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Abstract: Amid historically high food prices agriculture has made a comeback to the centre stage supported by donor resources in the backdrop of national governments' renewed emphasis on achieving food security for their populations. While this reinvigorated approach to agricultural development is attractive as a majority of poor people still rely on farm activities for their livelihoods with poverty reduction remaining the most important policy objective, there are concerns about low productivity of the sector along with its role in the process of development. This paper sheds some light on this debate by providing empirical evidence on farm-nonfarm growth linkages in the context of Bangladesh. By using time series data on sectoral outputs and cointegration techniques, it finds positive and statistically significant contribution of agriculture to overall and nonfarm output growth. The estimated 'spillover/externality' effects are found to be robust and consistent under different specifications. An important policy lesson that follows is that an agriculture-focussed development strategy may not compromise with a growth maximising objective while making a powerful dent in poverty incidence.

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I. Introduction

After more than two decades of neglect by academic and donor communities, agriculture is now occupying the centre stage amidst historically high food prices inflicting widespread food insecurity and threatening several years' of progress made on poverty reduction.[1] International donors have made fresh commitments for increased resources to be devoted in agriculture of developing countries with the possibilities of different desirable outcomes, as World Bank (2008) envisions agriculture being the main engine of growth in agriculture-based economies; a major instrument to reduce rural poverty in transforming countries where agriculture is already less important; and just like any other tradable sector operating on the basis of comparative advantage and thereby contributing to growth in advanced economies. Key questions that are being asked include, *inter alia*, how to formulate country strategies so that the sector can be used to support structural transformation of the economy under 'heterogeneous' conditions (de Janvry, 2010).

This renewed emphasis on agriculture has an interesting context and important policy relevance for Bangladesh. Despite registering agricultural output growth faster than that of population, the country faces formidable food-security challenges comprising food availability, accessibility and affordability by the poor.[2] Besides, agricultural growth is considered to be vital in tackling poverty. Notwithstanding its importance, Bangladesh economy has demonstrated significant transformation in which the share of agriculture in the gross domestic product has shrunk from more than 50% to 20% over the past three decades or so. Expansion in manufacturing and services activities, including manufacturing

[1] According to Timmer (2005), both the academic and donor communities lost interest in the sector, starting in the mid-1980s, mostly because of low prices in world markets for basic agricultural commodities. Low prices, made it difficult to justify policy support for the agricultural sector or new funding for agricultural projects.

[2] See for example, Asaduzzaman, A., Ringler, C., Thurlow, J, and Alam, S (2010). "Investing in crop agriculture in Bangladesh for higher growth and productivity, and adaptation to climate change, paper prepared for Bangladesh Food Security Investment Forum, 26-27 May 2010, mimeo.

export-orientation, is now an overwhelmingly salient feature of the country's output composition. Sustained economic growth with the on-going structural transformation, as reflected in the declining relative significance of agriculture, is generally considered to be a usual route to development. Nevertheless, addressing food insecurity and poverty would imply a continuously prominent role of agriculture. In this respect, an important issue that needs to be better understood is the implications of a reinvigorated agriculture-focussed growth strategy for the overall economy. As the majority of people still rely on farm activities for their livelihoods and poverty reduction remains a major policy objective, a farm-led growth strategy appears to be an attractive one.

However, the relevant policy choices involving agriculture, growth and poverty reduction may not be straightforward: the impact of agricultural growth on poverty-reduction is likely to be strong, but the effect on the overall economy is not clear. 'Agro-pessimists' argue that the farm economy in developing countries is often the least productive sector. Consequently, when resources are limited, a policy regime favouring agriculture might not constitute a growth-maximising strategy. There are also concerns about weak linkage effects of agriculture. In an open economy farm outputs provide mainly for import-competing consumption with the comparative advantage determining sectoral resource allocation. If the productivity in agriculture is low, non-farm sectors can be argued to be the most important vehicle for growth and poverty reduction.

The role of agriculture in the growth-poverty reduction nexus is one of the most critical medium to long-term policy issues for Bangladesh. The growth-poverty trade-off associated with agriculture-focussed development strategy is greatly mitigated if farm activities can exert strong linkage effects for the rest of the economy. Despite attracting so much attention, discussions surrounding it are often uninformed in nature due to lack of in-depth empirical investigations into the nature of linkages between agriculture and the rest of the economy. While medium term plan documents and planning tools provide 'multiplier' analyses to simulate economic impact of a policy intervention and/or increased farm activities, these exercises merely produce *ex ante* predictions from pre-determined inter-industry input-output flows. There are no empirical assessments on Bangladesh of the impact of the farm sector on non-farm activities with the help of *ex post* data. As the available studies in the literature offer mixed evidence, undertaking a country specific study is the best possible option to inform the policy discussion. The present study therefore aims to contribute to the relevant macro policy discourse more effectively by carrying out an econometric investigation. Using a popular dualistic model (Feder 1982) and its appropriate adaptation – as proposed by Gemmell et al (2000)– to deal with the trended time series data while identifying genuine long-run relationships between sectoral outputs, this paper captures the externality effects of agriculture on overall gross domestic product (GDP) and other sectors.

The paper is structured as follows: after this introduction in Section I, Section II provides a brief review of the literature; Section III outlines the analytical framework, econometric methodology and data sources; Section IV provides the results of empirical investigations; finally, Section V summarises the main findings of the paper and derive important policy implications.

II. A Brief Review of the Literature

There is a longstanding academic debate and inconclusive empirical evidence over the role of agriculture in economic growth and development.[3] Agriculture's primary role in economic development has traditionally been perceived as the source of low wage labour and cheap food to the modern sector. Improved productivity in agriculture and increased farm output can contribute to overall economic growth by releasing factors of production to other sectors in the economy (Lewis, 1954; Ranis and Fei, 1961; Schultz, 1964). Against this 'passive' view, others have considered an 'active' role in which in addition to labour and food supply, agriculture provides raw materials to non-agricultural production and demands inputs and consumption goods from the modern sector (e.g., Johnston and Mellor, 1961; Kuznets, 1964; Hazell and Roell, 1983). On the consumption side, in particular, a higher productivity in agriculture is to increase rural income, thereby creating demand for industrial outputs. Such linkage effects can increase employment opportunities in the rural non-farm sector, thereby indirectly generating rural income. Moreover, agricultural goods

[3] This has been depicted in numerous academic papers, and more recently summarized in Dethier and Effenberger (2011); Barrett, et al. (2010), and Timmer (2008 and 2005).

can be exported to earn foreign exchange in order to import capital goods. The importance of such linkages was further formalized in the analytical framework of agricultural demand-led industrialization (Adelman 1984).

A related issue is then the centrality of agriculture's role in the process of economic growth, about which Tsakok and Gardner (2007) quite appropriately observe two polar views. At one extreme, a large number of studies consider agriculture as necessary for overall economic transformation, however on the other extreme there is the view that economies can always bypass the process of agricultural development and focus on building an industrial base. As the majority of people in developing countries rely on farm activities for their livelihoods, a majority of development practitioners find an agriculture-focussed growth strategy attractive one. However there are others who argue otherwise. One key argument is that a large share of agriculture does not immediately call for agricultural demand-led industrialization, which is, amongst others, affected by the openness of a country to international trade (Gollin, 2010 as reported in Dethier and Effenberger, 2011). In a closed economy gains in agricultural productivity will lead to the aforementioned linkage effect. On the other hand, if the country has a comparative advantage in agriculture, openness to trade will draw resources away from the modern sector into agriculture, which might be less productive than industry (Matsuyama, 1992).[4] In a similar fashion, it has been argued that while both farm and nonfarm sectors can contribute to growth, a less productive agriculture sector would suggest that focusing efforts in other sectors might be more beneficial (Dercon, 2009). As agriculture is usually depicted as a traditional and low productive sector, agro-pessimists point out that a policy emphasis attached to it will not help developing countries achieve their growth-maximizing objective.

Empirical findings on agriculture's role and, growth and linkage effects are mixed. Humpheries and Knowles (1998) augment the Solow-Swan growth model to include the proportion of the population working outside the agricultural sector. Empirical results obtained from non-linear estimation techniques suggest that transfer of resources from agriculture to more productive sectors of the economy is significantly correlated with

[4] This discussion here draws on a recent comprehensive review of the literature by Detheir and Effenberger (2011).

economic growth across countries. That is, the passive role of agriculture is consistent with multi-sector growth model. Gollin *et al*. (2002) develop a model to demonstrate that low productivity in agriculture can delay industrialization, which in turn results in a country's lower per capita income relative to others. By contrast, improvement in agricultural productivity can hasten the process of industrailisation. They conclude that a greater understanding of the determinants of agricultural productivity will enhance our understanding of the development process of poor countries.

Approaches that have traditionally dominated the analyses to highlight the importance and linkages effects of agriculture are input-output models offering a snapshot of linkages across production sectors; expenditure-system models assessing the importance of rural households as source of demand for farm and non-farm goods; and SAM (social accounting matrix) based models providing numerical specification of input-output demands by production sectors, distribution of capital and labour across different sectors, and distribution of household expenditures across different consumption types. While based on *a priori* numerical specification, input-output and SAM based models can simulate the effects on agricultural growth on other sectors, they are not econometric methods based on *ex-post* data. The household consumption expenditure analysis can be a powerful tool but data limitations can prove to be a severe constrain for using such methods.

Amongst earlier studies, Chenery and Syrquin (1975) used regression analysis to find out the effects of agriculture on industrial activity. Hwa (1988) – in one of the earlier systematic empirical studies to assess the spillover effects of agriculture - employed an extended production function framework to include agriculture, exports and inflation. Then using cross-section data he found that inter-country variation in industrial growth is significantly associated with inter-country variation in agricultural growth. There was also the evidence that farm sector growth contributed to overall economic growth through its favourable impact on total factor productivity. Since then a large number of studies have been undertaken to empirically study the correlation between agricultural and non-agricultural activities.

Timmer (2002) presents evidence of a positive correlation between growth in agricultural and nonagricultural GDP using panel data (1960-85) for 65 developing countries. Similarly, Self and Grabowksi (2007) established a positive relationship between different measures of agricultural productivity and average growth of real GDP per capita over 1960 – 1995 for a cross-section of countries. However, using more recent data (1980-2001) for 52 developing countries, Gardner (2005) finds that agriculture does not seem to be a primary force behind growth in national GDP per capita. Since correlation does not imply causation, a few studies have studies have attempted causality assessment using statistical methods. Amongst others, Tiffin and Irz (2006) report agricultural value added exerting positive causal effects on GDP per capita in developing countries, while Bravo-Ortega and Lederman (2005) show regional differences in such relationship within the set of developing countries and reverse causality for developed countries. Shifa (2011) used instrumental variable method to identify the causal impact of agricultural growth on manufacturing growth under a panel regression framework. The results show a significant positive effect with the impact increasing with higher share of agriculture in national economy. In an economy with 50% agricultural GDP, a 1 per cent increase in agricultural output increases manufacturing output by about 1%.

Several studies have focused on individual country experiences using country-specific time series data. Using cointegration techniques and time-series data on Malaysia, Gemmell *et al*. (2000) find that it is the manufacturing output that have causal effects on agricultural value-added. Although the short-run expansion of manufacturing output adversely affects agriculture, the positive effect is borne out over the long-run. The growth in services is found to be associated with declining agricultural output. Matahir (2012) provides further evidence for Malaysia that the causal growth effects run from manufacturing to agriculture. For Thailand, however, Jatuporn *et al*. (2011) reports a bi-directional causality between agricultural and overall economic growth. Similar findings are also reported by Katircioglu (2006) for North Cyprus.

Chebbi and Lachaal (2007) use the Vector Autoregression (VAR) and cointegration analysis utilizing Tunisian data from 1961 to 2005 to suggest that important economic sectors tend to move together in the long run although in the short run agriculture had a limited role and its growth might not be conducive to non-agricultural sectors. Kumar and Kumar (2011) report positive contribution of agriculture to long-run economic growth in Fiji, but compared to manufacturing and services the effects are smaller. In the short run however the impact of farm growth is found to be mixed.

De Janvry and Sadoulet (2009), show that for China over 1980 – 2001, a 1 per cent agricultural growth had an effect on aggregate growth of 0.45 per cent, whereas the indirect effect through the non-agricultural sector represents half this effect. With declining relative significance of the farm economy, should one expect only to see its insignificant role in promoting growth in other sectors? Yao (2000) provides interesting evidence on China in this respect. Making use of time series data on sectoral outputs, he finds that despite diminished share in aggregate output, the farm sector still exerts significant spillover effects on non-farm activities.

Unlike the results of farm-nonfarm growth linkages, the poverty effects of agricultural growth appear to be more conclusive. Amongst others, in the context of India, Datt and Ravallion (1996) find that higher farm productivity reduces both absolute as well as relative poverty. Similarly, Loayza and Raddatz (2010) show for a cross-section of developing countries that growth in more labor-intensive sectors such as agriculture has a larger impact on poverty reduction than less labor-intensive activities. Also, Christiansen and Demery (2007) estimate that 1 percent per capita agricultural growth reduces poverty 1.6 times more than the same growth in industry and three times more than growth in the service sector.

It then follows that the relevant policy choices involving agriculture, growth and poverty reduction may not be straightforward: the impact of agricultural growth on poverty-reduction is likely to be strong, but the effect on the overall economy is not clear. When resources are limited and farming is not a productive sector, a reinvigorated focus on agriculture might not constitute a growth-maximising strategy. The growth-poverty trade-off associated with agriculture-focussed development strategy is greatly mitigated if growth in agriculture has strong linkage effects for the rest of the economy. Given that the findings from the empirical studies differ, and consequently country-specific case studies are the best possible option to guide policy directions.

III. Analytical framework, Econometric Methodology and Data

The analytical framework being used in the paper is grounded in dualistic models that have studied sectoral linkages within an economy. Earlier studies duly recognized the relationship between agriculture and industry as one of interdependence and complementarity. For example – as pointed out in Hwa (1988) – while providing inputs to industry, agriculture receives from it modern inputs, advanced technologies, and consumption goods to increase its productivity. Amongst others, Chenery and Syrquin (1975) and Hwa (1988) also used an extended production function framework by incorporating such factors as growth of inputs, agriculture, and price level (inflation) along with capital stock and labour force to investigate the intersectoral linkages.

Subsequently, the dualistic framework that has perhaps been most extensively used relates to assessing the impact of the export sector. One significant contribution in this regard is due to Feder (1982) who depicted the effects of exports on output explicitly as the sum of 'externality' and 'productivity differential' effects. This paper adapts the same Feder framework, but to deduce an empirical model to assess the linkage effects of agriculture. At the outset, it is generally accepted that agriculture has certain externality effects for the overall economy, as also pointed out in the discussion above. While it may be difficult to accept the farm sector's having (positive) productivity differential effects (compared to the rest of the economy), which is an important feature associated with the export sector in Feder's theoretical construct, it is argued that the sum of 'externality and production differential effects' arising from agriculture will be left for empirical verification, which is the main contribution of this paper.

Following Feder (1982), but adapting it to our case, the economy can be divided into two sectors – agriculture (A) and non-agriculture (NA). Then the specifications for aggregate output, farm production and non-farm activities can be given by:

$$Y = N + A; \quad N = f\ (K_N, L_{N,} A); \text{ and } \quad A = g\ (K_A, L_A)$$

Where, K_N and L_N are capital and labour in nonfarm sector, and K_A and L_A are the corresponding factors in the production of agriculture. As Feder originally assumed, exports to enter into the production fucntion of non-export sector and not vice versa, we do the same giving that prominent role to agriculture. While this can be considered as a limitation but a subsequent extension following Gemmell *et al* (2000) would address this issue. Total differentiation of the equations for N and A yields:

$$\dot{N} = f_K \dot{K}_N + f_L \dot{L}_N + f_A \dot{X}$$
$$\dot{A} = g_K \dot{K}_A + g_L \dot{L}_A$$

where f_k, f_L and g_K, g_L are marginal productivities with respect to the respective factors of production in the non-farm and farm sector, respectively; f_A represents the marginal externality effects of A on N. Since by definition$\dot{Y} = \dot{N} + \dot{A}$, assuming that productivity of inputs differ between sectors by a factor, γ so that$g_i = (1+\gamma)f_i$, one can obtain[5]:

$$\dot{Y} = f_K \dot{K}_N + f_L \dot{L}_N + f_A \dot{A} + (1+\gamma) f_K \dot{K}_A + (1+\gamma) f_L \dot{L}_A$$
$$\dot{Y} = f_K(\dot{K}_N + \dot{K}_A) + f_L(\dot{L}_N + \dot{L}_A) + f_A \dot{A} + \gamma (f_K \dot{K}_A + f_L \dot{L}_A)$$

Defining $\dot{K} = \dot{K}_N + \dot{K}_A \equiv \dot{K}$ and $\dot{L} = \dot{L}_N + \dot{L}_A \equiv \dot{L}$ and using the productivity differential effects, the adapted Federian equation becomes:

$$\dot{Y} = f_K \dot{K} + f_L \dot{L} + \left(\frac{\gamma}{1+\gamma} + f_A\right) \dot{A}$$

With some further manipulation, the empirical equation is derived as:

$$\frac{\dot{Y}}{Y} = \alpha\left(\frac{\dot{K}}{K}\right) + \beta\left(\frac{\dot{L}}{L}\right) + \left[\left(\frac{\gamma}{1+\gamma}\right) + f_A\right]\left(\frac{\dot{A}}{A}\right)\left(\frac{A}{Y}\right) \qquad (1)$$

It is then clear that if marginal productivities are equalized across sectors ($\gamma = 0$) and if there are no spillover effects from agriculture ($f_A = 0$), equation (1) becomes the traditional neo-classical formulation of the growth model. Feder noted that the parameter α is to be interpreted as the marginal productivity of capital in the non-agricultural sector, rather than as marginal productivity of capital in the economy as a whole.

The Feder equation has been the workhorse for empirical vindication of the export-led hypothesis, suggesting that exports offer significant positive externalities for non-export activities. One particular problem associated with equation (1) that has been highlighted

[5] Feder made the assumption $\gamma > 0$, to show productivities in the export sector higher. We can, however, keep this open for empirical verification.

with the advent of the modern time series analysis based on unit roots and cointegration is that by using the growth rates of variables it wipes off long-run information involving the variables. It is true that macro time series data can be non-stationary in nature and the use of such data could result in spurious regression results. While transformation into rates of growth can result in stationary data, but mere use of them would not result in long-run relationships. Gemmell et al (2000) provides a revised formulation of the Feder framework to capture the long-run relationship. While doing so, they also overcame a limitation by allowing the possibility of two-way spillovers.[6] When agriculture (A) and non-agriculture (N) are two sectors, Gemmell et al considers:

$$A = \theta_a + \alpha_a L_a + \beta_a K_a + \gamma_a M$$
$$N = \theta_n + \alpha_n L_n + \beta_n K_n + \gamma_n A$$

Clearly, now the externality effects, γ_i, allows for the possibility of bi-directional spillovers. Following Feder, they assume: $\frac{\alpha_n}{\alpha_a} = \frac{\beta_n}{\beta_a} = 1 + \delta$. If, $Y = A + N$, $L = L_a + L_n$, $K = K_a + K_n$ and $\theta = \theta_a + \theta_n$, then it can be shown that:

$$A = \frac{\theta(1+\delta)}{1+\delta-\gamma_n} + \left(\frac{1+\delta}{1+\delta-\gamma_n}\right)(\alpha_a L + \beta_a K) + \left[\frac{(1+\delta)\gamma_a - 1}{1+\delta-\gamma_n}\right] N \qquad (2)$$

Equation (2) is the equivalent in levels of the expression derived in the form of growth rates, in which $\left[\frac{(1+\delta)\gamma_a - 1}{1+\delta-\gamma_n}\right]$ captures the effects on agriculture of expansion in non-farm output for given factor endowments, and can be positive and negative. One important contribution of Gemmel et al. is that they showed that the model can be adapted to eliminate the aggregate inputs (capital and labour). If sectoral marginal products of labor are related to average productivity in the economy, such that, $\alpha_a = \alpha\left(\frac{Y}{L}\right)$ and making a similar assumption for capital productivity: $\beta_a = \beta\left(\frac{Y}{K}\right)$, (2) can be reduced to

$$A = \frac{\theta}{(1-(\alpha+\beta)-\frac{\gamma_n}{1+\delta})} + \left[\frac{(1+\delta)(\gamma_a + \alpha + \beta) - 1}{(1+\delta)\left(1-(\alpha+\beta)\right) - \gamma_n}\right] N \qquad (3)$$

The elimination of factors of production from (3) is particularly useful. Many developing countries do not have actual data on employment by sectors and as such most of the studies that have used the Feder framework have used either population or labour force as

[6] As pointed out above, Feder allowed only externality effects going from an export sector. Similarly in our case, we have considered on way flow of the same from agriculture.

a proxy for labour. Similarly, the data on capital stock is also problematic. There is usually data on investment, but not capital stock. Most studies usually construct a capital stock series using some statistical method, which may contain measurement errors. Equation (3) therefore overcomes serious data shortcomings. Gemmell et al. also extends the model to a three-sector economy.

Clearly equation (3) cannot be regarded as a traditional reduced form relationship since both A and N are potentially endogenous. This problem can be tackled by employing the recently popular econometric method of the Vector Autoregression (VAR) analysis. This approach is essentially a non-structural framework in which no particular relationships are imposed on the variables based on economic theory. As long as there is some hint of certain variables being related, a VAR model can be employed to investigate the empirical relationship. Time series econometrics (e.g. unit roots and cointegration methods) and wider recognition that many macro variables are jointly determined (i.e. endogenous) have made widespread use of VAR models. Given their very frequent applications, we only provide an outline of the VAR methodology that allows testing for long-run relationship.

A VAR framework with lag order p, VAR(p) can be represented as:

$$x_t = \mu D_t + \pi_1 x_{t-1} + \pi_2 x_{t-2} + \ldots + \pi_p x_{t-p} + \varepsilon_t$$

where D_t is (nx1) deterministic trend; x_t is an (n x 1) of endogenous variables; and ε_t is (n x 1) vector of independently distributed disturbances of zero mean and constant variance-covariance matrices. Following Johansen and Juselius (1990), if the level variables cointegrate, an equilibrium error-correction model can be represented by:

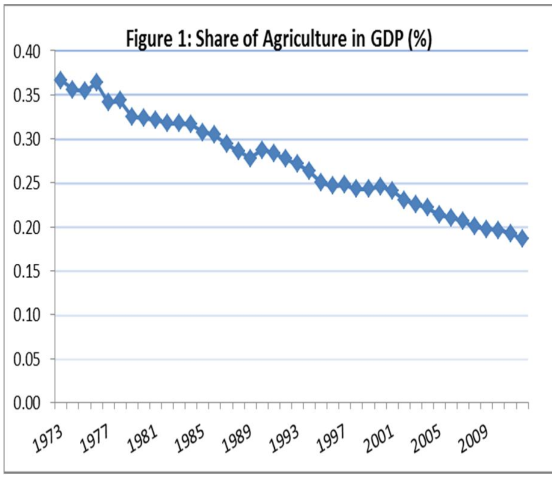
$$\Delta x_t = \mu D_t + \alpha\beta' x_{t-p} + \Sigma\theta_i \Delta x_{t-i} + \varepsilon_t$$

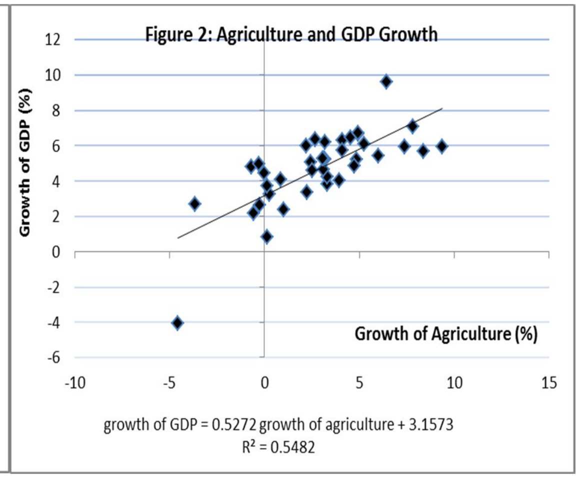
where comprises β (n x r) matrix of cointegrating vectors or long-run relationship, and α is a matrix of equilibrium correction coefficients of the same dimension, capturing the rate at which the disequilibrium is corrected in the system. Johansen and Juselius show that empirically the number of cointegrating vectors can be determined by using the Trace and Maximal Eigenvalue statistics. In the context of the present paper, as Gemmell et al. mention, we may consider equilibrium relationships as representing intersectoral linkages that bind sectors together in the process of economic development. When resource competition, productivity differentials or spillovers produce long-lasting effects, those long-run relationship will be evident through the coefficients of β.

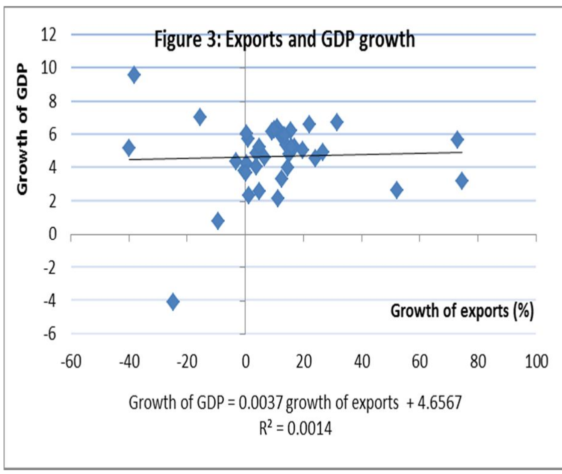
Data

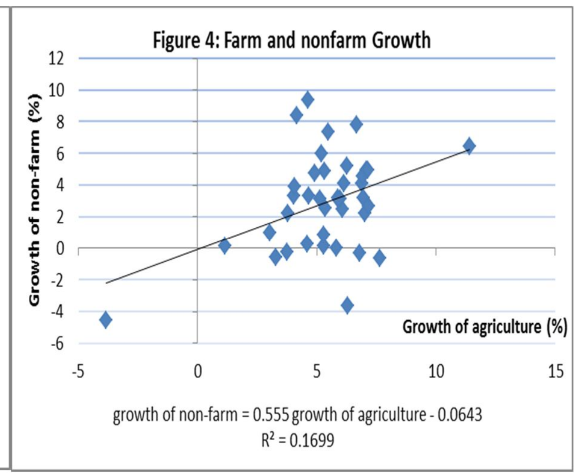
Empirical exercises will closely follow the analytical structures discussed above. The basic dual sector model as depicted in Feder (1982) as well as its adaptation will be analysed to assess the effects of farm output on nonfarm sector. The source of data that are used in the empirical exercise is mainly the Bangladesh Bureau of Statistics (BBS). It provides a GDP estimates and its breakdown by various sectors such as manufacturing, agriculture and services. BBS currently uses 1995-96 as the base year and all data for the period 1980-2012 are available at the same base year prices. BBS also has the relevant data for previous years since 1973 but with different prices. World Bank (2012) compiles these data to produce a longer time series and for this paper we use this data for 1973-2012. Two things need to be mentioned here. There is no information on actual employment. Hence we use the labour force data, defined as population aged between 15 and 64 years. The series on capital stock is constructed using the perpetual inventory method based on the information on investment-GDP ratio and allowing for some depreciation.

Before going to the results section, it might be useful to have a cursory glance at agriculture's significance in Bangladesh's economy through a few graphical expositions. First, Figure 1 shows that over the past 40 years, the relative significance of agriculture in GDP has almost halved: from about 37 per cent in 1973 to just 19 per cent in 2012. Nevertheless, the growth in agriculture fits very closely with the growth of overall output (Figure 2). Indeed, when a simple bivariate relationship is considered, a one-percentage point increase in farm output growth is associated with 0.53 percentage-point increase in GDP. Approximately 55 per cent of variation in overall economic growth can be explained by the variation in agricultural growth alone.









Agriculture's prominent role is highlighted further if one compares the bi-variate relationship between the growth of exports and overall GDP (Figure 3). Notwithstanding the facts that agricultural output are measured here as value-added while exports are not, and that the export sector comprised a relatively small activity until the mid-1990s, it may appear to be quite striking to observe an almost random relationship between export and GDP growth rates. Reducing the sample to from the mid-1990s to 2012 does not make much difference. These two simple graphs seem to suggest agriculture's much more prominent role in overall economic activity. Finally, Figure 4 provides the scatterplot of farm and nonfarm output growth rates. Clearly there are outliers, nevertheless about 17 per cent variation in nonfarm growth can be explained by variation in the farm output and a one percentage point increase in agricultural growth is associated with 0.55 percentage points increase in nonagricultural growth. Of course drawing causality inferences from bi-variate analysis is problematic but as the scatterplots in Figure 2 and Figure 4 are on growth rates they seem to suggest strong inter-sectoral linkages.

IV. Results

Let us first present the results of the standard Feder model. The basic estimating equation involved explaining the growth of GDP by rates of growth of capital, which is represented by the investment-GDP ratio, growth of labour, and growth of agriculture weighted by its share in GDP. Since these variables are growth rates, it is expected that they are unlikely to be non-stationary (i.e. to have unit roots) over a sufficiently long period of time.[7] Therefore, the ordinary least squares (OLS) regression results would be free from the shortcomings associated with non-stationary data.

Table 1 reports the OLS estimates of the basic Feder equation. An initial estimation revealed the problem of non-normality of errors, which would not allow for drawing valid statistical inferences. This problem could be tackled by incorporating an intercept dummy for 1974. Its inclusion however does not affect individual coefficient estimates.

Table 1: Estimation of the Basic Feder Equation

$$\underset{(s.e.)}{\frac{\Delta Y}{Y}} = \underset{(0.0224)}{-0.01809} + \underset{(0.0535)}{0.2254^{***}}\frac{\Delta K}{K} + \underset{(0.5185)}{0.4033}\frac{\Delta L}{L} + \underset{(0.2662)}{1.2534^{***}}\left(\frac{\Delta A}{A}\right)\left(\frac{A}{Y}\right) + \underset{(0.0152)}{0.633^{***}}\,D74$$

Adj. R^2 = 0.74 Serial Correlation: $\lambda^2(1)$ = 3.09
Functional form: $\lambda^2(1)$= 3.36 Normality : $\lambda^2(2)$= 3.37

$$\underset{(s.e.)}{\frac{\Delta Y}{Y}} = \underset{(0.013)}{-0.0053} + \underset{(0.0533)}{0.2153^{***}}\frac{\Delta K}{K} + \underset{(0.2993)}{1.2837^{***}}\left(\frac{\Delta A}{A}\right)\left(\frac{A}{Y}\right) + \underset{(0.006)}{0.5252^{***}}\,D74$$

Adj. R^2 = 0.74 Serial Correlation: $\lambda^2(1)$= 1.89
Functional form: $\lambda^2(1)$= 2.48 Normality : $\lambda^2(2)$ = 0.53

Note: All coefficients are estimated from heteroscedasctity adjusted variance-covariance matrix (White, 1980). *** indicates statistical significance at the one and five per cent 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 and functional form follow a chi-square distribution with one degree of freedom while the statistic for normality follows a chi-square distribution with 2 degrees of freedom.

[7] We tested for unit roots in the variables on their logarithmic transformation and on their first differences. The results suggested the level variables to be non-stationary, but their first differences are stationary. See Appendix A1 for details.

As reported in the upper panel of Table 1, the estimated coefficients on growth rates of capital, labour, and weighted agricultural growth are correctly signed. Although $\Delta L/L$ failed to register statistical significance, the other two parameters turned out to be statistically significant at the one per cent level. The effect of capital growth rate on the overall output growth appears to be modest, 0.225. However, the parameter of our interest – the externality effect of agriculture – is quite large: a one percentage point increase in the weighted farm growth is associated with a 1.25 percentage point increase in GDP.

The lack of significant effect of labour growth could be attributable to an overall surplus labour situation in Bangladesh. When the equation is re-estimated after dropping $\Delta L/L$, the results does not change much (see lower panel in Table 1). The coefficient on weighted agricultural growth increases slightly to 1.28 while the same on capital growth rate fell marginally to 0.215, with both the parameters maintaining their previous level of statistical significance.

As mentioned above, the Feder framework was originally developed and has mainly been used to assess the combined externality and productivity differential effects of exports. The same may be of particular interest here as well. The spillover effects of exports generally attract huge attention with numerous studies attempting to measure their significance for a large number of developing countries. Column 2 in Table 2 presents the estimation of the typical Feder equation in which the weighted agricultural growth variable has been replaced by the corresponding export variable.[8] The regression results are however subject to a massive non-normality of residuals problem, as indicated by a very large value of Jarque-Bera test statistics. This could only be overcome by incorporating three intercept dummies – one each for 1974, 1976 and 1978. Column 3 thus presents the results with normally distributed regression residuals, providing the basis for drawing usual inference. It is found that growth rates of both factors of production are correctly signed and statistically significant. The estimated coefficient on weighted export variable turns out to be 0.23 and

[8] The data on exports come from the same source as mentioned above. It is measured in constant 1995-96 prices.

becomes significant only at the 10 per cent level. This lower estimate of export externality effect is in sharp contrast to what is reported in Begum and Shamsuddin (1998).[9]

Table 2: Estimation of the externality effects: Exports and agriculture

Dependent Variable is ΔY/Y			
Constant	-0.059*** (0.0216)	0.0529*** (0.0151)	- 0.2025 (0.0178)
$\frac{\Delta K}{K}$	0.3539*** (0.0599)	0.2484* (0.1475)	0.1932*** (0.0459
$\frac{\Delta L}{L}$	1.1544** (0.5051)	-0.9179* (0.5568)	0.5701 (0.4576)
$\left(\frac{\Delta X}{X}\right)\left(\frac{X}{Y}\right)$	0.2320* (0.1339)	0.2076 (0.2193)	0.3079** (0.1187)
$\left(\frac{\Delta A}{A}\right)\left(\frac{A}{Y}\right)$	-	-	1.2624*** (0.2353)
$D74$	0.1239*** (0.016)	-	0.0723*** (0.0142)
$D76$	0.049*** (0.086)	-	-
$D78$	0.0524*** (0.0058	-	-
Adjusted R^2	0.7409	0.085	0.7801
Serial correlation: $\lambda^2(1)$	1.1052	6.0439	2.7955
Normality: $\lambda^2(2)$	1.3752	111.3242	1.8359
Functional form: $\lambda^2(1)$	2.4032	2.5340	3.5782

Note: ***, **, and * are for statistical significance at the 1, 5 and 10 per cent levels, respectively.

In Column 4 of Table 2, we include both the weighted export and agriculture growth rates to capture externality effects from both the sectors. The inclusion of agriculture improved the explanatory power of the model and made the problem non-normality of residuals less severe, as just one intercept dummy for 1974 ensured the usual classical properties involving the distribution of the error term. It is now observed that the impact of capital growth rate is considerably reduced to 0.19 although the effect remains statistically significant at the one per cent level. No significant influence of labour growth is found. The inclusion of agriculture helps improve the magnitude of externality effects arising from exports. The relevant coefficient, 0.31, is now statistically significant at the five per cent level. Finally, the magnitude and significance of agricultural spillover effects remain almost identical to those reported in Table 1. Therefore, agriculture's externality effects appear to be four times higher than exports'.

[9] They reported a coefficient representing the export externality effects of 0.96, rising up to 1.52 in different model formulations. The data used were for the period 1962 to 1992.

It needs to be pointed out that the much higher effect of farm output may not be considered as a surprise. Agricultural output considered here is a measure of value added, while exports are not. As the export sector is more heavily reliant on imported raw materials and inputs, its overall impact on the aggregate output is likely to be lower than agriculture. Furthermore, Bangladesh's overwhelming reliance on ready-made garments (apparels), a significant part of which has a very small domestic value-added content, would imply even relatively weaker contribution to economic growth. Therefore, although the export sector is portrayed as a source of spillover effects in terms of improved management, skill development, higher productivity gains, etc, - notwithstanding their relevance – the sector's overall contribution to growth remains much smaller than that of the farm sector.

Since the Feder framework has been criticized by pointing out that in a regression involving the growth rates of variables, the long-run information is lost while economic theories are about long-run relationships, which are most often captured through studying the behaviour of variables on their levels. This is despite the fact that many macroeconomic times series variables contain unit roots in which case the use of OLS regression method could result in spurious results. Instead, cointegration techniques should be used to tackle the problem on non-stationarity of data while finding out a valid long-run relationship.

One way of making the Feder's approach consistent with the modern time series methodology is to ascertain a long-run relationship (cointegration) amongst the variables first and then to estimate the equations as in Tables 1 and 2 above while incorporating the long-run information into it. To achieve this, we formally test the variable for unit roots, using the Augmented Dickey-Fuller test. Following the standard practice, these tests are carried out on the logarithmic transformation of the level variables and on their first differences as well. Appendix A1 provide the results of the unit root tests using the Augemented Dickey-Fuller (ADF) procedure. The results suggest that all our variables can be considered as integrated of order one on their levels. That is, their first difference will result in stationary time series.

Now corresponding to Table 1, we test cointegration amongst the level variables (on their logarithmic transformation denoted by y, k, l, and a) using the Johansen-Juselius (JJ) methodology. The JJ procedure starts with selecting a suitable order of VAR. As the results can be sensitive to the chosen lag lengths and therefore a great deal of caution is to be exercised to determine the optimal lag lengths. One can use statistical methods, such as Akaike Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC) for selecting the lag length. However, in small samples, it is difficult to initially set for sufficiently large VAR orders based on which testing down can be carried out. In our case with 39 years' data and 4 variables, overparameterisation of the model by choosing large lag orders is likely to be problematic. With the initial maximum lag length set at 3, AIC and SBC test results differed. However, irrespective of the choice of lag lengths (between 1 and 3), there is evidence of cointegration, as indicated by the *maximal eigenvalue* and *trace* statistics (Appendix A2). One salient feature of Johansen-Juselius methodology is that it can provide more than one cointegrating relationship in a model involving more than two variables in which case one has to identify the appropriate equations based on theoretical expectations about the size, sign and statistical significance of different variables. The coefficients under different CVs varied widely, particularly those ones associated with labour and capital, and we did not have much information *a priori* to identify the most appropriate equations.

Given the evidence of cointegration, however, one convenient way of estimating the long-run relationship normalized on *y* is to employ the Phillips-Hansen (PH) Fully Modified OLS methodology. This 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, correct for endogeneity in the regressors, and provides standard errors that follow standard normal distribution asymptotically and thus are valid for drawing inferences. The PH technique provided the long-run labour and capital elasticities of 0.25 and 0.16 respectively, along with a coefficient of 0.98 for agriculture. All these parameter estimates are statistically significant. Table 3 therefore represents the corresponding short-run dynamic adjustment of the model using the error-correction mechanism, which makes the Feder's basic equation consistent with the time-series cointegration technique. In this equation the short-run dynamics are being captured by first differenced variables, having contained the long-run information as the error-correction term (as the last right had side variables). We first built a 'general' model by including up to one lag of explanatory and dependent variables and then dropped the most insignificant ones to come up with the results presented in Table 3. The estimates initially suffered from residual non-normality, which was removed by inserting one intercept dummy for 1975.

Table 3: Estimation of the error-correction model with the corresponding Feder long-run equation

$$\Delta y_t = 0.047^{***} - 0.768\Delta l_t + 0.215^{***}\Delta y_{t-1} + 0.397^{***}a_t - 0.082^{***}D75 - 0.147^{**}(y - 0.816 - 0.246l - 0.156k - 0.986a)_{t-1}$$
$$(s.e.)\quad (0.011)\quad (0.327)\quad (0.069)\quad (0.056)\quad (0.011)\quad (0.059)$$

Adjusted R^2 = 0.85 Serial Correlation: $\lambda^2(1) = 0.114$ Functional form: $\lambda^2(1) = 2.39$
Normality : $\lambda^2(2) = 1.792$ Heteroscedasticity : $\lambda^2(1) = 2.095$

$$\Delta y_t = 0.0507^{***} - 1.0366\Delta l_t + 0.3289^{***}\Delta y_{t-1} + 1.3798^{***}\left(\frac{\Delta A}{A}\right)\left(\frac{A}{Y}\right) - 0.0884^{***}D75 - 0.114^{**}(y - 0.816 - 0.246l - 0.156k - 0.986a)_{t-1}$$
$$(s.e.)(0.109)\quad (0.310)\quad (0.0761)\quad (0.193)\quad (0.101)\quad (0.063)$$

Adjusted R^2= 0.86 Serial Correlation: $\lambda^2(1)$= 0.715 Functional form: $\lambda^2(1)$= 0.5653
Normality : $\lambda^2(2)$= 4.34 Heteroscedasticity : $\lambda^2(1)$ = 3.335

Note: ***, **, and * are for statistical significance at the 1, 5 and 10 per cent levels, respectively.

The upper panel in Table 3 shows that in the short-run the labour elasticity is significant although its sign is contradictory to our expectation although economic theories do not tell anything about the short-run behaviour of the variables. No significant effect of capital was detected and thus it was dropped from the equation. The short-run agricultural externality is estimated to be 0.40 and significant at the one percent level. The coefficient on the error-correction term is correctly signed and statistically significant. It once again validates the existence of a long-run relationship as postulated in our model. The coefficient suggests that about 15 per cent of disequilibrium errors are corrected within the first year after a deviation from the long-run path.

It needs to be pointed out that Δa_t is not exactly what Feder derived as the measure of spillover effects. It is the corresponding short-run variable which normally follows from the long-run relationship. However, it is possible to consider other variables in the short-run, given the long-run relationship as defined above. The lower panel in Table 3 thus replaces Δa_t with the weighted agricultural growth. Again, we started with a generalized model including up to one lag of explanatory and dependent variable and then dropped the insignificant ones. The results do not change much qualitatively except for the weigthed agricultural growth to capture a much bigger coefficient of 1.38 which is significant at the one per cent level. The error-correction term remains significant as before and the model diagnostics do not reveal any problems. Therefore, even after making the Feder framework consistent with the time-series cointegration technique, there is evidence of positive externality effects of agriculture.

Table 4 presents a set of error-correction models considering the long-run relationship only between y_t and a_t. This is in line with Gemmell *et al* (2000) modification of the Feder framework, as discussed above, which did not require data on capital and labour. For Bangladesh we used labour force data as a measure of labour. Again, in the absence of the information on capital stock, such data is generated through statistical means. Therefore, Gemmel et al representation would be a useful approach to get rid of the variables that may be subject to measurement errors.

We used the Maximal Eigenvalue and Trace tests to determine if y_t and a_t cointegrate. The test results for selecting the VAR orders proved to be inconclusive but an examination of residuals seemed to suggest that an order of 2 would suffice. There is however robust evidence of cointegration irrespective of the VAR orders chosen, and both the Eigenvalue and Trace tests supported this conclusion. Columns 2 and 3 in Table 4 are the error-correction models corresponding to a long-run relationship where the agricultural elasticity to output is estimated to be 1.60.[10] These models included both labour and capital in the short-run equations, but in none of the models was capital found to be significant and hence it was dropped.[11] We also found that the results did not change qualitatively if the significant labour variable was to be dropped. Column 2 shows that in the short run the agricultural elasticity to output is 0.42, which is statistically significant at the one per cent level. Column 3 replaces Δa_t with the weighted agricultural growth rate, and both it and its first lag turned out to be statistically significant. The net spillover effect is estimated to be 0.90. The error-correction terms in both the cases remain correctly signed and statistically significant.

[10] Along with the evidence of cointegration, the coefficient on a_t was also found to be statistically significant at the one per cent level.

[11] As we did before, along with Δa_t, both Δl_t and Δk_t and their first lags were included to begin with and then the models were tested down dropping the insignificant variables.

Table 4: Dynamic short-run equations involving Gemmell et al representation of Feder's dual sector model

Constant	0.042*** (0.008)	0.041*** (0.006)	0.057*** (0.008)	0.056*** (0.007)
Δy_{t-1}	0.232*** (0.054)	0.4612*** (0.061)	0.166*** (0.052)	0.378*** (0.0980)
Δl	-0.641*** (0.255)	-0.767*** (0.198)	-1.076*** (0.234)	-1.22*** (0.20)
Δa_t	0.42*** (0.049)	–	0.396*** (0.043)	–
$\left(\frac{A}{Y}\right).\Delta a_t$	–	1.3402*** (0.143)	–	1.275*** (0.135)
$\left(\frac{A}{Y}\right).\Delta a_{t-1}$	–	-0.437*** (0.22)	–	-0.405*** (0.159)
$D75$	-0.0810*** (0.008)	-0.0877*** (0.0029)	-0.081*** (0.008)	-0.086*** (0.066)
$D80$	-0.023*** (0.006)	-0.0300*** (0.003)	-0.019*** (0.006)	-0.024*** (0.0055)
$D95$	0.0191*** (0.006)	0.0157*** (0.002)	0.0175*** (0.005)	0.0142*** (0.005)
$y_t - (2.3109 - 1.605a_t)$	-0.09*** (0.03)	-0.093*** (0.031)	–	–
$y_t - (0.5373 + 1.08a_t + 0{,}015T)$	–	–	-0.093* (0.053)	-0.111** (0.049)
Adjusted R^2	0.91	0.93	0.93	1.91
Serial correlation: $\lambda^2(1)$	0.78	0.24	0.004	1.55
Normality :$\lambda^2(2)$	0.69	0.39	3.422	_
Heteroscedasticity:$\lambda^2(1)$	0.16	Adjusted (white procedure)	0.015	Adjusted (white procedure)

Note: ***, **, and * are for statistical significance at the 1, 5 and 10 per cent levels, respectively.

Given that a relatively large coefficient on agriculture is observed when the long-run relationship is defined only between y_t and a_t, a trend term is included in the cointegrating vector. The idea is that other factors could also have contributed to overall output and the trend factor will capture the combined effects of these variables.[12] The inclusion of the trend term reduces the long run coefficient on a_t to 0.8 (columns 4 and 5), which however maintains its statistical significance at the one per cent level. The trend term also achieves the same level of statistical significance. The short-run coefficients on Δa_t (in column 4) and Δa_t*(A/Y) (in cloumn 5) are similar to their counterparts in columns 3 and 4 respectively. The error-correction terms are correctly signed and statistically significant, as one would normally expect for a valid representation of the short-run dynamic equation.

[12] Note that there is no consensus on the inclusion of the trend term. However, in many instances it has been used, e.g. Yao (2000).

Finally, it would be of interest to see if farm and nonfarm outputs cointegrate. By deducting agricultural output from gross domestic value added, we generate a time series variable of 'non-agricultural output' and use its logarithmic transformation (na_t) in the cointegration test. Both the Akike Information and Schwarz Bayesian criteria suggested a VAR order of 2 to be used in the model specification (Appendix A5). There appears to be a robust evidence of cointegration as both the Maximal Eigenvalue and Trace statistics rejected the null-hypothesis of no-cointegrating relationship in favour of one cointegrating vector (Appendix A6). A trend term is inserted into the cointegration space as our careful examination of residuals indicated that it would be necessary to make the cointegrating relationship stationary.[13]

The Johansen technique - which treats na_t and a_t, jointly endogenous - estimates a coefficient on the latter as 1.27, when the long-run relationship is normalized on na_t. Therefore, a one per cent increase in agricultural output is associated with 1.27 percent increase in non-agricultrual output. The effect is statistically significant at the one percent level. This high elasticity is quite interesting given that unlike the previous cases when agricultural output was part of GDP, non-agricultural output is constructed by removing the farm value added from the overall output. Yet a strong externality effect of farm output on the non-farm sector is obtained.

The upper panel of Table 5 provides the relevant short-run dynamics considering the long-run relationship only between na_t and a_t. The short-run agriculture elasticity is estimated to be 0.17 which is significant at the five per cent level. The error-correction term is correctly signed and statistically significant at the one per cent level. It shows that it takes just about four years to correct for all disequilibrium errors. The lower panel in Table 5 replaces Δa_t with weighted agricultural growth variable. At this, like the previous cases, the farm spillover effect becomes more prominent: the estimated coefficient is 0.71 and highly significant. The error-correction term maintains its usual properties. On the whole, these results suggest agriculture to have important externality effects on the nonagricultural sector.

[13] Appendix A7 provide cointegration test results with the trend term included.

Table 5: Short-run model: farm-nonfarm sectoral linkages

$$\Delta na_t = 0.0479^{***} - 0.1728^{**}\Delta a_t + 0.0493^{***}D74 - 0.0759^{***}D75 - 0.2516^{***}\ (na - 1.273a + 0.016\ T)_{t-1}$$
$$(s.e.)\ \ (0.0029)\ \ \ (0.0747)\ \ \ (0.0117)\ \ \ (0.0129)\ \ \ (0.0621)$$

Adj. R^2*= 0.73* *Serial Correlation:* $\lambda^2(1)$*= 0.15* *Functional form:* $\lambda^2(1)$*= 0.5653*
Normality : $\lambda^2(2)$*= 0.16* *Heteroscedasticity :* $\lambda^2(1)$ *= 1.95*

$$\Delta na_t = 0.0487^{***} + 0.717\ ^{**}\Delta a\ \frac{A}{Y} - 0.0726^{***}D75 - 0.2745^{***}\ (na - 1.273a - 0.016\ T)_{t-1}$$
$$(s.e.)\ \ (0.0032)\ \ \ (0.2987)\ \ \ (0.0161)\ \ \ (0.078)$$

Adj. R^2= 0.66 Serial Correlation: $\lambda^2(1)$= 0.1.15 Functional form: $\lambda^2(1)$= 1.285
Normality : $\lambda^2(2)$= 5.23 Heteroscedasticity : $\lambda^2(1)$ = 0.87

Sectoral Interactions Involving Agriculture, Manufacturing and Services

Given the finding that farm and non-farm value added move together, the analysis is extended here further to disaggregate the latter into two principal components, viz., manufacturing and services, and then examine how agriculture interacts with them. This will be of particular interest to policymakers as non-farm activities are generally referred to as one of these two broad components. Following Gemmell *et al.* (2000), the sectoral interactions can be studied without the need for information on factors of production variables.[14] Along with a_t to consider logarithmic transformation of agricultural value-added, the corresponding transformations of services and manufacturing value-added in real terms (in 1995-96 prices) are denoted as m_t and s_t, respectively. Like a_t, m_t and s_t also appear to be I(1) variables (Appendix A1).

To test for cointegration a VAR (2) model is chosen. The selected VAR order is supported by various lag selection criteria including AIC and SBC (Appendix A8). Examinations of residuals associated with the 3 unrestricted VAR equations also seemed to have suggested an appropriate choice of lag lengths. As usual, the Trace and Maximal Eigenvalue test statistics were employed to test of cointegration and both the tests indicated the existence of one

[14] This is particularly advantageous in a VAR model that uses lag orders and as such inclusion of more variables could reduce degrees of freedom severely.

cointegrating vector (Appendix A9). Normalising on s_t suggests a long-run relationship between sectoral GDPs of the form:

$$s_t = -1.4822^{***} + 1.1327^{***}a_t + 0.2512^{***}m_t$$

The estimated equation implies that a 1 per cent increase in agricultural GDP leads in the long-run to a 1.14 per cent increase in services output. On the other hand the same increase in manufacturing GDP results in 0.25 per cent increase in service GDP. Both the coefficients along with the intercept are significant at the one percent error probability level. As the Johansen cointegrtion procedure considers all the variables as jointly determined, these estimates are free from the endogeneity problems.[15]

When variables cointegrate, there must be at least one causality relationship involving the variables (Granger 1988). While the concept of causality has a strict statistical interpretation, it would be of interest to know, in addition to positive long-run association if output increase in agriculture can have causal effects on other sectors. The other sectors can also have causal effects on agriculture but for the purpose of this paper finding of a causal effect running from agriculture to other sectors would be particularly significant. The popular Granger causality testing procedure needs to be adapted to the integrated properties of the variables. Sims et al. (1990) show that if the variables are cointegrated of order 1, Wald tests of Granger non-causality in levels VAR could be used based on the error-correction model. However, Toda and Yamamoto (1995) provides a simpler statistical procedure to test for Granger causality involving non-stationary level variables. Appendix A12 reports the Toda-Yamamoto Granger causality test results, which show that agriculture has causal effects both on services and manufacturing. Both these effects are significant at the 5 per cent level. On the other hand, services and manufacturing are also found to have causal effects on agriculture but at a reduced level of statistical significance (i.e. at the 10 per cent level). There is no evidence of manufacturing sector's any causal effect on services and vice versa (at the conventional level of statistical significance).

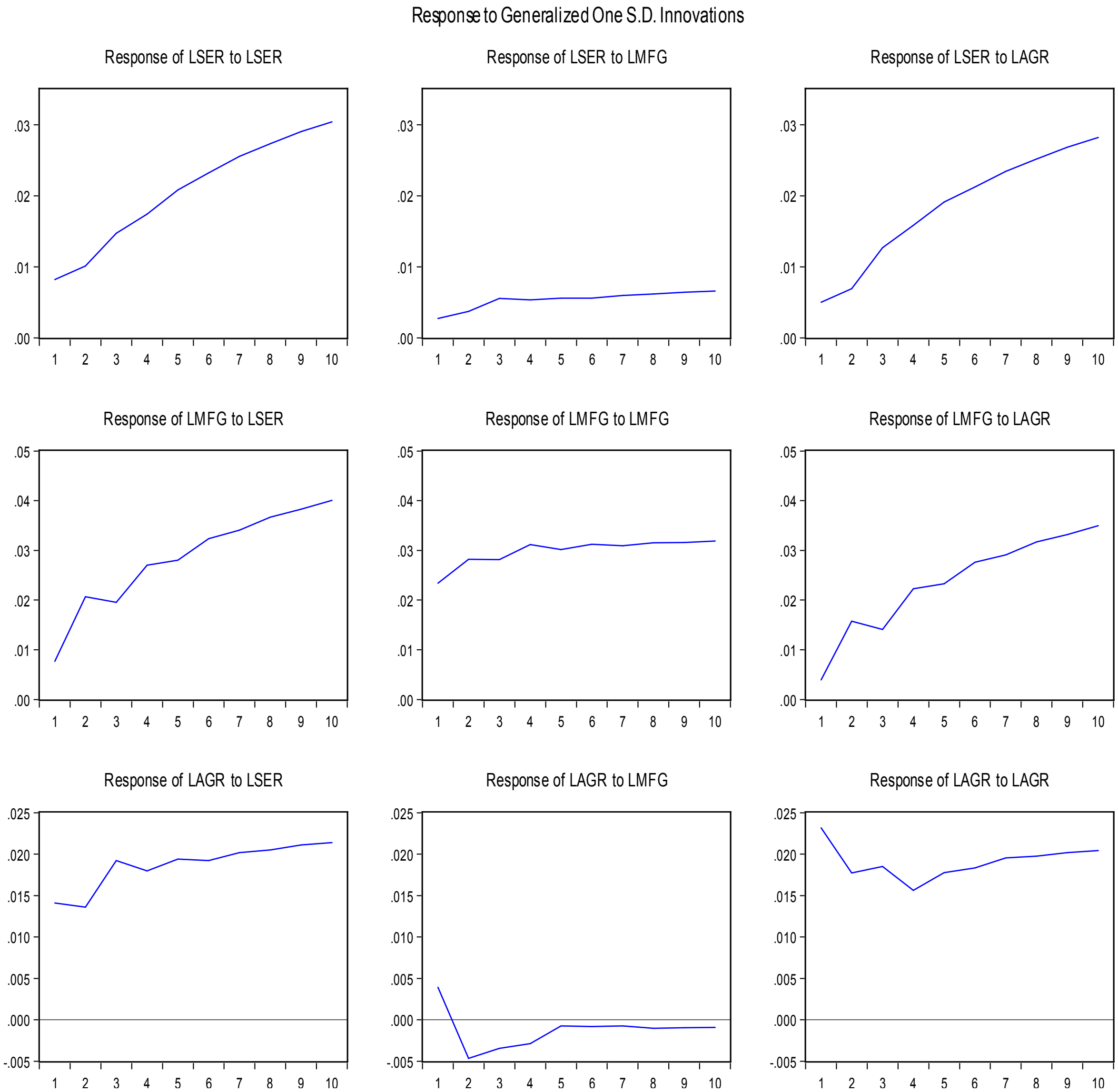
[15] However, that would also imply that the chosen normalization is arbitrary. This particular normalization was chosen because the services sector was the largest component of GDP, and we thought it would be interesting to know how agriculture and manufacturing activities contributed to services activities.

With 3 sectors included in our model, it is possible to construct vector error correction models (VECMs). However, like VAR models, the interpretations of VECMs may not be straightforward.[16] Instead, from policy perspective what is likely to be more useful is an impulse response analysis. An impulse response function (IRF) shows the dynamic behavior of a variable as given by its time path in response to a one-time exogenous random shocks given to this and other variables. For example, if the response of the services sector after a shock in manufacturing growth is positive, then presumably services will respond positively to innovations in services growth.

Following Pesaran and Shin (1998), the scaled generalized impulse response function is used here to trace the movement of a variable due to a shock. This approach may be considered as an improvement over the traditional and popular Cholesky approach that requires prior orthogonalization of the shocks, i.e. it can result in different impulse response functions depending on the order of the variables that are entered.

[16] In our case, the problem was further exacerbated by the fact that the weak exogeneity of agriculture and manufacturing was rejected by the data thereby not allowing us to use a single-equation based error correction model.

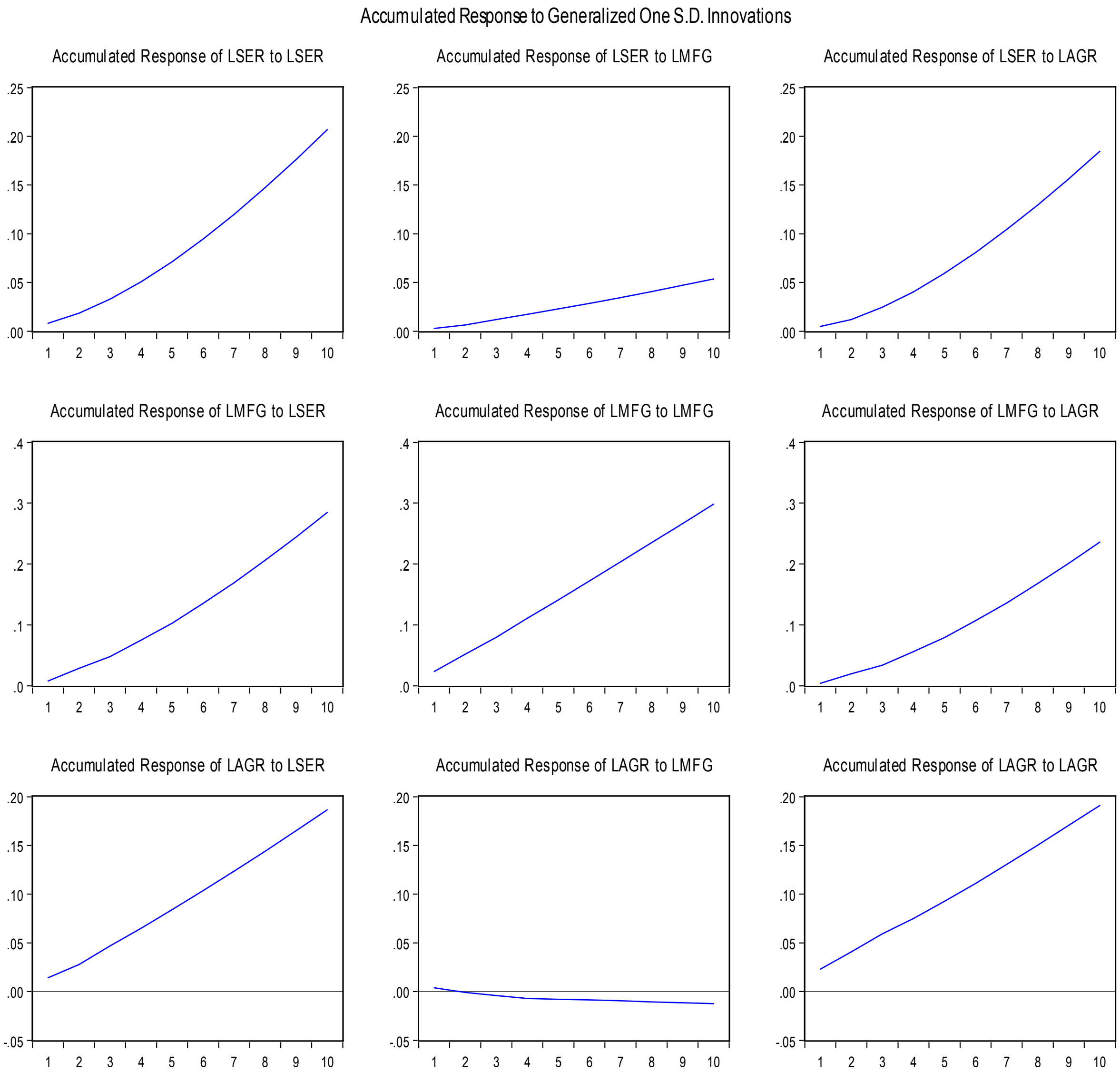
Figure 5: Generalised Impulse Response Functions



Note: LSER, LAGR, and LMFG represent for logarithmic transformations of services (s), agriculture (a), and manufacturing (m) outputs. The impulse response functions are derived from the estimated VECMs following the cointegrating equation. Impulse responses from a VECM, unlike from a stationary VAR, may not approach to zero as the horizon increases.

Figure 5 presents the responses to different sectoral GDPs to one standard deviation innovations in all three variables. Given the objective of this paper, our interest is to study the effects on other sectoral outputs following a one standard deviation shock on the agricultural sector. It is obvious that the services output depicts a strong and positive response due to a (positive) shock in agriculture. Almost a similar effect, although slightly fluctuating, is also exhibited by the manufacturing output. Amongst other inter-sectoral interactions, the responses of manufacturing and agriculture to services are also prominently positive. Figure 6 gives the accumulated responses over the time period considered. Clearly as the period specific responses to shocks in agriculture are positive for services and manufacturing, the accumulated further demonstrates sustained positive impacts.

Figure 6: Accumulated Impulse Response Function



Note: LSER, LAGR, and LMFG represent for logarithmic transformations of services (s), agriculture (a), and manufacturing (m) outputs.

V. Summary of Findings and Implications

This paper has made an attempt to understand the effect of farm production on the overall economic activity and sectoral outputs. The empirical assessment carried out here is based on a suitably adapted and extended dual sector analytical framework to deal with multiple intersectoral linkages and test for valid long-run relationship amongst variables using appropriate time series and econometric techniques.

The results associated with the dual sector model provide strong externality effects of agriculture. These effects are robust as they are maintained under different model formulations. When the the model is appropriately adapted to time series data properties, as proposed by Gemmell et al. (2000), there is evidence of 'cointegration' or a genuine long-run relationship between agriculture and overall economic output. The estimated long-run agricultural elasticity ranges from close to 1 to 1.60. The corresponding error-correction models satisfy usual properties and the positive effects of farm output growth are also borne out in the short run.

To clear out the natural growth-accounting effect of agriculture on overall GDP, empirical tests are carried out to ascertain a long-run relationship between farm and nonfarm outputs. There is also the evidence of cointegration between these two variables. The estimation of this relationship, using a methodology that treats both the variables as jointly determined, generates positive and highly significant effects of agriculture on the non-farm sector. In the short run as well, agriculture is found to be associated with the growth of the non-farm value added.

The paper also examines inter-sectoral linkages involving agriculture, manufacturing and services. It is found that these components of GDP move together and form a valid long-run relationship. The effects of agriculture on services are found to be quite large. Along with detecting causality effects running towards agriculture, more importantly for our case, there is also the evidence of agricultural growth causing outputs in other sectors. Finally, when impulse response functions are computed to trace the movements of different variables due to shocks provided to one sector, both agriculture and manufacturing are found to be depicting overtime growth in their activities due to a positive shock in agriculture.

There are important policy implications of the above findings. First and foremost, this paper provides consistent evidence that agriculture has significant positive spillover effects. As such, a policy emphasis to promote agriculture may not necessarily have adverse implications for other sectors.it is true that overtime the relative significance of the farm sector in the overall economy has declined quite considerably. Nevertheless, the findings of this paper should not be considered extraordinary. In the case of China, amongst others, agriculture is found to be exerting robust spillover effects despite the sector's greatly diminished significance (Yao, 2000).

Despite its output share being only about 20 per cent, agriculture provides for 45 per cent of employment opportunities in Bangladesh. Hence there exists an enormous scope of productivity improvement. Although the country has witnessed widespread adoption of HYV (high yield variety) technologies, it is widely recognised that agricultural production is still much less capital intensive compared to many other countries. Future productivity gains are likely to come from additional investment in this respect. This will not only bolster the firm sector's ability to provide food for population and raw materials for industrial sectors, but also tax revenues for government as well as saving generation for investment elsewhere.

It may be of great policy interest to understand the growth and productivity in the services sector, which is often regarded as a low productive area. The farm economy appears to exert large and significant positive influence on it. Movement of labour and saving out of agriculture to non-farm sectors can explain part of the inter-sectoral linkages. Recent evidence shows wages in agriculture are on the rise along with the growth in services, particularly the rural non-farm sector has flourished. This seems to indicate a more active role of agriculture in which it not only does provide capital and labour to other sectors, but also a huge market. The service-oriented rural non-farm and urban informal sectors have been considered to be the 'bridge' between commodity based agriculture and livelihood earned in the modern sectors, providing the transition from underemployment at farm tasks to regular wage employment in the local economy (Barrett et al., 2010). If services sectors are actually responding to increased demand of the farm economy, the farm-nonfarm linkages mark an important structural transformation process for Bangladesh.

Notwithstanding the spillover effects, an agriculture-focussed growth strategy will enhance the sector's ability to sustain a decent income growth for rural population thereby triggering immediate anti-poverty effects while ensuring a huge market for products and services for local industries. With its big population, Bangladersh has relatively a large domestic market, which implies that non-tradable and import-competing sectors are likely to be an important source of growth. Given its linkages, agricultural growth can boost economic activities in these sectors. Indeed, the findings of this paper suggest that an agriculture-focussed development strategy may not compromise with a growth maximising objective that will also make a powerful dent in poverty incidence.

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Appendix

Appendix A1: Unit Root Tests of the Variables

Variables	ADF test statistics		Variables	ADF test statistics		Conclusion
	Without trend	With trend		Without trend	With trend	
y_t (log of GDP)	6.4909**	-0.3196	Δy_t	-3.3054**	-4.7820**	$y_t \sim I(1)$
l_t (log of labour force)	-3.8968**	-2.8644	Δl_t	-3.0428**	-1.6395	$l_t \sim I(1)$
k_t (log of capital stock)	-1.1161	-3.5298	Δk_t	-3.4771**	-3.8044**	$k_t \sim I(1)$
a_t (log of agriculture)	2.7598	-1.0235	Δa_t	-4.1784**	-4.7776**	$a_t \sim I(1)$
na_t (log of non-agriculture)	4.8362**	-0.8181	Δna_t	-4.3152**	-5.4735**	$na_t \sim I(1)$
m_t (log of manufacturing)	3.9282**	-2.1181	Δm_t	-5.2709**	-6.6793**	$m_t \sim I(1)$
s_t (log of services)	7.3929**	1.5256	Δs_t	-3.0932**	-3.5827**	$s_t \sim I(1)$

Note: The 95 per cent simulated critical value using the corresponding sample size and 1,000 replications for the level variables without trend is -2.8607 and for models with trend is -3.6589. For first differenced variables, the comparable critical values are slightly different: -2.8738 and -3.5273 for models without and with the trend, respectively. ** indicates the rejection of null-hypothesis (i.e. the variable is non-stationary). All variables on their levels are strongly trended and this the ADF test with the trend term included is to be considered most appropriate in which none of the variables can reject the null hypothesis. All variables on their first difference reject the null-hypothesis of non-stationarity under the ADF test without trend. This is the most appropriate testing equation for growth rates (over a sufficiently long period of time, growth of any of these variables is unlikely to be trended) in which case we can conclude that all first difference variables are stationary. The inclusion of the trend term does not change this conclusion except in the case of growth of labour.

Appendix A2: Testing for cointegration in Johansen-Juselius procedure for variables *y, a, k,* and *l*

Null hypothesis	Alternative hypothesis	Computed statistics: Order of VAR = 1	Computed statistics: Order of VAR = 2	Computed statistics: Order of VAR = 3	95% Critical values
Cointegration LR test based on Maximal Eigenvalue of the Stochastic Matrix					
CV =0	CV =1	130.57	51.88	28.40	28.27
CV ≤1	CV =2	63.07	25.53	18.03	22.04
CV ≤2	CV =3	23.85	15.53	10.35	15.87
CV ≤3	CV =4	7.28	6.35	7.13	9.16
Cointegration LR test based on Trace of the Stochastic Matrix					
CV =0	CV =1	138.79	100.59	63.92	53.48
CV ≤1	CV =2	41.21	48.32	35.51	34.87
CV ≤2	CV =3	20.14	18.53	17.49	20.18
CV ≤3	CV =4	7.28	9.12	7.13	9.16

Note: CV stands for cointegrating vector.

Appendix A3: Testing for cointegration in Johansen-Juselius procedure for variables *y* and *a*

Null hypothesis	Alternative hypothesis	Computed statistics: Order of VAR = 1	Computed statistics: Order of VAR = 2	Computed statistics: Order of VAR = 3	95% Critical values
Cointegration LR test based on Maximal Eigenvalue of the Stochastic Matrix					
CV =0	CV =1	84.98	58.82	26.98	15.87
CV ≤1	CV =2	12.17	7.23	12.11	9.16
Cointegration LR test based on Trace of the Stochastic Matrix					
CV =0	CV =1	97.16	66.05	39.10	15.87
CV ≤1	CV =2	12.17	7.23	12.11	9.16

Note: CV stands for cointegrating vector.

Appendix A4: Testing for cointegration in Johansen-Juselius procedure for variables *na* and *a*

Null hypothesis	Alternative hypothesis	Computed statistics: Order of VAR = 1	Computed statistics: Order of VAR = 2	Computed statistics: Order of VAR = 3	95% Critical values
Cointegration LR test based on Maximal Eigenvalue of the Stochastic Matrix					
CV =0	CV =1	80.38	56.83	29.04	15.87
CV ≤1	CV =2	9.10	5.37	9.13	9.16
Cointegration LR test based on Trace of the Stochastic Matrix					
CV =0	CV =1	92.90	62.23	39.39	15.87
CV ≤1	CV =2	9.10	5.37	9.63	9.16

Note: CV stands for cointegrating vector.

Appendix Table A5: Choice of VAR order in the cointegration analysis between na and a

Order	Akaike Information Criterion	Schwarz Bayesian Criterion
3	193.56	183.89
2	197.11	190.67
1	192.76	189.55
0	112.03	112.07

Note: Based on 37 observations from 1976 to 2012. Order of VAR = 3

Appendix A6: Testing for cointegration in Johansen-Juselius procedure for variables *na* and *a*

Null hypothesis	Alternative hypothesis	Computed statistics: Order of VAR =2	95% Critical values
Maximal Eigenvalue test statistic			
CV =0	CV =1	56.85	15.87
CV ≤1	CV =2	5.37	9.16
Trace test			
CV =0	CV =1	62.23	20.18
CV ≤1	CV =2	5.37	9.16

Appendix A7: Testing for cointegration in Johansen-Juselius procedure for variables *na* and *a* (including a deterministic trend)

Sample (adjusted): 1976 2012
Included observations: 37 after adjustments
Trend assumption: Quadratic deterministic trend
Series: *na* and *a*
Lags interval (in first differences): 2

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.467785	23.74338	18.39771	0.0081
At most 1	0.010944	0.407178	3.841466	0.5234

Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.467785	23.33621	17.14769	0.0056
At most 1	0.010944	0.407178	3.841466	0.5234

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Appendix Table A8: VAR Lag Order Selection Criteria for the model with s, a, m

Sample: 1973 2012
Included observations: 37

Lag	LogL	LR	FPE	AIC	SC	HQ
0	133.3291	NA	1.75e-07	-7.044814	-6.914199	-6.998767
1	306.4684	308.8431	2.46e-11	-15.91721	-15.39475	-15.73302
2	324.2017	28.75677*	1.55e-11*	-16.38928*	-15.47498*	-16.06695*
3	329.7013	8.026418	1.93e-11	-16.20007	-14.89392	-15.73959

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan-Quinn information criterion

Appendix A9: Testing for cointegration involving s, a, m with a var order of 2

Included observations: 37 after adjustments
Trend assumption: Linear deterministic trend
Series: LSER LAGR LMFG
Lags interval (in first differences): 1 to 2

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.440763	35.69993	29.79707	0.0093
At most 1	0.211659	14.19622	15.49471	0.0777

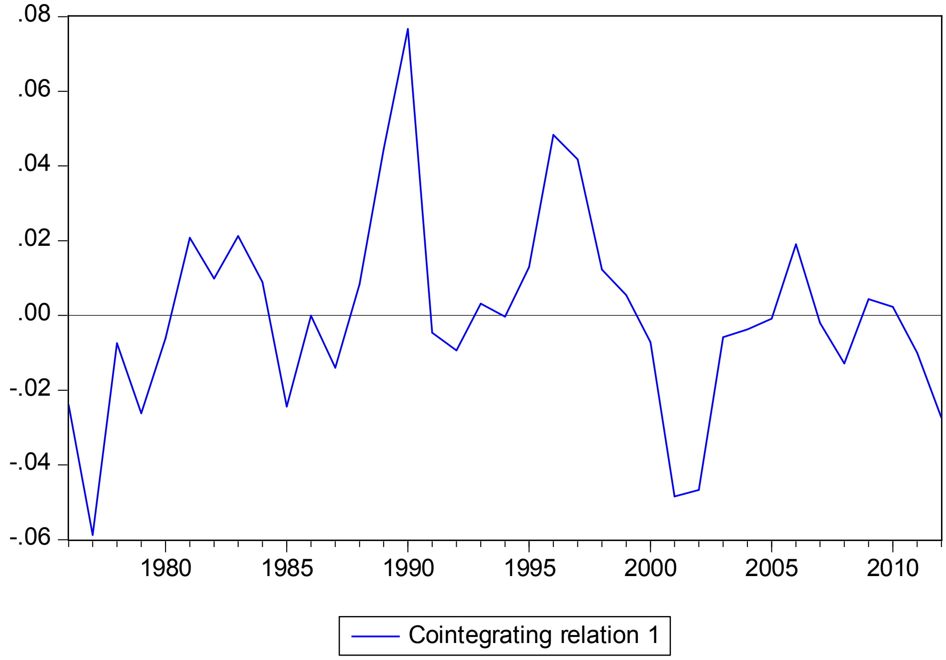
Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.440763	21.50371	21.13162	0.0443
At most 1	0.211659	8.799530	14.26460	0.3032

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Appendix A10: Cointegrating Relationship involving s_t, a_t, and m_t



Note: A trend term is inserted into the cointegration space.

Appendix A11: Cointegrating relationship and Vector Error Correction models

Vector Error Correction Estimates
Included observations: 37 after adjustments
Standard errors in ( ) & t-statistics in [ ]

Cointegrating Eq:	CointEq1		
LSER(-1)	1.000000		
LMFG(-1)	-0.251270		
	(0.06958)		
	[-3.61124]		
LAGR(-1)	-1.132727		
	(0.16019)		
	[-7.07113]		
C	1.482240		
Error Correction:	**D(LSER)**	**D(LMFG)**	**D(LAGR)**
CointEq1	-0.131355	-0.566388	0.508987
	(0.07432)	(0.21134)	(0.20952)
	[-1.76752]	[-2.67998]	[2.42932]
D(LSER(-1))	0.248765	1.146813	0.462023

	(0.15937)	(0.45321)	(0.44930)
	[1.56096]	[2.53043]	[1.02832]
D(LSER(-2))	0.231059	-0.409916	1.225075
	(0.16094)	(0.45770)	(0.45374)
	[1.43565]	[-0.89561]	[2.69992]
D(LMFG(-1))	-0.012031	0.021610	-0.289112
	(0.05543)	(0.15764)	(0.15628)
	[-0.21703]	[0.13708]	[-1.84997]
D(LMFG(-2))	0.008940	0.085208	-0.069919
	(0.03726)	(0.10597)	(0.10506)
	[0.23992]	[0.80407]	[-0.66553]
D(LAGR(-1))	-0.094991	0.203574	0.202336
	(0.08002)	(0.22756)	(0.22560)
	[-1.18710]	[0.89459]	[0.89690]
D(LAGR(-2))	0.064408	-0.098229	0.314386
	(0.07902)	(0.22471)	(0.22277)
	[0.81511]	[-0.43713]	[1.41123]
C	0.026373	0.019874	-0.041673
	(0.00921)	(0.02620)	(0.02598)
	[2.86226]	[0.75846]	[-1.60425]
R-squared	0.583224	0.542533	0.411801
Adj. R-squared	0.482623	0.432110	0.269822
Sum sq. resids	0.001960	0.015848	0.015575
S.E. equation	0.008220	0.023377	0.023175
F-statistic	5.797390	4.913229	2.900439
Log likelihood	129.6491	90.97864	91.29939
Akaike AIC	-6.575625	-4.485332	-4.502670
Schwarz SC	-6.227319	-4.137025	-4.154363
Mean dependent	0.047492	0.062837	0.029426
S.D. dependent	0.011428	0.031021	0.027121

Determinant resid covariance (dof adj.)	1.11E-11
Determinant resid covariance	5.36E-12
Log likelihood	322.6032
Akaike information criterion	-15.97855
Schwarz criterion	-14.80302

Appendix A12: Toda-Yamamoto (1995) Granger causality tests

VAR Granger Causality/Block Exogeneity Wald Tests
Date: 11/12/12 Time: 23:45
Sample: 1973 2012
Included observations: 37

Dependent variable: log(services)

Excluded	Chi-sq	df	Prob.
Log(agri)	7.347276	2	0.0254
Log (manu)	1.176402	2	0.5553
All	7.852451	4	0.0971

Dependent variable: LAGR

Excluded	Chi-sq	df	Prob.
Log (services)	5.361890	2	0.0685
Log(manu)	4.786962	2	0.0913
All	10.96604	4	0.0269

Dependent variable: LMFG

Excluded	Chi-sq	df	Prob.
Log(services)	4.330409	2	0.1147
Log(agri)	1.675048	2	0.0328
All	14.14037	4	0.0069

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