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Fiscal Sustainability in Major South Asian Economies: Evidences from Panel Data Analysis

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The paper examines the issue of fiscal sustainability for a panel of five major South Asian economies namely, India, Pakistan, Bangladesh, Sri Lanka and Nepal, for the period from 1985-2014. The results of panel cointegration tests by Pedroni (1999) and Westerlund (2007) confirm the presence of a long- run relationship between government revenue and expenditure. The panel auto regressive distributed lag (ARDL) estimates of the fiscal reaction function indicate a positive long-run response of the primary balance to the rising public debt ratio, thus confirming fiscally responsible behaviour in the region. However, the size of the cointegrating slope parameter between revenue and expenditure obtained from the group mean fully modified OLS (FMOLS) and the group mean dynamic ordinary least squares (DOLS) is significantly less than one, indicating weak form of fiscal sustainability. The weak sustainability underscores the need for commitment to long term fiscal discipline and justifies the ongoing efforts by the South Asian countries to strengthen their fiscal positions.

1. Introduction

The sustainability of public finances is currently a key policy issue in both developed and developing economies. An unsustainable fiscal position threatens both macroeconomic stability and the financial capacity of the state to deliver essential goods and services to citizens. Moreover, if fiscal positions are perceived to be unsustainable over the long term, then the

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reaction of the markets could trigger a fiscal crisis much sooner than might be expected by fiscal planners (Canagarajah et al., 2012).

Focusing on the issue of fiscal sustainability among developing economies, the case of South Asia deserves particular attention. Notwithstanding the high rate of real growth in the recent past, the fiscal situation in the region exhibits signs of distress. The reform process initiated in the late 1980s and the early 1990s in the region involved measures aimed at the simplification of tax systems, introduction of a value added tax and streamlining of expenditures. Due to the reforms, South Asia's fiscal deficits are decreasing gradually but remain consistently high relative to other developing regions (see Table 1). The persistence of the deficits may be attributed to the region's limited success in widening the tax base due to structural factors, such as large share of agriculture in national output, low literacy, large informal sectors, etc., coupled with populist spending raising prospects of unsustainable public finances.

Table 1: Fiscal Balance in South Asia vis a vis Developing Country Groupings

Country Groups	**2000-10**	**2011**	**2012**	**2013**	**2014**	**2016^{f}**
East Asia and Pacific	-1.6	.2	-.3	-2.3	-2.1	-2.1
Europe and Central Asia	-4.4	.7	-.6	-1.3	-1.5	-1.5
Latin America and the Caribbean	-2.6	-3.1	-3.6	-4	-5.2	-4.1
Middle East and North Africa	0.1	-4	-3.8	-6	-7.1	-5.3
South Asia	-7.4	-7.6	-7.2	-6.9	-6.7	-6.1
Sub Saharan Africa	-.6	-1.1	-1.7	-2.9	-2.5	-2.2

Source: Global Economic Prospects, January 2015
Note: f denotes forecast.

Against this backdrop, this paper aims to examine the sustainability of public finances for the five largest countries in South Asia namely, India, Pakistan, Bangladesh, Sri Lanka and Nepal, for the period from 1985-2014. The concept of fiscal sustainability implies the fulfilment of the so-called inter-temporal budget constraint (IBC), which states that the current level of debt in an economy should equal the present value of future fiscal surpluses. If this condition is to be met, economies cannot indefinitely issue debt to cover fiscal deficits, as the markets will observe a risk of bankruptcy (Carrion- i- Silvestre, 2015).

This paper seeks to employ two alternative approaches to test whether this condition is satisfied. First, we use co-integration techniques to assess whether there is a long-run relationship between the government revenues and expenditures. Second, we investigate whether the fiscal rule that relates the fiscal primary surplus and debt holds for the countries. The contribution of the paper to the existing literature is twofold. First, given the traditional problems related to the cointegration analysis using short-term data, we use panel data econometric techniques for assessment of fiscal sustainability. Single country time series estimation may suffer from shorter spans of data with associated smaller degrees of freedom and low power. Panel cointegration allows more variation in the data which could result in increased efficiency of the estimators. The empirical studies examining the issue of fiscal sustainability in a panel framework have mostly employed econometric techniques that impose homogeneity of slope coefficients across countries—an assumption that is likely to be violated. In this paper, we use panel approaches that allow for flexibility related to cross country heterogeneity. To the best of our knowledge, the present study is the first empirical application of these techniques in the assessment of fiscal sustainability for South Asian countries.

A further contribution is an evaluation of the degree of fiscal sustainability. This issue is usually disregarded in the existing literature pertaining to the selected countries, either because the analysis of sustainability is based on the assessment of interest –growth differentials (e.g., Ejaz and Javid, 2011; Mahmood, Arby and Sherazi,2014) or stochastic properties of public debt (e.g., Buiter and Patel, 2006; Deyshappriya, 2012) or because the estimates of the cointegrating vector between expenditure and revenue are not discussed (e.g., Jha, 2004; Kaur and Mukherjee ,2012; Munawar Shah et al., 2014).

The remainder of this paper is organized as follows: Section 2 provides a brief review of the literature. Section 3 discusses the data and methodology. Section 4 addresses the empirical assessment of fiscal sustainability, and finally, Section 5 offers concluding remarks.

2. Review of Literature

Measuring fiscal sustainability has been a highly contentious issue (Chalk and Hemming, 2000). Advancements in time-series techniques provided analysts with tools to test for the sustainability of fiscal policy via the fulfilment of IBC. The empirical contributions in this line can be broadly classified into two strands. The earlier studies following a univariate approach analyse the compliance to the budget constraint in terms of the mean-reverting behaviour of deficit and debt-GDP ratio series (e.g., Hamilton and Flavin, 1986; Wilcox, 1989). Second, the multivariate approach involves examining the long-run relationship between the flows of revenues and expenditures (e.g., Trehan and Walsh, 1988; Hakkio and Rush, 1991; Haug, 1991; Quintos, 1995).The subsequent studies refined the analysis by incorporating the possibility of structural changes that are associated with different degrees of sustainability (e.g., Quintos, 1995; Martin, 2000; Afonso, 2005) and generalized the definition of sustainability to distinguish between strict and weak sustainability (see the discussion below).

With advancements in cointegration techniques, the long run relationship between revenue and expenditure has been widely examined for assessing fiscal sustainability (e.g., Payne, 1997; Olekalns and Cashins, 1997; Hatemi-J,2002; Afonso and Jalles,2012; Dalgıç, İyidoğan and Balıkçıoğlu,2014).

However, Bohn's (1998, 2007) seminal work, challenges the time series literature on fiscal policy suggesting that stationarity-based sustainability tests are invalid because in an infinite sample, any order of integration of debt is consistent with the transversality condition, which implies that the IBC may be satisfied even if these particular time series tests are not. Moreover, the time series tests of fiscal sustainability do not explicitly identify the fiscal policy reactions underlying the data. As a result they do not shed much light on the kinds of policies that might deliver sustainability. (Adams, Ferrarini and Park, 2010)

Bohn (1998, 2007), therefore, suggested an alternative model-based approach to fiscal sustainability. In case fiscal authorities take corrective measures in response to deterioration in the debt position, rising debt ratios lead to higher primary surpluses relative to GDP, which indicates a tendency of debt towards mean reversion. Accordingly, a stable and strictly positive feedback from debt stock to primary surplus is a sufficient condition for fiscal sustainability. The model-based approach to fiscal sustainability has also been widely employed in the empirical literature (e.g., Prohl and Schneider, 2006; Ghatak and Sánchez-Fung, 2007; Ehrhart and Llorca, 2007; Afonso and Jalles, 2012; Carrion-i- Silvestre, 2015; Mackiewicz-Łyziak, 2015).

The empirical literature offers mixed results regarding the sustainability of fiscal positions of countries. The results vary according to the econometric techniques employed and the countries analysed. As noted by Westerlund and Prohl (2010) the weak empirical support for sustainability hypothesis by a number of studies may be due to the poor precision of commonly applied time series tests. To avoid this problem, some recent empirical attempts employ panel analysis for a group of countries since panel methods are considered more powerful than conventional time series methods(e.g., Ehrhart and Llorca(2007), Afonso and Rault (2007), Adedeji and Thornton(2008) ,Mercan (2013), Afonso and Rault (2013))

However, the empirical studies examining the intertemporal fiscal solvency constraint in the panel framework have almost exclusively focused on developed countries (OECD and EU members). The case of developing Asia and particularly that of South Asia is hardly explored. Moreover, the panel econometric techniques mostly employed in the aforementioned studies impose homogeneity of slope coefficients across countries. In view of this gap, the present study examines fiscal sustainability in major South Asian economies using panel approaches that allow for flexibility related to cross country heterogeneity.

3. Data and Methodology

The empirical strategy of the paper relies on two alternative approaches proposed by Hakkio and Rush (1991) and Bohn (1998) as discussed above. The theoretical framework underlying the first approach relies on the government's dynamic budget constraint and on the assumption that the interest rate follows a stationary stochastic process. Fiscal sustainability can then be tested through the following cointegrating regression[4]:

$$\mathbf{R_t = a + bG_t + u_t} \qquad \textbf{[1]}$$

where R_t and G_t are government revenue and expenditures inclusive of interest payments on debt; u_t is a stationary random variable and *a* and *b* are cointegrating parameters.

Following Quintos (1995), the fiscal sustainability exists in a "strong" form if and only if R_t and G_t are cointegrated and $b = 1$. The fiscal stance is instead only "weakly" sustainable if 0 <□*b* <1 in Equation [1]. Under this milder sustainability condition, government expenditures grow, on average, at a rate higher than government receipts. Finally, if the null hypothesis $b = 0$ cannot be rejected, the fiscal stance is unsustainable.
The alternative strategy proposed by Bohn (1998) involves estimating a fiscal reaction function (FRF) and checking whether there is a corrective response by the government to increases in the public debt. The model suggested by Bohn takes the following form:

$$\mathbf{S_t = \rho B_t + \alpha Z_t + \varepsilon_t} \qquad \textbf{[2]}$$

where S_t is the primary balance, B_t is the level of debt in the economy at the beginning of period t (which can be approximated by the level of debt in the period $t - 1$) and Z_t is a vector of explanatory variables that captures the economic cycles. The sufficient condition for sustainability requires $\rho > 0$ so that the government would take corrective actions – reducing the level of expenditure (excluding interest on debt) and/or increasing tax revenues – to offset the changes in the level of debt.

The empirical estimates in the paper are based on annual data on general government fiscal variables spanning from 1985-2014. The choice of the period as well as the selection of the five countries out of the eight in the South Asian region is based on the availability of consistent data. All the variables are measured in relation to GDP to obtain a more natural

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[4] see Afonso and Jalles (2012) for a formal proof

definition of sustainability that keeps pace with economic growth (Afonso, 2005) and to achieve similarly scaled series that offer easily interpretable information. The data are assembled from Key Indicators for Asia and the Pacific, the Asian Development Bank, World Economic Outlook Database and various International Monetary Fund (IMF) country publications and reports, such as IMF Article IV documents and Statistical Appendixes.

The stationarity properties of the data are tested using the Maddala and Wu (1999) test (M-W) and a second-generation panel unit root test, namely, the Pesaran's (2007) cross sectionally augmented IPS (CIPS) test. The M-W test combines the significance levels of individual Phillips- Perron or Augmented Dickey Fuller unit root tests for each cross-section *i* to construct an overall test statistic based on the test suggested by Fisher (1932):

$$\lambda = -2\sum_{i=1}^{N} \ln v_i \qquad [3]$$

where v_i is the p-value of the unit root test for country *i*. The statistic λ has a χ^2 distribution with 2N degrees of freedom where *N* denotes the number of panels. The null hypothesis tested is that all panels have a unit root versus the alternative that at least one panel is stationary. The M-W test is, however, based on a restrictive assumption that individual time series in the panel are cross-sectionally independent. The CIPS test relaxes this assumption and controls for the presence of cross-sectional dependence of the contemporaneous error terms. The test is based on a cross-sectionally augmented ADF (CADF) regression, which filters out the cross-sectional dependence by augmenting the ADF (p) regressions with the lagged cross-sectional mean and the lagged first differences of the cross -sectional mean. The estimated model is as follows:

$$\Delta y_{it} = a_{i0} + a_{i1}\, t + a_{i2} y_{i,t-1} + a_{i3}\bar{y}_{t-1} + \sum_{j=0}^{p} d_{ij}\, \Delta\bar{y}_{t-j} + \sum_{j=1}^{p} \delta_{ij}\, \Delta y_{i,t-j} + v_{it} \qquad [4]$$

Let $\tilde{t_i}$ denotes the t-ratio for a_{i1} in the above regression, the CIPS statistic is defined as:

$$CIPS = N^{-1}\sum_{i=1}^{N}\tilde{t_i} \qquad [5]$$

The null hypothesis of the test is the unit root for all the time series in the panel, while the alternative hypothesis is a stationary process for at least one of the time series.

To investigate the presence of a long-run relationship, both first and second -generation panel cointegration tests by Pedroni (1999) and Westerlund (2007) respectively are employed. Following Pedroni(1999), the estimated co-integration relationship between government revenue and expenditure is specified as follows:

$$R_{it} = \alpha_i + \delta_i\, t + \beta_i G_{it} + \varepsilon_{it} \qquad i=1,2,....N\,;\ t=1,2....T \qquad [6]$$

This formulation allows for considerable heterogeneity as follows: fixed effect (α_i), individual deterministic trend (δ_i) and a heterogeneous slope coefficient (β_i). The term $\varepsilon_{it} = \rho_i\ \varepsilon_{it-1} + v_{it}$ represents the estimated residuals from the panel regression. Pedroni (1999) developed seven statistics to test the null hypothesis of no cointegration. Panel v, panel rho, panel pp and panel ADF statistics are commonly referred to as within-dimension or panel cointegration statistics and the remaining three test statistics, group rho, group pp and group ADF are based on pooling along what is commonly referred to as between dimension or group mean panel statistics. For the panel cointegration statistics the alternative hypothesis is given by $H_a\colon\ \rho_i = \rho < 1\ ,\forall i$ and for the group mean cointegration test, $H_a\colon\ \rho_i < 1\ ,\forall i.$ The group mean cointegration tests ,thus, allow for a heterogeneous coefficient under the alternative hypothesis.

Along with the Pedroni test we also perform the Westerlund (2007) co-integration test which delivers robust critical values through the bootstrap approach even under the assumption of cross-section dependence. The test checks whether an error correction model has or not an error correction (individual group or full panel) based on the following equation:

$$\Delta R_i it = c_i(i) + \alpha_i i\,(R_i it - \beta_i i\, G_i(it-1)) + \sum_{j=1}^{pi} \theta_{ij}\, \Delta R_{it-1} + \sum_{j=1}^{pi} \gamma_j\, \Delta G_{it-j} + e_{it} \qquad [7]$$

where α_i is the speed of adjustment term. If $\alpha_i = 0$, there is no error correction and the variables are not co-integrated. If $\alpha_i < 0$, the model is error correcting implying that the variables are co-integrated. Westerlund (2007) developed four panel co-integration tests without any common-factor restriction. $P\tau$ and $P\alpha$ tests are designed to test the alternative hypothesis that the panel is co-integrated as a whole, whereas the two other tests, $G\tau$ and $G\alpha$ test whether at least one element in the panel is co-integrated.

The coefficients of the long-run relationship are obtained using group-mean estimators which allow for flexibility related to cross-country heterogeneity. In particular, we employ the group mean fully modified ordinary least squares (GM-FMOLS) and group mean dynamic ordinary least square (GM-DOLS) estimators. The GM -FMOLS estimator, derived by Pedroni (2000) uses the group mean of individual FMOLS estimators and corrects for endogeneity and serial correlation by estimating the long-run covariance directly. The estimator has satisfactory size and power properties even for small panels, if T is larger than N (Perdroni,2000). The group-mean panel FMOLS estimator for Eq. [6] can be written as:

$$\hat{\beta}_{GFM} = \frac{1}{N}\sum_i \left[\sum_{t=1}^{T} \frac{(G_{it} - \bar{G}_i) R_{it}^{*} - T\hat{\gamma}_i}{\sum_{t=1}^{T} (G_{it} - \bar{G}_i)^2} \right] \qquad [8]$$

where $R_{it}^{*} = (R_{i,t} - \bar{R}_i) - \frac{\hat{\Omega}_{21,i}}{\hat{\Omega}_{22,i}} \Delta G_{i,t}$ and $\hat{\gamma}_i = \hat{\Gamma}_{21,i} + \Omega^0_{21,i} - \frac{\hat{\Omega}_{21,i}}{\hat{\Omega}_{22,i}} (\Gamma\hat{\ }_i(22,i) + \Omega_i(22,i)^{\dagger}0)$. Here, $\hat{\Omega}_i = \hat{\Omega}^0{}_i + \hat{\Gamma}_i + \hat{\Gamma}_i{}'$ is the estimated long-run covariance matrix of the stationary vector consisting of the estimated residuals from the cointegration regression and the differences in government expenditures. $\Omega^0_{21,i}$ is the long-run covariance between the stationary error terms ($\varepsilon_i it$ **in equation 6**) and the unit root autoregressive disturbances. $\Omega_{22,i}^{0}$ is the long-run covariance among the difference in government expenditures. $\hat{\Gamma}_i$ is the weighted sum of the autocovariances and a bar over the letters denotes the mean for i members. The associated t-statistic for the between-group FMOLS estimator takes the following form:

$$t_{\hat{\beta}_{GFM}} = 1/\sqrt{N} \sum_{i=1}^{N} \left(\hat{\beta}_{FM,i} - \beta\right) \left(\Omega_{11,i}^{-1} \sum \left(G_{it} - \bar{G}_i \right)^2 \right)^{1/2} \qquad [9]$$

where β is a value under the null hypothesis. The above t-statistic is standard normal as T and N approach infinity.

Next, the GM-DOLS panel cointegration estimator is considered. DOLS uses the past and future values of ΔG_{it} as additional regressors to correct for endogeneity and serial correlation.. The between-group panel DOLS regression can be written as follows:

$$R_{i,t} = \alpha_i + \delta_i\, t + \beta_i G_{i,t} + \sum_{k=-ki}^{ki} \gamma_{ik} \Delta G_{i,t-k} + u_{i,t}^{*} \qquad [10]$$

$$\hat{\beta}_{DOLS} = \left[\frac{1}{N} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} Z_{i,t} Z_{i,t}' \right)^{-1} \left(\sum_{t=1}^{T} Z_{i,t} \hat{R}_{i,t} \right) \right]_1 \qquad [11]$$

where $Z_{i,t}$ is the 2(K+1)1 vector of regressors; $Z_{i,t} = \left(G_{it} - \bar{G}_i, \Delta G_{i,t-k}, --- \Delta G_{i,t+k} \right)$ and $\breve{R}_{i,t} = R_{i,t} - \bar{R}_i$. A bar over the letters denotes the mean, and the subscript 1 outside the bracket denotes the first element of the vector used to obtain the pooled slope coefficient. The associated t statistic for the group mean estimators is constructed as follows:

$$t_{\hat{\beta}_{DOLS}} = 1/\sqrt{N} \sum_{i=1}^{N} \left(\hat{\beta}_{D,i} - \beta\right) \left(\frac{1}{\hat{\sigma}_i^2} \sum \left(G_{it} - \bar{G}_i \right)^2 \right)^{1/2} \qquad [12]$$

where $\hat{\sigma}_i^2$ is the long run variance of the residuals from the DOLS regression and $\hat{\beta}_{D,i}$ **is the conventional DOLS estimator.** The t statistic is standard normal as T and N approach infinity. Both GM-FMOLS and GM-DOLS allow the control of the likely cross-sectional dependence by including common time dummies in the model.

For the robustness of results, we employ the common correlated effect mean group (CCEMG) estimator. CCEMG is a generalization of mean group estimator by Pesaran and Smith (1995) and is consistent in the presence of unobserved common factors proxied by the inclusion of cross sectional averages of dependent (R^- **in our case)** and independent variables ($\bar{G}$) in the regression setup. The model considered is:

$$R_{it} = \acute{\beta}_i G_{it} + c\,\acute{\square}_i\,\bar{R}_t + d\,\acute{\square}_i\,\bar{G}_t + \varepsilon\,\acute{\square}_{it} \qquad [13]$$

In the presence of a mix of stationary and non-stationary series, the fiscal reaction function may be estimated through the panel autoregressive distributed lag (ARDL) approach (e.g., Asiama, Akosah and Owusu-Afriyie, 2014; Akosah 2015; Waheed, 2016 etc.). In panel ARDL form the FRF equation may be written as:

$$\mathbf{PS_{it} = \alpha_i + \varphi_i B_{it-1} + \theta_i Z_{it-1} + \sum_{j=1}^{pi-1} \psi_{ij} \Delta PS_{i,t-j}} +$$
$$\mathbf{\sum_{j=0}^{ki} \varphi_{ij}^{*} \Delta B_{i,t-j} + \sum_{j=0}^{li} \theta_{ij}^{*} \Delta Z_{i,t-j} + \qquad + \mu_{it}} \qquad \mathbf{[14]}$$

where i = 1, 2, ..5, t = 1985,..2014 and μ_{it} is the error term assumed to be independently distributed across i and over t. The terms φ_i, θ_i represent long run coefficients; $\psi_{ij}, \varphi_{ij}^{*}, \theta_{ij}^{*}$ represent short run coefficients, and series PS, B and Z are as defined earlier. Following the basic specification of Bohn (1998), Z includes temporary factors impacting the primary balance including output gap (Ygap) and government expenditure gap (Ggap). The former captures the cyclical conditions and the latter accounts for unexpected expenditures, unrelated to the economic cycle, such as military expenditures. The output (expenditure) gap is measured as a ratio to potential output (expenditure) i.e. actual output (expenditure) less potential output (expenditure) divided by potential output (expenditure) where the potential output is proxied by the trend obtained with the Hodrick- Prescott filter. The possible non-linearities in the relationship are investigated by adding the quadratic term $(B_{t-1})^2$ as an explanatory variable.

Pesaran and Shin (1999) suggest two estimators for the ARDL model, the mean group estimator (MGE) and the pooled mean group estimator (PMGE). The PMGE allows short-run coefficients, including the intercepts, the speed of adjustment to the long-run equilibrium values, and error variances to be heterogeneous while the long-run slope coefficients are restricted to be homogeneous across countries.

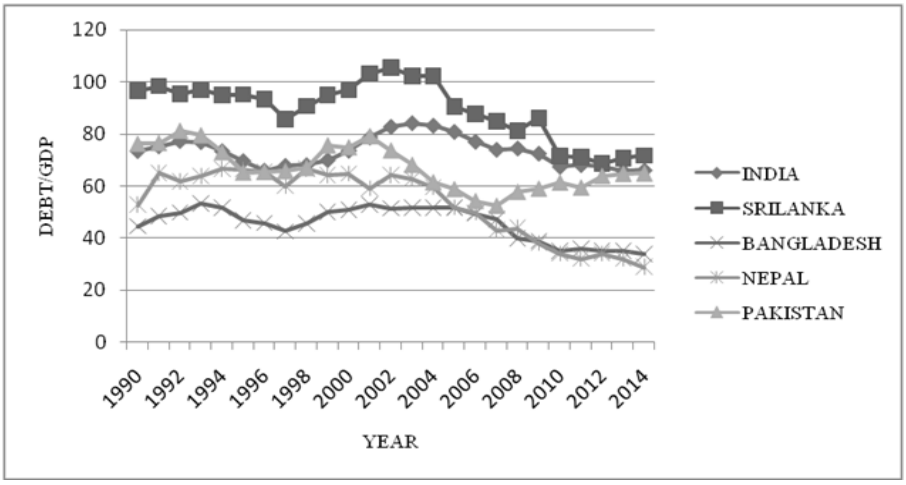
The MGE requires estimating separate regressions for each country and calculating the coefficients as unweighted means of the estimated coefficients for the individual countries. It allows for all coefficients to vary and be heterogeneous in the long-run and short-run.

The homogeneity test for long-run parameters is performed using the test suggested by Hausman (1978). Under the long run homogeneity assumption, both MGE and PMGE are consistent estimators, but only PMGE is the efficient estimator.

4. Empirical Analysis

Before the formal econometric analysis of fiscal sustainability, a brief characterization of the data may be appropriate at the outset. Generally speaking, the debt to GDP ratio exhibits a declining trend after 2000 for all the countries under considerations (see Figure 1). In case of India, the debt ratio registers a decline from 2003 which marks the period of implementation of the Fiscal Responsibility and Budgetary Management (FRBM) Act. An uptick in the debt ratio during 2007-08 may be observed in the case of India, Pakistan and Sri Lanka reflecting the effect of the Global Financial Crisis.

Figure 1: Public Debt to GDP: 1990-2014

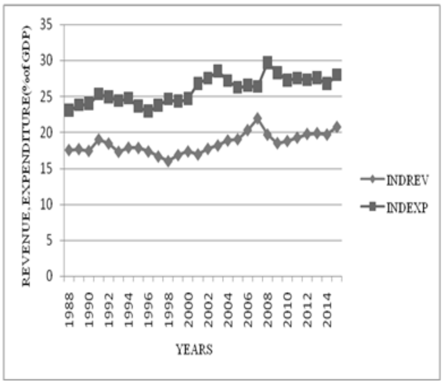


Source: World Economic Outlook Database, October2016.

Figures 2-6 capture the trends in revenue/GDP and expenditure/GDP for the countries. Among all economies, only Nepal has experienced a notable rise in its revenue/GDP ratio during the 1985-2014 period. The ratio has declined for Pakistan and Sri Lanka, and oscillated in a narrow range for India and Bangladesh. A preliminary inspection of the graphs indicates a synchronized behaviour of revenue and expenditure for all countries hinting at a possible long-run relationship. However, expenditure as a share of GDP on average exceeds the revenue share for all the countries which seemingly supports the weak sustainability hypothesis.

Figure 2: Revenue Expenditure (%of GDP)

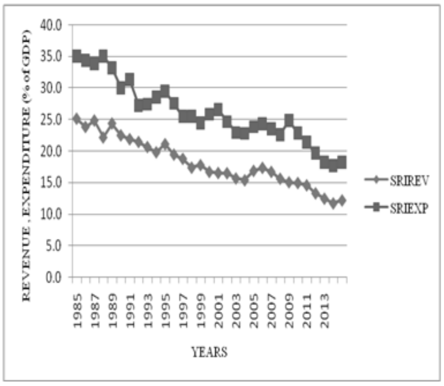
India



Source: World Economic Outlook Database

Figure 3: Revenue Expenditure (%of GDP)

Sri Lanka



Source: World Economic Outlook Database

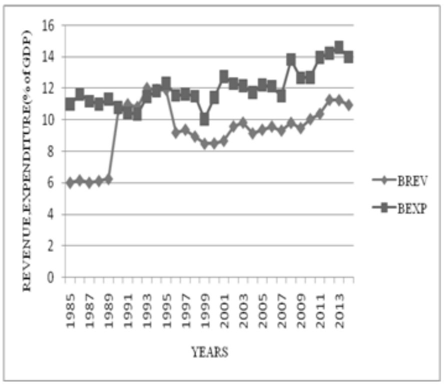
Figure 4: Revenue Expenditure (%of GDP)

Pakistan

Source: World Economic Outlook Database

Figure 5: Revenue Expenditure (%of GDP)

Bangladesh



Source: World Economic Outlook Database

Figure 6: Revenue Expenditure (%of GDP)

Nepal

Source: Key Indicators for Asia and the Pacific

Next, we focus on the econometric examination of sustainability conditions as discussed in Section 3. To set the stage for cointegration analysis, the series are subjected to the investigation of unit roots. The results of M-W and CIPS unit root tests reported in Table 2 show that the null hypothesis of I (1) series is rejected only for the PS series under both the tests[5]. The null is rejected for series R and G under the M-W test for the 0th lag only. The CIPS test, however, fails to reject the null for R, G and B series consistently for all lags. Therefore, whereas series R, G and B may be considered integrated of order one, series PS is I(0).

Table 2: Results of Panel Unit Root Tests

Test	R	G	B	PS
M-W	37.56*** (0) 20.00 (1) 10.85 (2) 14.02 (3)	42.59*** (0) 22.72 (1) 16.09 (2) 8.22 (3)	4.17 (0) 7.47 (1) 8.51 (2) 11.01 (3)	33.6*** (0) 20.2** (1) 22.7*** (2) 14.6 (3)
CIPS	0.48 (0) 2.047 (1) 3.043 (2) 3.064 (3)	-1.09 (0) -.378 (1) 1.95 (2) 2.62 (3)	.49 (0) .125 (1) .99 (2) .85 (3)	-4.0*** (0) -1.78** (1) -1.90** (2) -1.01 (3)

[5] At the outset the cross- sectional dependence in the data was tested using the Breusch Pagan LM test. The null of cross sectional independence was rejected for all series except revenue /GDP indicating the need to employ second generation panel unit root and cointegration tests.

*Note:****,** show significance at 1%and 5% respectively. Figures in parenthesis are lag lengths. An intercept and trend are included. Figures in parentheses are p-values. H0: series is I(1) for all panel, H1: series is stationary for some panel .

As a first step towards evaluating fiscal sustainability, the long-run relationship between revenue and expenditure is examined. The results of Pedroni (1999)'s panel cointegration test reported in Table 3 indicate that majority of within and between dimension tests reject the null hypothesis of no cointegration.

Therefore, it may be concluded that revenue and expenditure share a long run stochastic trend in the selected sample of South Asian countries for the period from 1985-2014.

Table 3: Results of Pedroni's Panel Cointegration Test

Panel cointegration (within dimension) statistics	
Panel v- statistic	1.04
Panel rho- statistic	-2.02***
Panel PP- statistic	-2.91***
Panel ADF- statistic	-3.04***
Group mean (between dimension) statistics	
Group rho- statistic	-1.00
Group PP- statistic	-2.59***
Group ADF- statistic	-2.68***

Note: Deterministic intercept and trend included. Automatic lag length selection based on SIC with a max lag of 5. Newey-West automatic bandwidth selection and Bartlett kernel. ***,** show significance at 1%,5% respectively.

Since the Pedroni cointegration test fails to take into account the cross-country dependence, a second generation cointegration test proposed by Westerlund (2007) is also employed . In small datasets, with T=30, Westerlund (2007) warns that the results of the tests may be sensitive to the specific choice of lag and lead lengths. Hence, to avoid overparameterization and the resulting loss of power, we hold short-run dynamics fixed by taking lead and lag=1. The results of the Westerlund (2007) test reported in Table 4 indicate that all four tests confirm presence of cointegration at 5%. The findings from second generation panel cointegration test thus corroborate those of first generation test.

Table 4: Results of Westerlund (2007) Panel Cointegration Test

Test	Value	P	Robust p
$G\tau$	-2.12	0.00	0.01
$G\alpha$	-9.6	0.00	0.00
$P\tau$	-4.6	0.00	0.02
$P\alpha$	-7.58	0.00	0.01

Note:The test takes no cointegration as null. The test regression is fitted with a constant and one lag and lead. The Kernel bandwidth is set according to rule $4(T/100)^{2/9}$. The p values are for one-sided test based on normal distribution. Robust p values are for one-sided test based on bootstrap replications.

After confirming cointegration, the next step is to estimate the coefficients of long-run relationship and draw inferences regarding the degree of sustainability. The estimates of long run coefficients are reported in Table 5.

Table 5: Long Run Coefficients under Alternative Estimators (Dependent Variable: Revenue)

	MG-DOLS		**MG-FMOLS**		**CCEMG**	
	Coeff	Chi square (H_0: β=1)	Coeff	Chi square (H_0:β =1)	Coeff	Chi square (H_0: β=1)
Expenditure	0.73	11.32 (0.00)	0.61	8.27 (0.00)	0.76	18.61 (0.00)
ECT	-0.243***		-0.236***		-0.218***	

Note: Figures in parenthesis are p values. *** indicates significance at 1%. To control for cross sectional dependence, MG-DOLS and MG-FMOLS models are estimated including time dummies.

As evident from the table, all three estimators, namely, MG-DOLS, MG-FMOLS and CCEMG indicate a positive slope coefficient. However, the chi square restriction on the coefficients consistent with 'strong' fiscal sustainability (b=1) is rejected at high significance levels for all three estimates indicating that the sustainability exists only in the 'weak' form. The 'weak' fiscal sustainability implies that government expenditures are systematically higher than government revenues. Therefore, although the IBC is satisfied in the strict sense (because the bubble term goes to zero), the upward pressure on the stock of debt is likely to increase the risk of default, forcing the government to offer higher interest rates to service its debt (Quintos, 1995).

The error correction term (ECT) estimated by inserting long run coefficients in short run dynamic specification of the models is negative and statistically significant under all estimators. The negative ECT shows that the system is driven to its long run cointegration path with a speed of adjustment of approximately 22-24% per year.

After examining the cointegration between revenue and expenditure we proceed to evaluate fiscal sustainability through the alternative strategy proposed by Bohn (1998, 2007). Given that the debt series is I(1) while the primary surplus series is I(0), the fiscal policy reaction function is estimated using panel ARDL model. As shown in Table 6 ,the long run coefficients of public debt is positive and statistically significant under both the MG and PMG estimators, indicating that fiscal authorities react systematically to rising public debt ratio to ensure fiscal sustainability.
The estimates indicate that the primary balance improves on average between 0.03 and 0.04 percentage point in response to a one percentage point increase in debt ratios thus confirming fiscally responsible behaviour.

Table 6: Results of Panel ARDL Estimates of Fiscal Reaction Function

Variable	**MGE**	**PMGE**
Long Run Coefficients		
$\mathbf{B_{t-1}}$	0.033***(0.009)	0.043***(0.016)
$\mathbf{(B_{t-1})^2}$	0.039***(0.012)	0.062***(0.011)
Ygap	0.07**(0.03)	0.063**(0.03)
Ggap	-0.029****(0.007)	-0.017***(0.0016)
Short Run Coefficients		
$\mathbf{\Delta PS_{t-1}}$	0.33*** (0.053)	0.41***(0.123)
$\mathbf{\Delta\ B_{t-1}}$	0.013*(0.007)	0.038*(0.017)
$\mathbf{\Delta(B_{t-1})^2}$	0.021*(0.0115)	0.0109*(0.0059)
$\mathbf{\Delta\ Ygap_t}$	0.113**(0.053)	0.144**(0.067)
$\mathbf{\Delta Ggap_t}$	-0.058*(0.032)	-0.074**(0.033)
ECT	-0.37***(0.047)	-0.33***(0.05)
Hausman Test	56.16***	

Note: MGE: Mean group estimator, PMGE: Pooled mean group estimator. Optimal lag structure based on Akaike's Information Criterion is ARDL (1,1,1,1,1). The null hypothesis under the Hausman test is that the difference in the estimated coefficients between the MG and PMG is not significantly different and that PMG is more efficient. *, **, *** indicate 1%, 5% and 10% levels of significance, respectively. Standard errors are shown in parentheses.

The response of primary balance to the squared debt term is also positive, significant and larger than the coefficient of lagged debt implying that the fiscal response is stronger when debt-to- GDP ratios are higher. The coefficient on output gap is positive implying that a positive shock to the output gap raises primary surplus by a factor of 0.06 -0.07 on average. The coefficient of expenditure gap is negative under both estimators indicating that an increase of expenditure above its potential lowers the primary surplus.

The short-run error correction models linked to the long run functions show that the explanatory variables maintain their long run signs in the short run as well. The coefficient of debt, however, is weaker in terms of statistical significance. The sign and significance of the estimated coefficient of lagged primary balance under both the estimators indicates that fiscal policy exhibits a strong degree of inertia in the region, causing the sign and magnitude of primary budgets in one year to depend substantially on the previous years' budgetary outcome and decisions. This should come as no surprise, as government budget plans typically run over several years, and many of the revenue and expenditure items are irreversible in the short term. The Hausman test statistics enables us to make a choice between the two estimators. The null hypothesis that MGE and PMGE are consistent estimators, but only PMGE is the efficient estimator is rejected implying that the appropriate estimator is MGE.

5. Conclusion

The study examines the issue of fiscal sustainability for a panel of five major South Asian economies namely, India, Pakistan, Bangladesh, Sri Lanka and Nepal, for period from 1985-2014, using two alternative approaches. First, we assess whether there is a long-run relationship between the government revenues and expenditures. Second, we investigate whether the fiscal rule that relates the primary surplus and debt holds for the countries.

The results of panel cointegration tests by Pedroni (1999) and Westerlund (2007) confirm the presence of a long- run relationship between government revenue and expenditure. The panel ARDL estimates of fiscal reaction function indicate a positive long-run response of primary balance to rising public debt ratio. However, since the size of the cointegrating slope parameter between revenue and expenditure is significantly less than one, fiscal sustainability exists only in weak form. The weak sustainability implies compliance to IBC in the strict sense but points to difficulties in marketing of future debt.

Though not indicating sustainability in a strong form, our results point to an improvement in the fiscal outlook for South Asia compared to a number of individual country studies. Ejaz and Javid (2011), Mahmood, Arby and Sherazi (2014) relying on an eclectic approach rejected the sustainability hypothesis for South Asian countries. Olekalen and Cashin (1997) for the period from 1951-1998 could not establish a long- run relationship between government revenue and expenditure in the case of India. Jha (2004) rejected the sustainability hypothesis for Nepal (for the period from 1960-1996) and Pakistan (for the period from 1956-1999) in view of the absence of a cointegrating relationship between revenue and expenditure. In contrast to the aforementioned studies, our study finds robust evidences in favour of compliance to IBC by using two alternative strategies. The evidence in favour of cointegration between revenue and expenditure, in contrast to the previous studies, may be attributed to the use of more powerful cointegration tests based on panel data, and to the coverage of fiscal data in the post reform period where governments in the region are taking a more cautious approach towards fiscal management.

To conclude, our results establish that South Asian countries have adhered to their intertemporal budget constraints in the post reform period. However, the sustainability exists only in a weak form which underscores the need to reinforce commitments to long-term fiscal discipline and justifies the ongoing efforts by the countries to strengthen their fiscal positions.

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