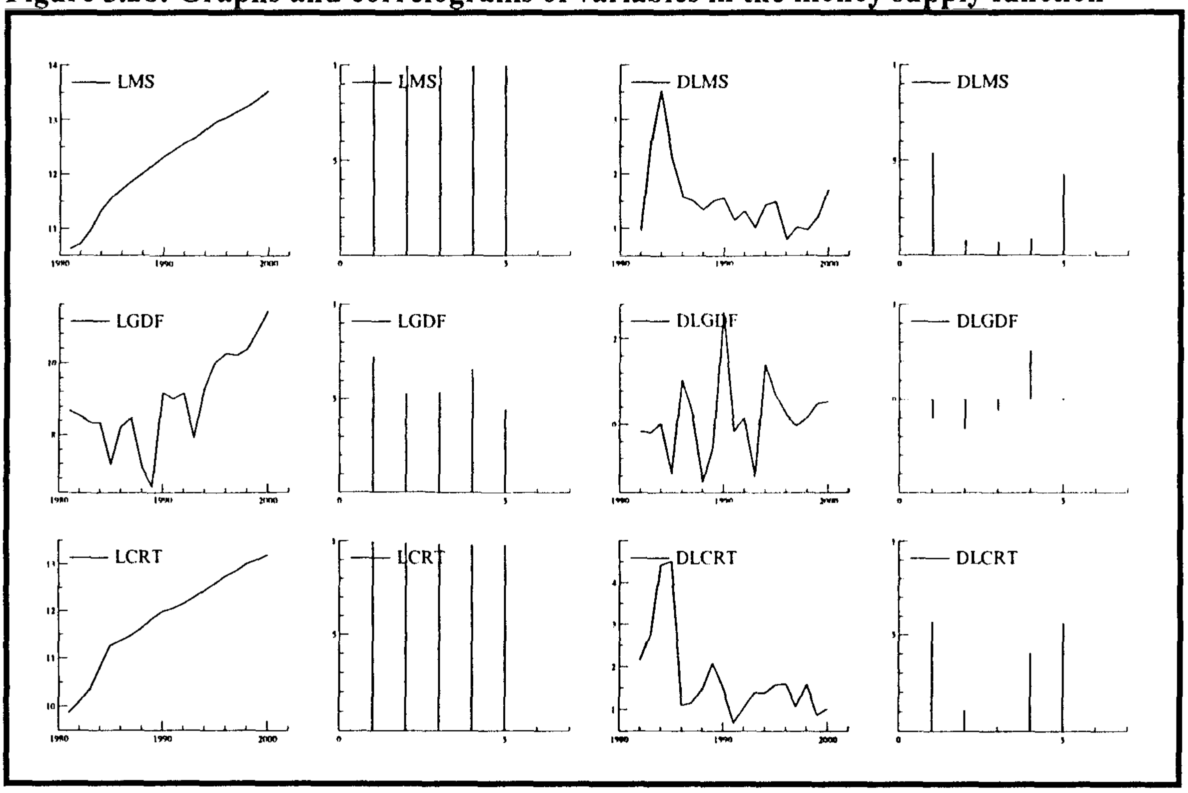
Money supply (MS) is considered to be a linear function of the budget deficit financed by government borrowing (GDF) and total credit to the private sector (CRT). The results of DF-ADF tests in Table 5.23 shows that only for lnGDF there is some conclusive evidence to suggest its order of integration to be *I(1)*, as all the test statistics on the level variables fall short of the critical values while the former exceed the latter in the case of the first differenced variable, ΔlnGDF. Such a conclusion is also supported by the graphical plots and correlograms, as given in Figure 5.28.

Table 5.23: Unit root test of the variables in the money supply equation

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
lnMS	-4.57	-3.25	-4.77	-5.71
ΔlnMS	-2.11	-3.11	-1.87	-3.49
lnGDF	-1.00	-0.65	-2.96	-2.61
ΔlnGDF	-4.81	-4.38	-4.90	-4.80
lnCRT	-3.85	-2.55	-3.29	-3.54
ΔlnCRT	-2.10	-2.44	-2.58	-3.42

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

Figure 5.28: Graphs and correlograms of variables in the money supply function



In the case of lnMS, the DF-ADF tests appear to reject the non-stationarity hypothesis on its level but not on its first difference counterpart, ΔLMS, making it a complicated variable. Correlograms of ΔlnMS also do not show any random movement in the autocorrelation coefficients. Finally, the integration order of lnCRT is also not clear as the ADF tests can reject unit root neither on its level nor first difference. The second difference of lnCRT was tested for stationarity without any success. The indeterminacy of the integration order of lnCRT is also reflected in the correlograms of Figure 5.28.

5.6.1.2. Estimation of the Money Supply Function

Table 5.24: Estimates of Money Supply Equation

Money Supply Equation: Long-run PHFMOLS estimates

lnMS =	1.797^{***} +	0.0382^{***} lnGDF +	0.853^{***} lnCRT
(s.e)	(0.157)	(0.012)	(0.017)
t-ratio	11.40	3.18	48.57

Adjusted R^2=0.996

Money supply function: Short-run estimates

ΔlnMS =	0.0517^{**} +	0.0177^{***} ΔlnGDF +	0.579^{***}ΔlnCRT -	0.30^{**} RPHMSUP $_{t-1}$
(s.e)	(0.02)	(0.0059)	(0.137)	(0.11)
t-ratio	2.49	3.00	4.20	-2.63

Diagnostic Test

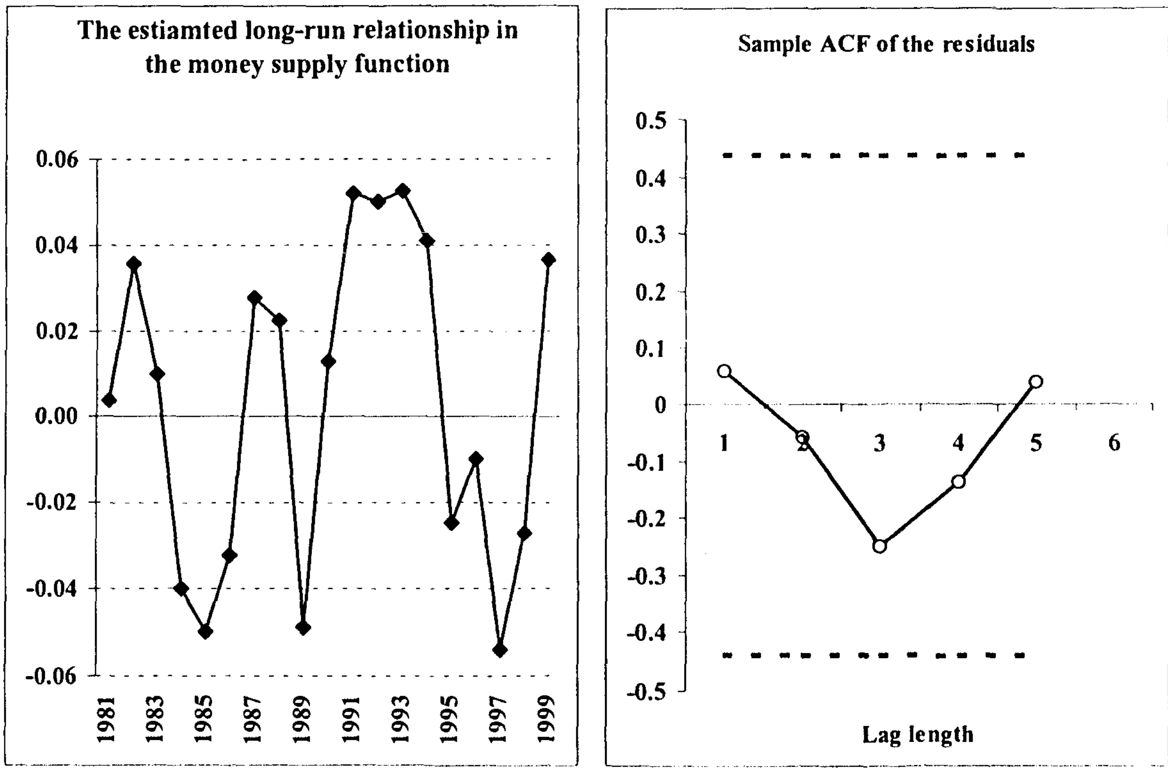
Adjusted R^2=0.687

Serial Correlation: $\chi^2(1)$= 0.56	Functional Form: $\chi^2(1)$=0.40
Normality: $\chi^2(2)$= 0.37	Heteroscedastisity: $\chi^2(1)$=0.37

Note: Statistical significance at the one and five per cent levels are indicated by ***, and ** respectively.

The top row in Table 5.24 provides the estimated long-run money supply function. All the variables in the estimated equation come out to be highly significant. A 10 per cent increase in budget deficit financed by government borrowing is found to be associated with about 0.4 per cent rise in money supply. However, credit given to the private sector has much stronger effect on money supply, as the estimated elasticity is 0.85. Although the integration orders of lnMS and lnCRT could not be determined, the cointegrating relationship, as presented in Figure 5.29, seems to be stationary. Individual sample autocorrelation coefficients (ACF) of the residuals upto the 5^{th} lag length are not statistically significant and all the ACFs fall within the 95 per cent error bands. Furthermore, the ADF test statistic on the cointegrating relationship was computed at –4.10 against its 95 per cent critical value of -3.74. All this implies that we have found a valid long-run cointegrating relationship in the money supply equation.[50]

Figure 5.29: The long-run relationship in money supply equation and its sample ACF



The bottom row in Table 5.24 gives the short-run money supply equation. This is estimated following the error-correction modelling technique. In the short-run too all variables turn out to be statistically significant and the short-run government deficit and private sector credit elasticities are estimated to be respectively 0.018 and 0.58. The error-correction term, $RPHMSUP_{t-1}$, is correctly signed and shows that it takes more than 3 years for short-run deviations to converge to the long-run steady-state path. The correct sign and the significance of the error-correction term is indicative of a valid relationship in the first step long-run equation.

[50] The stationarity of the residuals will actually imply that either LMS and LCRT are individually or their linear combination is ~I(1) as LGDF is found to be ~I(1).

5.6.2. Inflation

The last estimating equation in the model is the one for inflation. It was previously specified that the rate of inflation be determined by the rate of growth of money supply and the rate of growth of income. Statistically, however, estimating an equation on the growth rates of variable causes a serious problem, as it is tantamount to wiping out all long-run relationship. This is despite the fact that many time series are non-stationary and transformation in terms of growth rate makes them stationary. In the equation for inflation, therefore, it is essential to see whether the level variables, i.e., the price level (P), money supply (MS) and income (Y) are cointegrated and then the appropriate short-run model can capture the relationship in terms of their growth rates.[51]

5.6.2.1. Unit root test of variables

Table 5.25 provides the results of the unit root tests involving the three variables on their levels and first differences. For lnY and lnMS the tests were actually carried out in previous sections. In section 5.5 it was argued that lnY was most likely to be *~I(1)*, while in section 5.8 the problem concerning the determination of the order of integration of lnMS was discussed. Thus, lnP is the only variable which needs to be tested for unit root. Clearly, the DF-ADF test results for lnP are ambiguous and cannot reject the null of unit root even for ΔlnP. The graph of ΔlnP is strongly downward trended and its correlograms (see Figure 5.30) do not give random movement as expected for a stationary variable. The second difference of lnP also failed to reject the null hypothesis under the DF-ADF tests making it extremely difficult for us to establish its order of integration. Since the unit root properties of lnMS and lnP are also dubious, cointegration can only be found if the residuals from the long-run relationship become stationary.

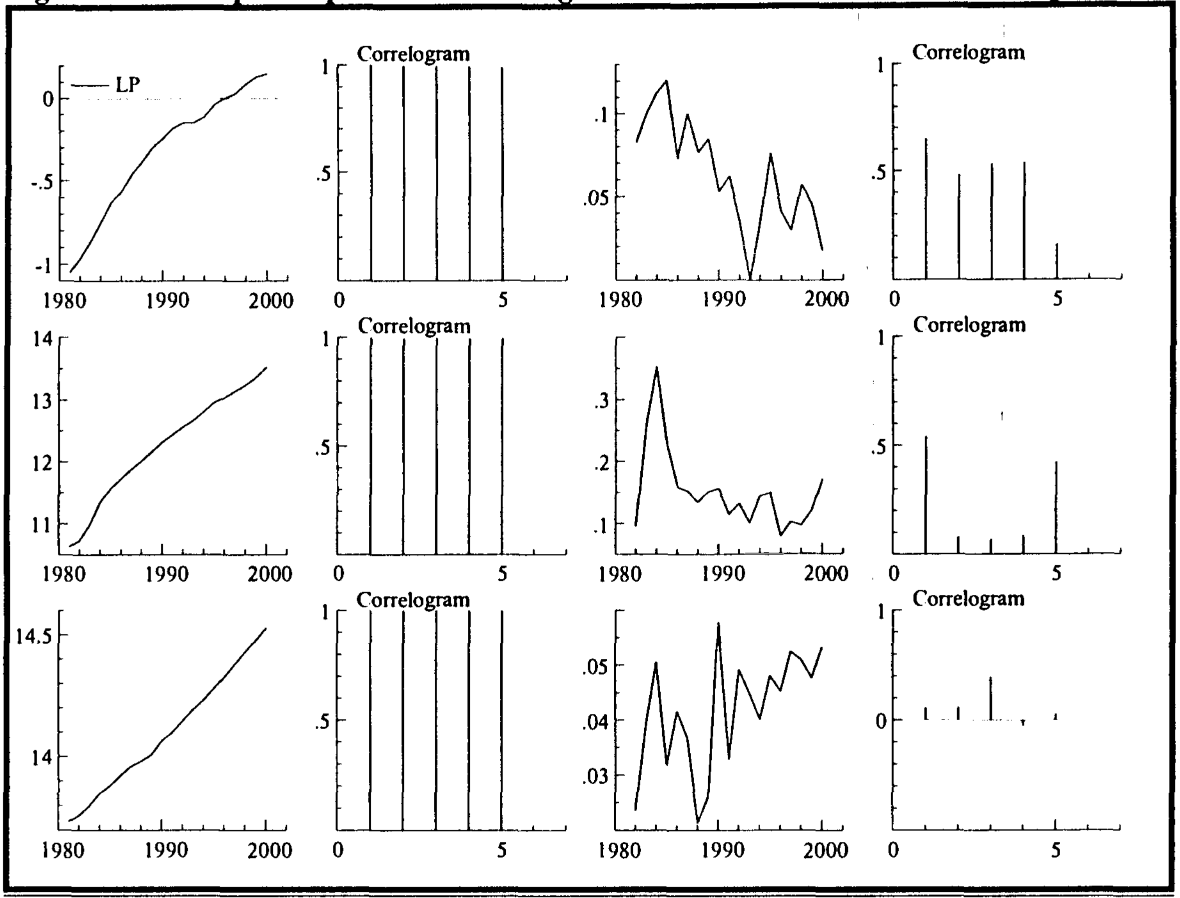
Table 5.25: DF-ADF tests of variables in the inflation equation

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
lnP	-5.70	-3.48	-2.06	-1.99
ΔlnP	-1.72	-1.32	-3.07	-2.81
lnMS	-4.57	-3.25	-4.77	-5.71
Δ lnMS	-2.11	-3.11	-1.87	-3.49
lnY	2.18	2.24	-0.30	-0.04
ΔlnY	-3.43	-2.28	-4.41	-3.46

Note: The critical value for the DF-ADF tests without the trend term is –3.04 while the comparable value for the regression including the trend terms is –3.69.

[51] Note that when variables are transformed on the logarithmic scale, their first differences indicate growth rate.

Figure 5.30: Graphical plots and correlograms of variables in the inflation equation



5.6.2.2. Estimating the price level and inflation equation

The PHFMOLS estimate of the static long-run price equation, given in the top row of Table 5.26, shows that a one per cent increase in money supply increases the price level by 0.53 per cent while the same change in output (lnY) depresses the price by 0.43 per cent. Both the signs of the explanatory variables are theoretically plausible and the estimated coefficients are highly significant.

Table 5.26: Long-run and short-run estimates of the price equation

Long-run equation: PHFMOLS estimate

lnP =	- 0.781	+ 0.535^{***} lnMS	- 0.43^{**} lnY
(s.e)	(1.25)	(0.03)	(0.12)
t-ratio	-0.62	15.65	-3.67

Short-run error-correction equation: OLS estimate

ΔlnP =	0.038^{**}	+ 0.29^{***} ΔlnMS	-1.26^{***} ΔlnY	+ 0.0368^{***} D93	- 0.38^{***} $RPHP_{t-1}$
(s.e)	(0.017)	(0.048)	(0.33)	(0.009)	(0.12)
t-ratio	2.17	6.01	-3.84	3.70	-3.15

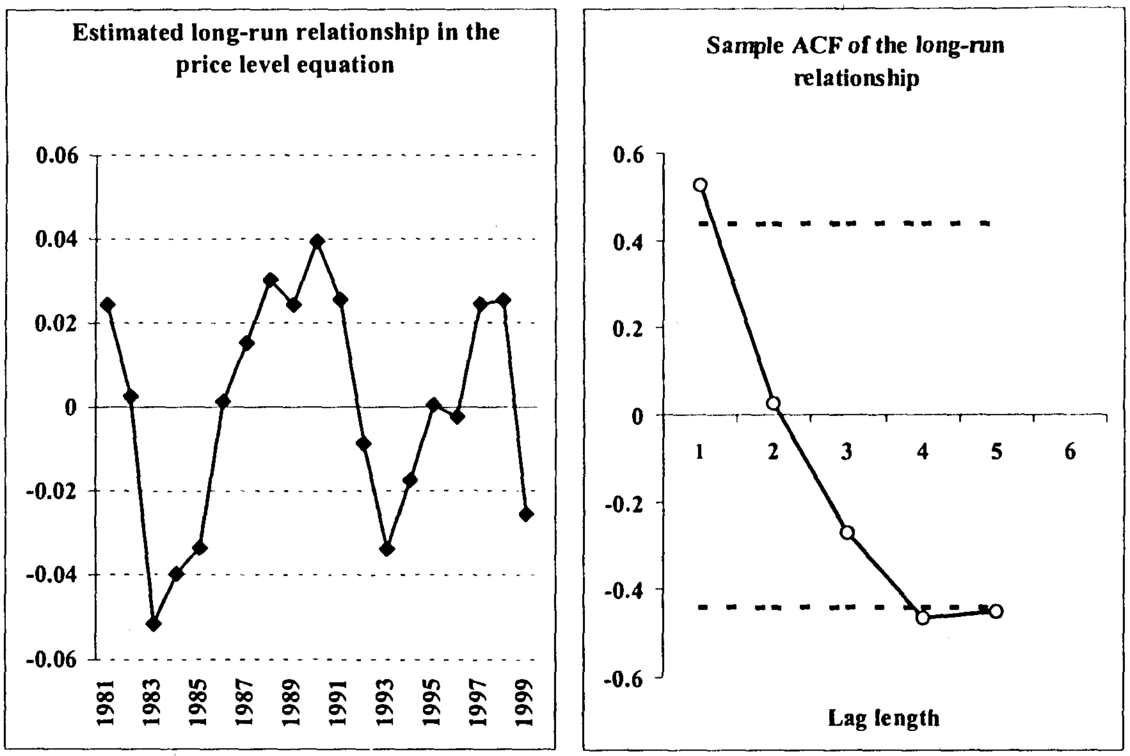
Diagnostic Test

Adjusted R^2=0.85

Serial Correlation: $\chi^2(1)$= 1.21	Functional Form: $\chi^2(1)$=0.004
Normality: $\chi^2(2)$= 3.17	Heteroscedastisity: $\chi^2(1)$=0.72

Note: Statistical significance at the one and five percent levels are indicated by ***, and ** respectively.

Figure 5.31: Long-run relationship in the price level equation and its sample ACF



Unfortunately the estimation of the long-run equation could not provide very strong evidence for cointegration. The ADF test statistic for residual stationarity was found to be –3.04 as against its 95 per cent critical value of –3.34 thus upholding the non-cointegration hypothesis. The graph of the residuals (Figure 5.31) also does not seem to support its stationarity and the sample autocorrelation coefficients at the first, fourth and fifth lag lengths were individually significant and fell outside the 95 per cent error bands – all characterizing a non-stationary variable.

Despite the lack of evidence for cointegration, the short-run error correction model corresponding to the above-mentioned long-run relationship is estimated in the bottom row of Table 5.25. Like the long-run model, the short-run money supply growth is positively and output growth is inversely related with the growth in price level. The short-run output elasticity emerges out to be significantly higher than the similar effects in the long-run equation. This might suggest that in the context of Bangladesh the 'demand-pull' has dominated the movement in the price level and an increase in output supply has a very big influence on the changes in the price level. Interestingly, as a sharp contrast to the non-cointegration result the error-correction term, $RPHP_{t-1}$, appears to be correctly signed and highly significant. In the literature often a well-behaved (i.e., significant, negatively signed, and a value less than 1 absolutely) error-correction term is considered to be an evidence for a valid long-run relationship, in which case our short-run results contradict with the finding of the cointegration test.

Since our objective is to estimate an equation for inflation, i.e., the rate of change in the price level, and not the price level as such, a convenient way of modelling it might be to follow the unrestricted error-correction modelling technique (Banerjee *et al.* 1993). The estimation of an UECM involves the estimation of just one equation that regresses first difference of the dependent variable on the differenced variables and the lags of the variables on their levels (instead of the lag of the residuals in the Engle-Granger second step). The unrestricted model in our case can be specified as:

$$\Delta \ln P = C + \beta\, \Delta \ln MS + \gamma\, \Delta \ln Y + \Gamma_1 \ln P_{t-1} + \Gamma_2 \ln MS_{t-1} + \Gamma_3 \ln Y_{t-1} + error$$

From the above equation the short-run parameters are directly obtained as the coefficient on the differenced variables (i.e., the estimates of β and γ in this case). The long-run solution, on the other hand, is obtained by noting that in equilibrium there is no change in the steady state, so that $\Delta \ln P = \Delta \ln MS = \Delta \ln Y = 0$. Therefore:

$$0 = C + \Gamma_1 \ln P + \Gamma_2 \ln MS + \Gamma_3 \ln Y$$

And thus: $\ln P = -\left[\dfrac{C}{\Gamma_1}\right] - \left[\dfrac{\Gamma_2}{\Gamma_1}\right] \ln MS - \left[\dfrac{\Gamma_3}{\Gamma_1}\right] \ln Y$

The coefficient on lnMS and lnY will then give the long-run coefficients. Following the "general to specific" approach to modelling we first constructed a general model by incorporating first lag of the differenced variables and then eliminated the insignificant terms. The resultant parsimonious UECM is given in Table 5.27.[52]

Table 5.27: Modelling inflation by Unrestricted Error Correction Model

$\Delta LP =$	$0.042 +$	$0.236^{***}\Delta LMS$ -	$0.89^{**}\Delta LY$ +	$0.037^{***}D93$ +	$0.20^{***} LMS_{t-1}$ -	$0.21^{**}LY_{t-1}$ -	$0.364^{***}LP_{t-1}$
(s.e)	(0.68)	(0.064)	(0.38)	(0.009)	(0.06)	(0.078)	(0.11)
t-ratio	0.61	3.68	-2.29	3.90	3.17	-2.68	-3.21

Diagnostic Test

Adjusted R^2=0.86

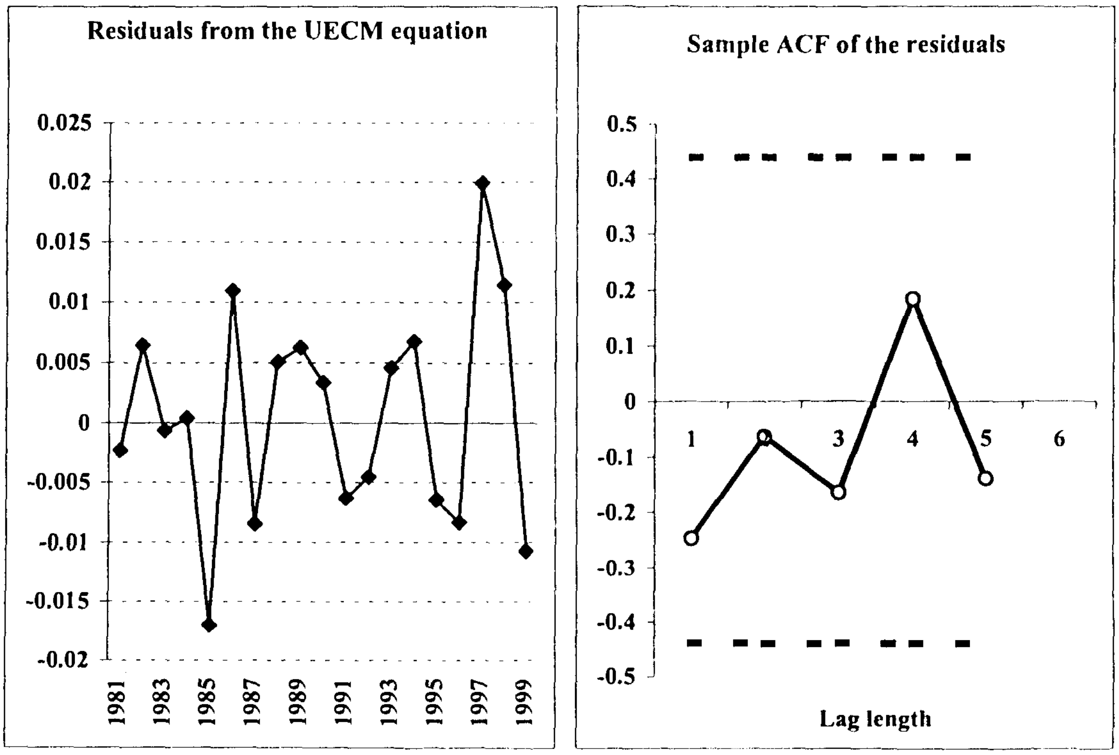
Serial Correlation: $\chi^2(1)$= 2.19	Functional Form: $\chi^2(1)$=0.13
Normality: $\chi^2(2)$= 0.19	Heteroscedastisity: $\chi^2(1)$=1.25

Note: Statistical significance at the one and five percent levels are indicated by ***, and ** respectively. D93 is a dummy variable with 0 for 1993 and 1 otherwise.

[52] As evident from Table 5.26 the estimated UECM model is estimated with a dummy for 1993. The dummy variable was required to capture some unexplained movement in that particular year. Note that such a dummy was not required in estimating the earlier static steady-state relationship but was included in the short-run error-correction model of Table 5.25. Since the UECM in Table 5.26 uses the same dependent variable as in Table 5.25, it is no surprise that the 1993 dummy is again required.

It is now obvious that the short-run money supply and output elasticities are respectively 0.24 and –0.89. The corresponding long-run elasticities are estimated to be 0.56 and -0.58. In the UECM the coefficient on the lagged level dependent variable is the error-correction term, which is almost identical to the one estimated in Table 5.26. The UECM model explains 86 per cent variation in the rate of inflation and diagnostics do not point toward any problem. The residuals from the UECM regression look a like stationary variable and the sample ACF fall well within the critical bands (Figure 5.32). Therefore, despite the indeterminate orders of integration for two of the variables, stationarity of the residuals is obtained. This means the variables in our long-run model cointegrate. Under the UECM the test for cointegration can be *directly* undertaken by employing the Pesaran *et al.* (2001) test. In the present case the computed F-statistic was found to be 5.15, which exceeded the critical upper bound critical value of 4.34 thus providing evidence for cointegration.

Figure 5.32: Residuals from the UECM equation and sample ACF



6. Conclusion

The basic objective of this paper has been to develop a small macroeconometric model for Bangladesh in order to capture, amongst others, the nexus between output, government deficit, external sector, money supply, and price level. It is envisaged that the model can be used to examine the effects of both domestic and external shocks to the economy. The model developed here first disaggregates the economic activities into 6 blocks viz., (i) production, (ii) investment, (iii) private consumption expenditure, (iv) external sector, (v) government and (vi) monetary and price blocks. Then under each block equations have been formulated to explain the variation in the relevant endogenous variables. While the present paper provides the specification of various sectoral equations and their econometric estimates, the tasks of model validation and policy simulations based on the estimated equations are undertaken in another paper.

There are several distinctive features of the modelling exercise that has been undertaken in this paper. It uses the recently published revised national income accounting data of the BBS and, therefore, the estimates of the sectoral equations are themselves important to know how various theoretical specifications fit the new data. The supply side of the economy has been given special importance in this paper by specifying separate equations for value added in agriculture, manufacturing and services. Similarly, one important component of the demand side in the economy, private investment demand, has been modelled separately for agriculture, manufacturing, and services sectors along with the private consumption and government consumption expenditures for the overall economy. Thus, the present paper also captures the important dynamics of the demand side of the economy.

Another important feature of the present exercise has been to explicitly consider the time series properties of the data and to use estimation techniques suitable for dealing with the non-stationary data. As most of the variables appear to be non-stationary on their levels, cointegration techniques have been used to verify the long-run relationship as postulated in the theoretical specifications. The short-run equations are estimated following the error-correction modelling technique, which captures the short-run dynamics considering the long-run information. Although the determination of the integrating order of variables was somewhat an involved task, partly because of a relatively short sample size, convincing evidence for valid long-run relationship was obtained for a majority of the equations.

To conclude, it might be useful to summarize some of the interesting results that come out of the present paper.

(1) Sectoral capital stocks are found to have significantly influenced the long-run production of agriculture, manufacturing, and services sectors.[53] The effect of capital stock on output is the lowest in agriculture (the elasticity is 0.26), followed by manufacturing (with an elasticity of 0.90) and services (with an elasticity of 1.73). While agricultural production continues to depend on such factors as irrigated area and weather, additional investible resources leading to the growth of capital stocks in manufacturing and services sectors hold great promises for output growth.

(2) Private investment demands in agriculture, manufacturing and services sectors appear to have been subject to structural break as the estimated relationships for the 1990s are at variance with those of the 1980s. For the 1990s credit given to the private sector is positively related to private investment in each sector. In the manufacturing sector the effect of credit is found to be the highest with an elasticity of 1.15 compared to such comparable elasticities of 0.36 and 0.79 for agriculture and services sector, respectively. In the 1990s public investment has been complimentary to private investment in the agriculture and services sectors but appears to have crowded out private sector contribution in manufacturing. Why the effect of public investment on private investment behaviour in manufacturing has been different from other sector is an issue worth investigating in future research.[54]

(3) One of the most striking findings of the present paper is associated with the estimate of the income elasticity of private consumption expenditure. A structural break in the relationship between private consumption expenditure and disposable income is noticed as the new national accounts data reveals that compared to the 1980s there has been a drastic decline in the ratio of private consumption expenditure to total disposable income in the 1990s. The average propensity of consumption in the 1980s stabilized just above 0.85, but in the following decade it came down dramatically to less than 0.7. This results in an income elasticity of private consumption expenditure

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[53] This result is important as there no study in Bangladesh that provides estimates for disaggregated production functions for agriculture, manufacturing and services sectors.

[54] Note that such an investigation is beyond the scope of the present study.

for the 1990s to as low as 0.69 in comparison with as high as 0.85 for the 1980s. Further research is needed to understand why elasticity of consumption fell so rapidly and whether and how this trend can be consistent with the development of a low-income country like Bangladesh.

(4) In the foreign trade block, the estimated equation for supply function of jute and jute goods encountered a structural break. Consequently, the price elasticity of supply for the 1980s was estimated to be 1.09 but for the 1990s a perverse response was observed (with an elasticity of –0.15). There have been significant changes in the demand for jute and jute goods as synthetic products have replaced the use of the former to a great extent. It might be that because of the declining demand for jute, the supply response to price changes is very low. The capacity utilization ratio in the estimated equation failed to register statistical significance, which was consistent with the excessive supply capacity in the industry. In the case of RMG, the estimation of the equation was constrained by a very short sample size because of lack of data on export prices. However, the long-run price elasticity of supply of RMG is computed to be 0.85 and the capacity utilization elasticity comes out as 1.27. It is worth mentioning here that the market for RMG is highly regulated and that some significant import liberalization measures in the developed countries are currently being undertaken may help explain this phenomenon. The Gradual removal of quantitative restrictions mainly in the EU and North American markets has resulted in increased competition generating a downward pressure on prices. Hence, further liberalization of RMG trade regime in future might bring in dramatic changes to the estimated relationship. In the case of other export goods, the long-run price elasticity of supply is found to be 1.47. The elasticity with respect to capacity utilization with regard to this export category is very high: 3.2.

(5) The estimated import demand function for intermediate goods and raw materials revealed highly price inelastic nature of the demand for these products. This finding is in line with the general perception that production in the domestic economy is critically dependent on imported inputs and that there has been a very low backward integration into Bangladesh's economy. For all other importables, the demand was found to be fairly price elastic. Both types of imports were also income elastic.

(6) A one-to-one correspondence between total imports and revenue from imports is found. Revenue from internal taxes and non-tax sources are also highly responsive to rise in nominal income; in both cases the long-run elasticity being greater than one. The government consumption expenditure elasticity with respect to nominal income is estimated to be almost 1.

(7) In the monetary and price block the regression results are satisfactory as well. Both the government budget deficit and the credit pumped into the economy through the private sector were found to be influencing the money supply significantly. Finally, in the equation for price level the money supply acted as a contributing factor to inflation while increased aggregate output exerted a downward pressure on it.

It follows from the above that the estimated sectoral equations in general yield results that are plausible. These results also identify a number of issues (e.g., structural breaks, low income elasticity of consumption, low price response of exports,) that may require further research. When the objective is to build a macroeconomic model there exists hardly any scope for in-depth investigation into a particular sector. Nevertheless, the present study has unearthed a number of important macroeconomic dynamics in the context of Bangladesh and shows that applied research with the revised national income accounting data generates results that are plausible and theoretically consistent.

References

Ahmed, S. M., Hoque, M.N. and Talukder, M.S.I. (1993), "Estimating Export Demand Function for Bangaldesh: An Application of Cointegration and Error-Correction Modelling", *Bangladesh Development Studies*, vol.21, no.4, pp.89-104.

Athukorala, P.C. and Riedel, J. (1994), "Demand and Supply Factors in the Determination of NIE Exports: A Simultaneous Error-Correction Model for Hong Kong: A Comment" *Economic Journal*, vol.104, pp.1411-14.

Athukorala, P.C. and Riedel, J. (1996), "Modelling NIE Exports: Aggregation, Quantitative Restrictions and Choice of Econometric Methodology", *Journal of Development Studies*, vol.33, no.1, pp.81-98.

Banerjee, A., Dolado, J.J., Galbraith, J.W. and Hendry, D.F.(1993), *Cointegration, Error-correction, and the Econometric Analysis of Non-stationary Data*, Oxford University Press.

Bangladesh Bank (various years), *Economic Trends*, Statistics Department, Bangladesh Bank, Dhaka.

Bangladesh Bureau of Statistics (BBS), *The Statistical Yearbook of Bangladesh*, Dhaka, Bangladesh, Various Issues.

Bangladesh Bureau of Statistics (BBS) (2001), *National Accounts Statistics: Gross Domestic Product 2000-2001*, June, BBS.

Bangladesh Bureau of Statistics (BBS) (2000), *National Accounts Statistics of Bangladesh (Revised Estimates, 1989-90 to 1998-99)*, December, BBS.

Bayes, A., Hussain, I. And Rahman, M. (1995) "Trends in the External Sector: Trade and Aid", Chapter 8 in Sobhan, R. (ed.) *Experiences with Economic Reforms: A Review of Bangladesh's Development 1995*, Centre for Policy Dialogue and University Press Limited, Dhaka, pp.243-297.

Bhuyan, A. R. and Rashid, M.A. (1993), *Trade Regimes and Industrial Growth: A Case Study of Bangladesh*, Bureau of Economic Research, Dhaka University and International Centre for Economic Growth, San Francisco, USA.

Charemza, W.W. and Deadman, D.F. (1992), *New Directions in Econometric Practice*, Edward Elgar, England.

Dickey, D.A. and Fuller, W.A. (1979), "Distribution of the Estimators for Autoregressive Time Series With a Unit Root", *Journal of the American Statistical Association*, vol. 74, pp.427-81.

Dickey, D.A. and Fuller, W.A. (1981), "Likelihood Ratio Statistics for Autoregressive Time Series With A Unit Root", *Econometrica*, vol. 49, pp. 1057-72.

Doornik, J.A. and Hendry, D. (1997), *Modelling Dynamic Systems Using PcFiml 9 for Windows*, London: Timberlake Consulting.

Engle, R.F. and Granger, C.W.J. (1987), "Cointegration and Error Correction: Representation, Estimation and Testing", *Econometrica*, vol. 55, pp.251-76.

Engle, R.F. and Yoo, S. (1987), "Forecasting and Testing in Cointegrated Systems", *Journal of Econometrics*, vol.35.

Granger, C.W.J. (1969), "Investigating Causal Relations by Econometric Models: Cross Spectral Methods", *Econometrica*, vol.37, pp.424-38.

Granger, C.W.J. (1988), "Some Recent Development in A Concept of Causality", *Journal of Econometrics*, vol.39, pp.199-211.

Granger, C.W.J. and Newbold, P. (1974), "Spurious Regressions in Econometrics", *Journal of Econometrics*, vol.2, pp.111-20.

Gujarati, D. (1995), *Basic Econometrics*, Third Edition, McGraw Hill.

Hall, S.G. (1986), "An Application of the Granger and Engle Two-Step Estimation Procedure to the United Kingdom Aggregate Wage Data", *Oxford Bulletin of Economics and Statistics*, vol.48, p.238.

Hall, S.G. (1991), "The Effects of Varying Lag Length VAR Models on the Maximum Likelihood Estimates of Cointegrating Vectors", *Scottish Journal of Political Economy*, vol.38, pp.317-23.

Harris, R.(1995), *Using Cointegration Analysis in Econometric Modelling*, Pentice Hall and Harvester Wheatsheaf.

Hendry, D. (1995), *Dynamic Econometrics*, Oxford University Press.

Hendry, D. (1999), "An Econometric Analysis of US Food Expenditure, 1931-1989", Chapter 17 in Magnus, J.R. and Morgan, M.S. (eds), *Methodology and Tacit Knowledge: Two Experiments in Econometrics*, John Wiley and Sons.

Hendry, D. and Richard, J.F. (1983), "The Econometric Analysis of Economic Time Series", *International Statistical Review*, vol.51, pp.3-33.

Holden, D. and Perman, R. (1994), "Unit Roots and Cointegration for the Economists", in B. Bhaskara Rao (ed.) *Cointegration for the Applied Economists*, St. Martin's Press.

Hossain, M.Akhtar (1995), *Inflation, Economic Growth and the Balance of Payments in Bangladesh: A Macroeconometric Study,* Oxford University Press, Delhi.

Hussain, I., Rahman, A. and Rahman, M. (1997), "Current External Sector Performance and Emerging Issues", Chapter 6 in Sobhan, R. (ed.) *Growth or Stagnation? A Review of Bangladesh's Development 1996*, Centre for Policy Dialogue and University Press Limited, Dhaka, pp.161-219.

Islam, Nurul (1965), *A Short-Term Model for Pakistan Economy: An Econometric Analysis,* Oxford University Press, Dhaka.

Johansen, J. (1988), "Statistical Analysis of Cointegrating Vectors", *Journal of Economic Dynamics and Control*, vol.12, pp.231-54.

Johansen, S. and Juselius, K. (1990), "Maximum Likelihood Estimation and Inference on Cointegration With Application to the Demand For Money", *Oxford Bulletin of Economics and Statistics*, vol.52, pp.169-210.

Kabir, H. (1997), "Consumption Functions for Bangladesh: A Time Series Analysis", *Social Science Review*, vol.XIV, No.1, pp.191-204.

MacKinnon, J. (1991), "Critical Values for Cointegration Test", in Engle, R. and Granger, C. (eds) *Long-Run Economic Relationshis: Readings in Cointegration*, Oxford University Press.

Maddala, G.S. (1992), *Introduction to Econometrics*, 2nd Edition, McMillan: New York.

Mallick, S.K. (1999), *Modelling Macroeconomic Adjustment With Growth in Developing Economies: The Case of India*, Ashgate.

Musila, J.W. (2002), "An Econometric Model of the Malawian Economy", *Economic Modelling*, vol.19, pp.295-330.

McKinnon, J. (1991), "Critical Values of Cointegration Tests", Chapter 13 in Engle R.F. and Granger, C.W.J. (eds), *Long-Run Economic Relationships: Readings in Cointegration*, Oxford University Press.

Panagariya, A., Shah, S. and Mishra, D. (2001), "Demand Elasticities in International Trade: Are They Really low?" *Journal of Development Economics* vol.64, pp. 313-342 and also *Policy Research Working Paper 1712*, World Bank, Country Operations Division, Country Department I, South Asia, 1996.

Pesaran, M. H., Shin, Y. and Smith, R.J. (2001), "Bounds Testing Approaches to the Analysis of Long Run Relationships", *Journal of Applied* Econometrics, Special Issues in Honour of J.D. Sargan on the Theme, *Studies in Empirical Macroeconometrics*, edited by D.F. Hendry and M.H. Pesaran, pp.289-326 and also University of Cambridge DAE Working Paper No.9907, Accessed from http://www.econ.cam.ac.uk/faculty/pesaran/pss1r1.pdf.

Phillips, P. (1995), "Fully Modified Least Squares and Vector Autoregression", *Econometrica*, vol.63, 1023-78.

Phillips, P.C.B. and Hansen, B. (1990), "Statistical Inference in Instrumental Variables Regression With I(1) Processes", *Review of Economic Studies*, vol.57, pp.99-125.

Phillips, P.C.B. and Loretan, M. (1991), "Estimating Long-Run Economic Equilibria", *Review of Economic Studies*, vol.58, pp.407-36.

Rahman, Sultan H. and F. J. Shilpi (1996), *A Macro Econometric Model of the Bangladesh Economy: Model, Estimation, Validation and Policy Simulation,* Research Monograph No. 17, Bangladesh Institute of Development Studies, Dhaka.

Rashid, M.Ali (1981), A Macro Econometric Model of Bangladesh, *Bangladesh Development Studies* , Vol. IX, No. 3.

Razzaque, A. and Ahmed, N. (2000), "A Reexamination of Domestic Saving-Foreign Aid Relationship in the Context of Bangladesh", *Bangladesh Development Studies*, vol.26, no.4., pp. 1-37.

Razzaque, M.A. (2002), "A Reassessment of Small Country Assumption in World Trade: Some Implications for Bangladesh's Exports", Chapter 5 in *Essays on Bangladesh's Exports*, Unpublished PhD thesis, University of Sussex.

Senhadji A. and Montenegro, C. (1998), "Time Series Analysis of Export Demand Equations: A Cross-Country Analysis", *IMF Working Paper*, WP/98/149.

Senhadji, A. (1998), "Time Series Estimation of Structural Import Demand Equations: A Cross-Country Analysis", *IMF Staff Papers*, vol.45, no.2, pp.236-267.

Shilpi, F.J. (1990), *Estimating Income and Price Elasticities of Imports and Exports of Bangladesh*, Research Report Number 122, Bangladesh Institute of Development Studies, Dhaka.

White, H. (1980), "A Heteroscedascity Consistent Covariance Matrix Estimator and a Direct Test of Heteroscedasticity", *Econometrica*, vol.48, pp.817-818.