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# Growth-Maximising Fiscal Rule Targets in India*

KRISHANU PRADHAN

## Abstract

The paper estimates growth-maximising debt-to-GDP ratio (d*) in the Indian context from a theoretical framework of Ramsey-type growth model developed by David Alan Aschauer, and Cristina Checherita-Westphal and others. The output elasticity of public capital (α) plays a crucial role in determining d* under golden rule of budgetary deficit. Based on the estimate of α, the computed value of d* is around 65%–67% range, significantly lower than the current debt-to-GDP ratio (73% in 2016). Since a large share of India's fiscal deficit is due to the persistent revenue deficit, the effective value of d* would be even lower. The study results have some policy implications. Since the value of d* is significantly lower than the current level, and as the large revenue deficit is persisting, fiscal tightening by central and state governments should target to reduce not only fiscal deficit but more importantly the revenue deficit.

**Keywords**: public debt, fiscal deficit, macroeconomic stability, economic growth

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The importance of debt- and deficit-related fiscal rules, in India or elsewhere in the world, is firmly established both in the academic literature and in the policy arena. Most of the fiscal rules address macroeconomic vulnerability emanating from loose fiscal policy. The underlying compulsion of such rules or targets is the fiscal prudence essential for macroeconomic stabilisation. Although, in terms of intertemporal policy assignments, the monetary authority is largely responsible for macroeconomic stabilisation, the focus of fiscal policy should be on long-term structural issues like infrastructure development, provision of public goods, and economic growth. Thus, from the perspective of intertemporal policy choice, rigid fiscal rule targeting macroeconomic stabilisation is misaligned.

Most countries' fiscal and economic structures or compulsions are different, but there is a remarkable convergence of fiscal rule targets among countries.[1]

To attain the numerical fiscal rule targets concerning fiscal deficit or overall borrowing, governments are curtailing capital expenditures, as it is difficult to control committed and recurring expenditures. For instance, in India, the combined central- and state-government spending on capital expenditure came down from 5.2% of Gross Domestic Product (GDP) in 1990–1991 to less than 3% in 2002–2003 and remained below 4% of GDP till 2014–2015 (Government of India, 2010, 2018). Such a decline in public investment has occurred despite Fiscal Responsibility and Budget Management Act (FRBM) guidelines asking government borrowing for only public investment. Such stagnation or decline in public investment has detrimental effects on economic growth.

A major criticism of Maastricht Treaty convergence criteria has been that the numerical fiscal rule targets are not grounded in any theoretical framework of growth-maximising public debt ratio. Most of Maastricht

[1] For instance, Maastricht Treaty convergence criteria ask for eurozone countries to keep the fiscal deficit within 3% of GDP and debt-to-GDP ratio within 60%. Although the Indian economy is no match to eurozone countries, India has adopted the same numerical targets of deficit and debt-to-GDP ratios. Whether India's FRBM targets are just a mere coincidence, plagiarised, or derived from any logical explanation has been debated (Government of India, 2017; Gurumurthy, 2016).

Treaty's fiscal rule targets are criticised to be a set of arbitrary numbers. Macroeconomic stabilisation is required for sustained long-run growth, but decline in public investment to maintain fiscal rule targets might jeopardise macroeconomic stabilisation through lower economic growth. In this context, two relevant questions are: Can stagnation or decline in public investment to maintain fiscal rule targets ensure that public debt remains at a sustainable level?[2] If not, what would be the growth-maximising public-debt-to-GDP ratio that India can target through fiscal rules for public debt sustainability?

These questions are important as a drop in public investment reduces the availability of public goods necessary for private capital formation or the private sector. Provision of public goods, even though financed by government borrowing, might keep public debt at sustainable levels through the beneficial impact of higher economic growth. Apart from such provision, productivity of public capital, which varies across countries, is also crucial in determining the sustainable level of public debt. That is why, to derive meaningful fiscal rule targets for maximising economic growth (subject to macroeconomic stability), information on the productivity of public capital is crucial in determining an optimal level of public-debt-to-GDP ratio. Productivity of public capital may well determine the growth-maximising public debt ratios (Checherita-Westphal et al., 2012). Therefore, it is important to estimate the growth-maximising public debt ratios from the productivity of public capital. These debt ratios or fiscal rule targets would no longer be arbitrary numbers as the fiscal rules literature suggests, but they would be obtained from the long-run optimising behaviour of the private sector. Thus, the estimated optimal debt-to-GDP ratio may also examine the consistency of official fiscal rule targets from the perspective of growth maximisation in India.

[2] Literature on fiscal rules across countries increasingly recognises the importance of targeting stock concept of fiscal imbalance—that is, public debt—rather than flow concepts, such as year-on-year fiscal deficit or revenue deficit. See Checherita-Westphal et al. (2012) and the N. K. Singh committee report (Government of India, 2017) for further discussion. That is why the focus of this study is to determine growth-maximising public debt ratio in India.

Studies by Reinhart and Rogoff (2010), Checherita-Westphal and Rother (2010), Cecchetti et al. (2011), and Baum et al. (2013) obtained the non-linear relationships between an economy's debt burden and its rate of growth. Another strand of study by Reinhart et al. (2003) showed that the estimated 'debt intolerance' level of external public debt varies remarkably, for groups of countries, from less than 40% to over 100% of GDP. The correlation between economic growth and public debt ratio in general, or external debt default in particular, suggests that there exists a non-linear relationship between public debt ratio with economic growth and debt default. The essence of the non-linear relationship highlights that there is an optimal level of public debt ratio for each economy beyond which either economic growth starts declining or the country heads towards debt default. Being purely empirical in nature, these studies did not provide any theoretical foundation for determining optimal public debt ratios. Thus, determining the optimal level of public debt for countries from purely empirical literature may fail to adequately explain what determines the optimal public debt ratios. Checherita-Westphal et al. (2012) highlighted the limitations of such arbitrarily determined fiscal rule targets in different countries. Their study also mentioned the limitations of empirical literature which identifies the non-linear relationship between public debt ratios and economic growth without providing an idea about what that optimal level of public debt depends upon.

In this context, the objective of the paper is to estimate the growth-maximising public debt ratios for India from a simple theoretical framework developed by Aschauer (2000) and Checherita-Westphal et al. (2012). The recent N. K. Singh Committee Report (Government of India, 2017) too stressed the importance of targeting general government debt level with fiscal deficit as the operational targets, from medium- to long-term perspectives in management of fiscal policy.

## Theoretical Framework

In this paper, the theoretical framework to explain the link between public debt ratios and economic growth is based on Aschauer (2000) and Checherita-Westphal et al. (2012). Aschauer (2000) modelled the growth-optimising behaviour of the ratio of public capital to private capital and to national income. Aschauer (2000) modelled the determination of these ratios under the framework of Ramsey–Cass–Koopmans-type growth model. Checherita-Westphal et al. (2012) extended Aschauer's work of determining the optimal ratio of public capital to private capital and income to the determination of optimal public-debt-to-income ratio. It is assumed that under the golden rule of budgetary deficit, borrowing is made for the creation and accumulation of only public capital. That means, borrowing for public consumption or revenue expenditure is not allowed. In other words, revenue deficit in the budget should be kept nil.

The derivation of equation-optimal public-debt-to-GDP ratio has used Ramsey-type growth model with a three-input production function, namely labour (L), private capital ($K_p$), and public capital ($K_g$) (Checherita-Westphal et al., 2012). The production function can be written as

$$Y = \{[\ L^{\beta}\ K_p^{\ (1-\beta)}\ ]\}^{(1-\alpha)}\ [K_g]^{\alpha} \qquad (1)$$

where $0 < \alpha, \beta < 1$ and the constant return to scale between public and private sector inputs and between private inputs is assumed for sake of the steady state solution to be convergent.

According to Checherita-Westphal et al. (2012), the optimal public-debt-to-GDP ratio ($d^*$) is equivalent to the expression of optimal public capital stock ($K^*_g$) to GDP under the assumptions of golden rule of budgetary deficit, and it depends on the output elasticity or productivity of public capital stock. Following Checherita-Westphal et al. (2012), this can be expressed as[3]

$$(D/Y)^* = d^* = \{\alpha\ /\ (1 - \alpha)^2\}^{1 - \alpha} \qquad (2)$$

where $\alpha$ and $(1 - \alpha)$ are the output elasticity of public capital stock and private inputs respectively, Y is GDP, D is the nominal stock of debt, and d is the debt-to-GDP ratio. The * indicates the optimal level of debt and $d^*$ is the long-run optimal debt ratio when all inputs and input ratios take their optimal values. As the optimal value of $d^*$ depends on

[3] For detailed derivation and discussion on the determination of optimal public-debt-to-GDP ratio, see Aschauer (2000) and Checherita-Westphal et al. (2012).

optimal value of $K^*_g$-to-GDP ratio,[4] the next section briefly describes the evolution of private and public capital stocks to GDP in India from 1981 to 2012.

## Evolution of Capital-Stocks-to-GDP Ratio (1980–1981 to 2011–2012)

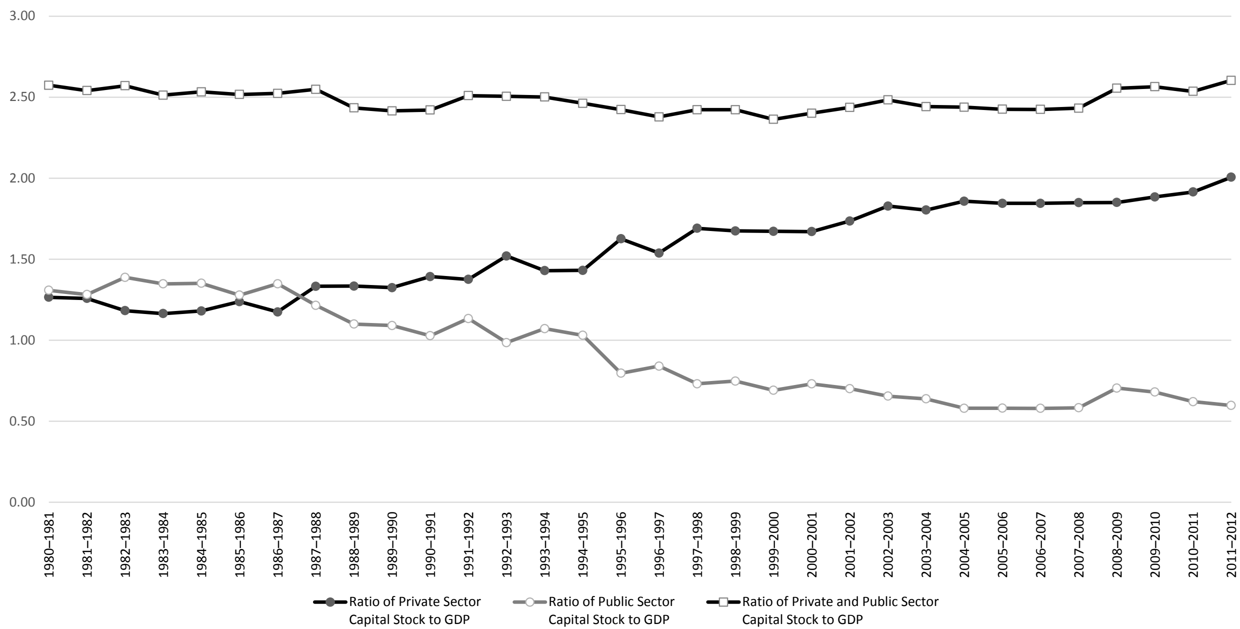
Figure 1 shows that the ratio of public sector capital stock to GDP experienced a secular decline, from 1.3 in the 1980s to 0.7 in 2003. The decline in public investment, which started since the balance of payment (BoP) crisis in 1990–1991, accelerated after 1994, owing to International Monetary Fund (IMF)–led structural adjustment programme and the limited fiscal capacity of government in financing investment. Figure 1 also reveals that public sector investment dominated overall investment in the country till around the mid-1980s. This was due to a significant increase in the number of central public sector enterprises (CPSEs), from fewer than 50 in 1961, to 179 in 1981, and 244 in 1990 (Government of India, 2011). Mirroring the increase in CPSE units, government investment increased from less than ₹40 billion in 1961 to over ₹180 billion in 1981 and ₹1,000 billion in 1990 (Government of India, 2011). Thus, the evolution of public investment has largely been shaped by investment in CPSEs during the 1980s and fiscal capacity constraint after the mid-1990s.

On the other hand, the ratio of private sector capital stock to GDP has experienced a persistent increase since the late 1980s. The average share of private sector in capital formation has been twice that of public sector after 1995–1996 and mostly thrice after 2004–2005. In recent years, the role of private sector capital formation by household sector and corporate sector has been a significant driving force behind India's impressive growth performance. The overall ratio of capital-stock-to-GDP ratio hovered around 2.4 to 2.5 during 1981 to 2012. Therefore, what has changed is the transition from government-supported capital

[4] It depends on the ratio of private inputs to public capital. The private capital in Aschauer (2000) incorporates all types of capital, including human capital.

**Figure 1**

*Trend and Pattern of Private, Public, and Aggregate Capital-Stock-to-GDP Ratio (1981 to 2011–2012)*




*Note*. Author's compilation using RBI (2014, 2016) and Das et al. (2015).

formation to private sector–led investment in India. That is why public goods or investment which are crucial inputs for the private sector should not compete or replace private capital or investment. Their optimal level should be determined by their productivity.

## Study Period, Data, and Variable Description

The period of this study is from 1980–1981 to 2011–2012. The study period is not extended beyond 2011–2012 as shifting of base year from 2004–2005 to 2011–2012 in estimating GDP with new methodology is underway in India. Not all National Accounts Statistics (NAS) data have been updated or revised. Also, debates on the revision of base year and the new methodology of estimating GDP are yet to be settled.

Table 1 describes the variables and lists the data sources.

## Empirical Framework

To compute the steady-state ratio of debt-to-GDP derived from Equation 1 in determining the growth-maximising debt-to-GDP ratio, the estimated value of output elasticity ($\alpha$) of public capital ($K_g$) is required. Taking the natural log (ln) transformation of Equation 1 we get

$$\ln(Y) = (1 - \alpha)\ \{\beta \ln(L) + (1 - \beta) \ln(K_p)\} + \alpha \ln(K_g) \qquad (3)$$

Expressing Equation 3 using the condition $\gamma = \beta(1 - \alpha)$ and $\delta = (1 - \beta)(1 - \alpha)$, we get

$$\ln(Y) = \gamma \ln(L) + \delta \ln(K_p) + \alpha \ln2(K_g) \qquad (4)$$

Expected sign: (+) (+) (+)

The expression ($\gamma + \delta$) is the share of two private inputs, L and $K_p$ in the economy, that is, ($1 - \alpha$). The estimation of parameters from Equation 4 would be done by application of time series econometric estimation techniques. The predicted sign of estimated $\gamma$, $\delta$, and $\alpha$ would be positive because they are respectively, the output elasticity of labour (L), private capital ($K_p$), and public sector capital stock ($K_g$).

**Table 1**

*Description, Measurement of Data Sources of Variables*

| Variable | Description | Data sources |
|---|---|---|
| GDP (Y) | Gross domestic product at constant market prices with base year 2004–2005. | RBI (2016) |
| Private Sector Capital Stock ($K_p$)<br><br>Public Sector Capital Stock ($K_g$) | Data on public and private sector capital stock (in constant price 2004–2005 base year) are not available separately for India. National Accounts Statistics of the Central Statistical Office (CSO) provides the database on capital stock by types of institutions in 1999–2000 prices from 1981 to 2008. On the other hand, database on the net aggregate non-financial capital stock at constant price (2004–2005 base year) from 1980–1981 to 2011–2012 is made available by the India KLEMS Database 2015. Using data on gross capital formation from RBI's H*andbook of Statistics on Indian Economy 2014*, the share of private and public sector capital formation is determined. Applying the share of public and private capital formation, the stock of public and private sector capital is derived from the India KLEMS database 2015. | RBI (2014); Das et al. (2015) |
| Labour employment (L) | In India, continuous time series data on organised sector employment are available, while that of unorganised sector are available quinquennially. As a consequence, continuous time series data on overall employment in India are yet to be produced by any government agency. To overcome this difficulty, the data on labour employment in the India KLEMS database are used. It is argued by Checherita-Westphal et al. (2012) that it is more relevant to consider data on labour hours worked annually than just the number of employed. In this regard, database of PENN World Table for India on average annual hours worked by persons engaged is used. | Das et al. (2015); Feenstra et al. (2015) |

## Techniques of Estimation

### Unit Root Testing

Table A1 reports the descriptive statistics of variables under study—ln(Y) ln($K_g$), ln($K_p$), and ln(L). Other than of ln($K_g$), kurtosis values of variables are well below 3. That means variables other than ln($K_g$) do not suffer from the problem of heavier tail distribution. Kurtosis values of ln($K_g$) are slightly higher than 3, which indicate a modest leptokurtic distribution. Similarly, other than of ln(L), skewness values of variables are positive but less than one, and hence slightly positively skewed, whereas ln(L) is slightly negatively skewed.

Since the variables of interest in the study are time series in nature, one needs to check for their stationarity (i.e., presence or absence of stochastic trends). Otherwise, we may end up with spurious regression estimates. Following Dickey and Fuller (1979, 1981) and Phillips and Perron (1988), the existence of stochastic trends or unit roots in ln(Y), ln($K_g$), ln($K_p$), and ln(L) is checked by Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests, with different specifications in deterministic trends. If a variable is stationary at level, then it is integrated of order zero, that is, I(0). If the same variable becomes stationary after first or second difference, then it is integrated of order 1 or 2, that is, I(1) or I(2). The results of the ADF and PP tests statistics (reported in Table A2) indicate that the null hypothesis of unit root of variables in level cannot be rejected. However, the variables become stationary after first difference. In other words, all variables are non-stationary at level while stationary after first difference.

### Cointegration

Since the variables are I(1) in level, an application of ordinary least squares (OLS) techniques to estimate the parameters may be spurious unless the variables are cointegrated or serial correlation problem is addressed. Existence of cointegration (CI) would provide long-run relationship among the variables. However, if serial correlation is detected among the residual errors, we need to correct it. Otherwise, the estimates, though unbiased, would not be efficient.

We can check the presence of CI by either the Engle–Granger (1987) CI test or the Johansen (1988) CI test. The Engle–Granger (E–G) CI is a two-step procedure which tests whether the linear combination of I(1) variables is stationary. In other words, the E–G test suggests that variables which are non-stationary, that is, I(1) in levels, may be cointegrated if their linear combination is stationary, that is, I(0). The E–G CI test is based on the Dickey–Fuller (D–F) or the ADF tests for unit roots in the residuals from single equation estimates. According to Stock (1987) and Enders (2014), if variables are cointegrated, an OLS regression would provide a *super consistent* estimator of the parameters. On the other hand, the Johansen CI test is more generalised than the E–G test as the former permits to check the existence of more than one cointegrating relationship among variables. In other words, the Johansen CI test can be applied to more than one cointegrating vector, whereas the E–G test is applicable for one cointegrating vector. However, if only one cointegrating vector is found when more than two variables are involved, the E–G and Johansen CI tests would provide the same conclusion. There would be no substantive complication; the only difference would be that the level regression and Error Correction Model (ECM) would involve more variables (Patterson, 2002). Johansen (1988, 1991) proposed two types of tests, namely, trace tests and maximum eigenvalue tests, to determine the number of CI relations.

Parameters of Equation 4 would be estimated by applying OLS and fully modified ordinary least squares (FMOLS) after ensuring that CI exists following E–G and Johansen CI tests. Further, if the serial correlation problem is found among the residual errors in OLS estimation, it would be corrected by incorporating an autoregressive (AR) term in the regression. It would help us to compare the estimated value of coefficient of $K_g$ (i.e., $\alpha$) which is crucial in estimating the growth-maximising debt-to-GDP ratio under OLS and FMOLS estimations and serial correlation–corrected estimates.

## Engle–Granger Cointegration Tests and OLS Estimates

Equation 4 is estimated by OLS techniques, and the estimated results are reported in Table 2. The results of the OLS estimates suggest that the coefficients of $\ln(K_p)$ and $\ln(K_g)$, that is, $\delta$ and $\alpha$, are positive and statistically significant while that of ln(L), that is, γ, is positive but statically insignificant. The output elasticity of public capital (α) is 0.285. Then the computed output elasticity of two other inputs (1 - α), namely, L and $K_p$, taken together would be 0.715.

**Table 2**

*Estimates of Equation 4 by Ordinary Least Squares*

| Variable | Coefficient | Coefficient value | t-statistic | Prob# |
|---|---|---|---|---|
| ln(L) | γ | 0.068 | 0.141714 | 0.8883 |
| ln($K_p$) | Δ | 0.637* | 20.52020 | 0.0000 |
| ln($K_g$) | A | 0.285* | 8.635859 | 0.0000 |
| Constant | C | 0.11 | 0.980634 | 0.3352 |
| $R^2$ | 0.998870 | | | |
| Adj. $R^2$ | 0.998749 | | | |
| DW stat | 1.22 | | | |
| F-statistic | 8253.165 | | | |
| P-value | 0.000 | | | |

*Note.* * indicates coefficient is significant at 1% level of significance.

The estimated value of the DW statistic (1.22) is greater than adj. $R^2$ = 0.9987; and as per the 'rule of thumb' one can reject the presence of spurious regression (Ghatak & Ghatak, 1996). However, recent literature on spurious regression, especially the classic review article by Santaularia (2009), has highlighted problems in applying rule-of-thumb criteria to conclude about spurious regression. In the present study, the estimated DW value for k = 3 (for three explanatory variables excluding the intercept) and N = 32 (number of observations) at 5% significance level is lower than the tabulated lower level (i.e., tabulated $DW_L$ = 1.244) DW value. The decision rule is that if the estimated DW value is positive but lower than the lower value of tabulated DW statistic (i.e., $0<DW<DW_L$), then we reject the null hypothesis of no positive auto-correlation. It means that residual errors follow first-order serial correlation. Moreover, the Breusch–Godfrey serial correlation Lagrange Multiplier (LM) test results from Table 3 reveal the presence of first-

**Table 3**

*Breusch–Godfrey Serial Correlation Lagrange Multiplier Test Results*

| | | | |
|---|---|---|---|
| Test for first-order serial correlation | | | |
| F-statistic | 4.295467 | Prob. F(1,27) | 0.0479 |
| Obs *R-squared | 4.392168 | Prob. Chi-Square(1) | 0.0361 |
| Test for second-order serial correlation | | | |
| F-statistic | 2.280866 | Prob. F(2,26) | 0.1223 |
| Obs *R-squared | 4.776412 | Prob. Chi-Square(2) | 0.0918 |

*Note*. Optimal lag length is 1 after considering the AIC and SIC statistics.

order serial correlation among residual errors. The second-order serial correlation is not significant at 5% level, though it is significant at 10% level. In the present study, since 5% significance level is considered as standard, and the optimal lag length is 1, after taking into account Akaike information criterion (AIC) and Schwarz–Bayesian information criterion (SIC) statistics, the second-order serial correlation is not taken into account for further analysis.

Therefore, though the rule of thumb is satisfied, the DW test statistic and LM test results do not rule out the presence of first-order serial correlation among the residual errors at 5% level of significance. Under this circumstance, we have two choices—either go for E–G CI test or correct the OLS estimates by including an autoregressive term of order 1, that is, AR (1), in the analysis.

E–G CI tests by applying the D–F tests to the estimated residual errors is attempted. We need to check whether the estimated residual errors denoted as $\mu_t$ from Equation 4 follow unit root. The values of estimated errors are nothing but the deviations from the long-run relationship among the variables concerned. If the estimated errors from Equation 4 are found to be stationary, then variables are cointegrated. Following E–G CI procedure, we perform the unit root test on residual errors of estimated Equation 4. Equation 5 represents the D–F unit test results of estimated errors of Equation 4.

$$\Delta\mu_t = -0.63\mu_{t-1} \quad (5)$$

SE = (0.17) and ‘$\tau$’= -3.8, adj. $R^2 = 0.33$, DW = 1.89 and SIC(7) = -5.2.

The coefficient of $\mu_{t-1}$ (-0.63) being statistically significant at 5% level implies the absence of serial correlation among the residual errors, and hence we may reject the possibility of spurious regression. Thus, the E–G test of CI suggests that ln(Y), ln(L), ln($K_p$), and ln($K_g$) are cointegrated.

Since DW test statistic and LM test indicate the presence of first-order serial correlation, we need to correct the OLS estimates by incorporating the AR(1) term in the regression analysis. Table 4 reports the estimated value of coefficients after correcting for serial correlation. The estimated value of AR(1) coefficient (Ø) being 0.38 (i.e., $<1$) and statistically significant indicates that the convergence is achieved. There has been a slight increase in the estimated value of ln ($K_g$), that is, α, to 0.29. Other coefficient values too have changed slightly after correcting for serial correlation. The computed output elasticity of other inputs (i.e., L and $K_p$) after correcting serial correlation is 0.71, which is not very different from 0.715 under OLS estimation.

**Table 4**

*Serial Correlation Corrected Estimates of Equation 4*

| Variable | Coefficient | Coefficient value | t-statistic | Prob. |
|---|---|---|---|---|
| ln(L) | γ | 0.06 | 0.62 | 0.5431 |
| ln($K_p$) | δ | 0.63* | 16.3 | 0.0000 |
| ln($K_g$) | α | 0.29* | 7.6 | 0.0000 |
| Constant | C | 0.15 | 0.145 | 0.8856 |
| AR(1) | Ø | 0.38** | 2.13 | 0.0432 |
| $R^2$ | 0.99899 | | | |
| Adj. $R^2$ | 0.9984 | | | |
| DW stat | 1.9 | | | |
| F-statistic | 6481.5 | | | |
| P-value | 0.000 | | | |
| Convergence achieved after 10 iterations | | | | |

*Note*. ** and * indicate coefficients are significant respectively at 5% and 1% level of significance.

## Johansen Cointegration Test and FMOLS Estimates

The results of the Johansen CI tests are reported in Table 5. The null hypothesis of no CI is rejected by both the trace and maximum eigenvalue tests at 5% level of significance. This implies that there exists one CI relationship among the variables.

**Table 5**

*Johansen Cointegration Test Results*

| Unrestricted Cointegration Rank test (Trace) | | | | |
|---|---|---|---|---|
| Null Hypothesis | Eigen Value | Trace statistics | 5 percent critical value | Prob.# |
| r = 0* | 0.716810 | 59.83082 | 47.85613 | 0.0025 |
| r ≤ 1 | 0.460146 | 22.08751 | 29.79707 | 0.2937 |
| r ≤ 2 | 0.088657 | 3.593806 | 15.49471 | 0.9335 |
| r ≤ 3 | 0.026598 | 0.808737 | 3.841466 | 0.3685 |
| **Unrestricted Cointegration Rank test (Maximum Eigenvalue)** | | | | |
| Null Hypothesis | Eigen Value | Max-Eigenvalue statistics | 5 percent critical value | Prob.# |
| r = 0* | 0.715810 | 37.74332 | 27.58434 | 0.0018 |
| r ≤ 1 | 0.460146 | 18.49370 | 21.13162 | 0.1125 |
| r ≤ 2 | 0.088657 | 2.785069 | 14.26460 | 0.9600 |
| r ≤ 3 | 0.026598 | 0.808737 | 3.841466 | 0.3685 |

*Note*. * indicates rejection of null hypothesis at 1% level of significance, while # indicates MacKinnon–Haug–Michelis (1999) p-value. r measures the number of CI relationships.

Equation 4 is estimated by FMOLS to obtain the optimal estimate of cointegrating regression. The FMOLS modifies the least square estimates to accommodate for serial auto-correlation in residual errors and for the endogenous relations among regressors resulting from variables being I(1) at level. This method was developed by Phillips and Hansen (1990) and modified by Phillips (1995).

The results of FMOLS estimates are reported in Table 6. They suggest that the coefficients of $\ln(K_p)$ and $\ln(K_g)$, that is, $\delta$ and $\alpha$, are positive and statistically significant while that of $\ln(L)$, that is, $\gamma$, remains

**Table 6**

*Estimates of Equation 4 by Fully Modified Ordinary Least Square*

| Variable | Coefficient | Coefficient value | t-statistic | Prob# |
|---|---|---|---|---|
| $\ln(L)$ | $\gamma$ | 0.056 | 0.671412 | 0.5077 |
| $\ln(K_p)$ | $\delta$ | 0.643* | 17.57980 | 0.0000 |
| $\ln(K_g)$ | $\alpha$ | 0.281* | 7.428225 | 0.0000 |
| Constant | C | 0.23 | 0.245430 | 0.8080 |
| $R^2$ | 0.9988 | | | |
| Adj. $R^2$ | 0.9986 | | | |
| DW stat | 1.22 | | | |
| Long-run variance | 0.000456 | | | |

*Note*. * indicates coefficients are significant at 1% level of significance.

statistically not significant.[5] A comparison of estimated coefficient values of parameters under OLS, FMOLS, and serial correlation–corrected estimates indicate that the output elasticities of inputs differ very marginally. The coefficient of $\ln(K_g)$ under FMOLS is close to that of $\ln(K_g)$ under OLS and serial correlation–corrected estimates. Similar are the cases of $\ln(K_p)$ and $\ln(L)$. Hence, the computed output elasticity of other inputs, namely, L and $K_p$, under FMOLS is 0.719 and is close to 0.715 under OLS estimation.

## ECM Results

Since there exists only one cointegrating vector as revealed by Johansen CI test, single-equation ECM representation is considered. The single-equation ECM representation can be expressed as

$$\Delta\ln(Y_t) = \Omega + \beta U_{t-1} + \pi_1\Delta\ln(K_{pt}) + \pi_2\Delta\ln(K_{gt}) + \pi_3\Delta\ln(L_t) + £_t \quad (6)$$

[5] It is not surprising that the coefficient of $\ln(L_t)$ expectedly has positive sign but is statistically not significant. This could be attributed to measurement problems in labour employment data in the unorganised sector, although the large share of Indian economy comes from this sector.

In Equation 6, $£_t$ is the error term which follows the *white noise* process, β is the speed of adjustment (long-run causality) parameter associated with the lagged value of the estimated error residuals of Equation 4, and $\pi_1$, $\pi_2$, and $\pi_3$ are the short-run (causality) parameters. Ω is the constant term. $U_{t-1}$ is the lagged 'disequilibrium term', and β is the error correction or adjustment coefficient. A meaningful ECM implies

**Table 7**

*Estimated Error Correction Model Results and Diagnostic Checks*

| Coefficients | Estimated coefficients | Standard error | t-statistic |
|---|---|---|---|
| Ω | 0.0085 | 0.0166 | 0.515 |
| β | -0.6026* | 0.213 | -2.822 |
| $\pi_1$ | 0.521* | 0.136 | 3.844 |
| $\pi_2$ | 0.283* | 0.1002 | 2.824 |
| $\pi_3$ | 0.104 | 0.1834 | 0.564 |
| $R^2$ | 0.43 | | |
| Adj. $R^2$ | 0.34 | | |
| F-statistic | 4.85 | | |
| Prob# | 0.005 | | |
| DW | 1.8 | | |
| AIC | -5.11 | | |
| **Diagnostic Checks** | | | |
| Breusch–Godfrey Serial Correlation Lagrange Multiplier Test | | | |
| Optimal Lag length is 1 after considering SIC and AIC statistics | | | |
| F-statistic | 1.433 | Prob. F(1,25) | 0.2425 |
| Obs*R-squared | 1.681 | Prob. Chi-Square(1) | 0.1948 |
| Heteroskedasticity Test: Breusch-Pagan-Godfrey | | | |
| F-statistic | 2.28 | Prob. F(4,26) | 0.088 |
| Obs*R-squared | 8.04 | Prob. Chi-Square(4) | 0.09 |
| Residual Normality Tests | | | |
| Jarque–Bera (J–B) Statistics | 1.09 | Prob | 0.58 |

that β should be negative and statistically significant. The estimated ECM of Equation 6 is presented in Table 7.

Estimated ECM results and diagnostic checks of Equation 6 reported in Table 7 reveal that the estimated β is negative and statistically significant. Its value (-0.6026) suggests that, on an average, disequilibrium error is adjusted by 60% annually. Except labour coefficient ($\pi_3$), short-run dynamic effects of private- and public-sector capital stocks ($\pi_1$ and $\pi_2$) have significant positive impact on output (lnY) in the short run. The diagnostic checks show that the ECM estimation does not suffer from the problem of serial correlation and heteroscedasticity in residual errors. Also the Jarque–Bera (J–B) statistics suggest that the residuals are normally distributed.

## Growth-Maximising Debt-to-GDP Ratio (d*)

Output elasticity of public capital (α)—under OLS, serial correlation–corrected estimates, and FMOLS—is estimated. As mentioned, the output elasticity of $K_g$ is crucial in the computation of growth-maximising public-debt-to-GDP ratio (d*). Higher the value of estimated α, higher would be the value of the d* and vice-versa. The computed values of d* under OLS, serial correlation–corrected estimates, and FMOLS, are presented in Table 8. It shows that the growth-maximising debt-to-GDP in India should hover around 65%–67%, which is modestly higher than the medium-term, sustainable debt-to-GDP ratio (60%) estimated by the

**Table 8**

*Computation of Growth-maximising Debt-to-GDP Ratio (d*) in Percentage*

| Models | Formula for calculating d* | Value of d* (%) |
|---|---|---|
| OLS Estimate of α = 0.285 | | 66 |
| Serial correlation corrected estimate of α = 0.29 | $d^* = \{\alpha / (1 - \alpha)^2\}^{1 - \alpha}$ | 67.5 |
| FMOLS Estimate of α = 0.281 | | 65 |

N. K. Singh Committee (Government of India, 2017) for central and state governments combined. However, the estimated d* value is substantially lower than the all-time-high debt-to-GDP ratio (89% in 2003–2004) and the recently revised estimate (73% in 2015–2016) (Government of India, 2010, 2018). Therefore, the need for fiscal tightening by both central and state governments is paramount not only to ensure macroeconomic stability but also to accelerate growth.

## Some Caveats

The results presented above are subject to some estimation and inferential limitations.

First, data-related problems may affect the estimated value of coefficients and may rule out getting an exact estimation. For example, India does not have reliable estimates of labour employment data over the years, especially because reliable data on employment is difficult to obtain due to large-scale unorganised and informal employment. Although the study has used labour hours worked annually, the number of labourers engaged annually is difficult to gauge in India. Moreover, significant productivity and wage differences exist not only between formal and informal sector but also within a particular sector across states (or regions). This may be one reason why the labour coefficient is statistically insignificant in all estimations.

Second, data on private- and public-sector capital stock in constant price with 2004–2005 base year is derived from the aggregate capital stock data provided by KLEMS India on the basis of share of gross capital formation by these two sectors, respectively. Besides, productivities of different types of capital stock are not the same. Even within the private sector, productivities of the household sector and the private corporate sector capital stock would be different. Unless all these data and measurement issues are addressed, the estimated results may not be an exhaustive estimate for d*.

Third, whatever estimates have been obtained are point estimates with their respective standard errors. Hence the value of the output elasticity of capital ($\alpha$) and growth-maximising d* would be indicative, not exhaustive.

Fourth, the aggregate one-sector production function with three inputs (L, $K_p$, and $K_g$) is used to estimate the output elasticity of these inputs under the assumption of constant return to scale, not only between public and private inputs but also between two private inputs (L and $K_p$).

Since a large chunk of the combined fiscal deficit (Revenue deficit was 48% of fiscal deficit in 2014–2015) is due to persistent revenue deficit, the effective value of d* would be even lower than 65%–67% range in the Indian context.

## Conclusion and Implications

The contributions of this research work are the following:

First, following the theoretical model and empirical estimation by Aschauer (2000) and Checherita-Westphal et al. (2012), the present paper has estimated an optimal value of debt-to-GDP ratio to maximise economic growth in the Indian context, under the assumption that government borrowing is dedicated to creation of capital stock and not for consumption or revenue expenditures (i.e., golden rule of budget deficit).

Second, based on the estimated value of the output elasticity of public sector capital stock (α), the value of growth-maximising debt-to-GDP ratio (d*) is computed to be around 65%–67%.

As the value of d* is substantially lower than the current value of debt-to-GDP ratio (73% in 2015–2016), the importance of fiscal tightening by central and state governments (combined) should be strengthened. In the Indian context, since a large chunk of fiscal deficit is on account of revenue deficit, the effective value of d* would be much lower. The central and state governments should not only target reducing the fiscal deficit, but more importantly reduce the revenue deficit more aggressively. Eliminating the revenue deficit or limiting the primary deficit is necessary not only to ensure macroeconomic stability but also to accelerate growth.

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# Appendix

**Table A1**

*Descriptive Statistics of Cointegrated Variables (1980–1981 to 2011–2012)*

| | LNY | LNK | LNKG | LNL |
|---|---|---|---|---|
| *M* | 16.83076 | 17.31114 | 16.74996 | 13.58246 |
| Median | 16.79944 | 17.31019 | 16.70329 | 13.58405 |
| Maximum | 17.84675 | 18.60019 | 17.38918 | 13.87879 |
| Minimum | 15.97462 | 16.25339 | 16.28687 | 13.13846 |
| *SD* | 0.556733 | 0.729874 | 0.287035 | 0.250507 |
| Skewness | 0.209906 | 0.113371 | 0.910365 | -0.291900 |
| Kurtosis | 1.904012 | 1.778890 | 3.249741 | 1.824702 |
| Sum | 538.5845 | 553.9566 | 535.9986 | 434.6388 |
| Sum Sq. Dev. | 9.608490 | 16.51420 | 2.554066 | 1.945374 |
| Observations | 32 | 32 | 32 | 32 |

## Table A2

*Stationarity Checks ADF and PP Test Statistics*

| | ADF tests statistics | | PP tests statistics | | | ADF tests statistics | | PP tests statistics | | Whether I(0) or I(1) in level |
|---|---|---|---|---|---|---|---|---|---|---|
| Variable | Intercept | Intercept & Trend | Intercept | Intercept & Trend | Variable | Intercept | Intercept & Trend | Intercept | Intercept & Trend | |
| $\ln(Y)$ | 2.55 | -0.44 | 5.46 | 0.22 | $\Delta\ln(Y)$ | -4.58* | -5.61* | -4.58 | -6.64 | I (1) |
| $\ln(K_g)$ | 1.05 | -0.48 | 1.24 | -0.49 | $\Delta\ln(K_g)$ | -5.44* | -5.87* | -5.5* | -5.87* | I (1) |
| $\ln(K_p)$ | 9.38 | 3.89 | 14.7 | 6.3 | $\Delta\ln(K_p)$ | 1.78 | -4.45* | -1.12 | -4.52* | I (1) |
| $\ln(L)$ | -1.1 | -0.79 | -0.99 | -1.45 | $\Delta\ln(L)$ | -3.87* | -3.99** | -3.92* | -3.97** | I (1) |
| $\ln(EL)$ | 0.69 | -4.32* | -0.81 | -1.96 | $\Delta\ln(EL)$ | -4.2* | -4.33* | -2.14 | -2.1 | I (1) |

*Note.* * and ** indicate the rejection of null hypothesis of unit roots of variables are rejected at 1% and 5% level of significance respectively.

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