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Investigating Short Run Causality between Real GDP and Government Expenditure in India Since 1950s

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Abstract
This paper has examined short run causality between government expenditure and GDP in India during 1951-2013 using a Toda-Yamamoto (1995) modified Granger causality approach under VAR environment. Exponentially detrended annual time series data on GDP and government expenditure at constant prices are used. Structural break point unit root tests are conducted besides the usual unit root tests to determine the order of integration of each variable. Tests for structural breaks reveal significant breaks in both time series around the period 2001-04. Government expenditure is found to significantly Granger-cause real GDP but the converse is insignificant implying that Wagner's law is inapplicable. The study thus suggests uni-directional causality from government expenditure to GDP. Moreover government expenditure in India has a long-run co-integrating relationship with real GDP and therefore short run causal relations may be anticipated.
Keywords: Real GDP, government expenditure, Granger Causality, Toda-Yamamoto approach and VAR.
JEL Classification: C32, E50, E52, E59.

1. Introduction and Objectives

The association between government expenditure and GDP has long been debated by macroeconomists both at theoretical and empirical levels. Two lines of thought have naturally dominated the debate. The first is the Keynesian approach which argues that it is public spending or government expenditure that influences GDP, leading to the all important policy proposition that public spending may be stepped up during phases of low GDP growth. The second is that of Wagner's law which states that it is GDP or national income that influences the amount of public spending. The law was formulated by Wagner on the basis of his empirical observation for several advanced capitalistic industrial nations where he found a long run tendency of government expenditure to increase as per capita income rises. In public economics this law also came to be known as the law of increasing state activities.

Wagner identified GDP growth as the key determinant of public expenditure growth. Keynes on the other hand maintained that government expenditure is the principal driver of GDP growth. Thus according to Keynes government expenditure acts as an instrument of fiscal policy and is effective in maintaining short-run macroeconomic stability as well as in promoting long-run macroeconomic growth. Hence the standard Keynesian prescription during times of poor growth is that of fiscal intervention and this plays a pivotal role in the path of macroeconomic recovery and growth. For developing nations such as India fiscal expansion is associated with both growth and development as social overhead spending promotes both. During macroeconomic recession the standard Keynesian prescription is to borrow funds from the private sector, non-bank public and the commercial banking sector (or sometimes the financial sector in a broad sense) via issue of government bonds (or its different variants) and ultimately to spend on the people through the government's numerous spending programmes in the form of fiscal expansion. Developing nations like India mostly emphasise on infrastructure development as a part of its fiscal expansionary programmes which directly leads to macroeconomic growth as well as development. Government expenditure is treated as an exogenous variable or a policy instrument aimed at short-run stabilization of the economy. Thus according to Keynesian approach macroeconomic causality between GDP and government expenditure should run from government expenditure to GDP and not the other way round.

The literature on empirical testing and validity of Wagner's law and the Keynesian hypothesis is quite extensive and hence a detailed discussion on empirical contributions in this field is beyond the scope of this paper. Although numerous empirical studies have been carried out on the topic over the years in both developed and the developing countries, the findings of these studies are rather mixed and no clear consensus on the direction of causality seems to emerge. A few influential works are cited as a build-up to the present study.

Econometric causality between government expenditure and national income was examined by Singh and Sahni (1984) on the basis of Indian data. Their results support neither Wagner's law nor the Keynesian hypothesis. Ahsan et al. (1992) did not find any causality between government expenditure and national income for the United States. No evidence of Wagner's law was found by Afxentiou and Serletis (1996) and Ansari et al. (1997) on the basis of cross-country data. Bohl (1996) tested Wagner's law for the post-World War II period for G7 countries and found it to be valid only for the UK and Canada. In case of Turkey, Bagdigen and Cetintas (2003) did not find any causality between national income and government expenditure. Neither Wagner's law nor its converse was established by Frimpong and Oteng-Abayie (2009) for the West African Monetary Zone nations. According to Verma and Arora (2010) there is no instantaneous impact of increase in government expenditure on GDP in case of India. Similarly, Taban (2010) found no significant relationship between government spending and GDP growth for Turkey.

At the other extreme, studies by Chletsos and Kollias (1997) for Greece, Ghali (1998) for the ten OECD nations, Demirbas (1999) for Turkey, Thornton (1999), besides Chang (2002) for six up-and-coming nations, Kolluri et al. (2000) for G7 nations, Islam (2001) for USA, Al-Faris (2002) for Gulf Cooperation Council countries, Aregbeyen (2006) for Nigeria, Sideris (2007) for Greece, Kalam and Aziz (2009) for Bangladesh and Rehman et al. (2010) for Pakistan found causality between national income and public expenditure. Grullón (2012) and Salih (2012) found Wagner's law to be valid in case of the Dominican Republic and Sudan, respectively. Contrary to these, the studies by Jiranyakul and Brahmasrene (2007) in case of Thailand, Pradhan (2007) in case India, Babatunde (2008) for Nigeria, Magazzino (2010) for Italy and Ighodaro and Oriakhi (2010) for Nigeria established the Keynesian result that public expenditure causes national income. Ayo et al. (2011) in an exceptional result established bi-directional causality between government expenditures and economic growth both in the short and the long run for Nigeria.

The paper is presented in the following sections. After a brief discussion of theoretical and empirical works on the topic in the introduction, data sources and econometric issues are discussed in section 2. Section 3 presents the analysis of empirical results, followed by summary and concluding remarks in section 4.

2. Methodology and Data

2.1 Data detrending

The study period is 1951 – 2013 as annual government expenditure figures for India are available from 1951 onwards. This gives us 60 annual time point observations which is a typical macroeconomic long run. The variables considered are government expenditure (G) and Gross Domestic Product at 2004-05 prices which is constant price GDP or synonymous to real GDP. The 'G' figures are also WPI deflated. The entire data set is compiled from Reserve Bank of India: *Handbook of Statistics on the Indian Economy*, 2014 available at the RBI website.

Long run macroeconomic data is most likely to have a trend – linear or non-linear. A glance at the time series line plots for each variable during 1951-2013 (not presented) reveals strong non-linear trends in all three variables. Both parabolic and exponential curves are fitted to each variable and the goodness of fit statistics are compared (see Table A1 of the Appendix). The results are strongly suggestive of exponential trends in each of the three variables. Accordingly, the exponentially detrended series on each variable are preferred for analysis. The detrended data is generated using the following steps. First, the natural logarithm of the variable is regressed linearly on a constant and time, i.e., the linear regression $ln(y_t) = ln(\alpha) + \beta.t + error,$ is run where y_t is the variable to be detrended. This is a log-linear form of the exponential growth (or smoothing) function $y_t = \alpha.\exp(\beta.t).$ Second, the parameters α and β are estimated using OLS and predicted $ln(y_t)$ series is generated. Third, anti-log of predicted $ln(y_t)$ is generated, which is predicted y_t in non-logarithmic form. Finally $e_t = y_t - \widehat{y_t}$ is the residual from the exponential smoothing (or curve fitting) in non-logarithmic form and is thus the part of y_t that is free from any exponential trend (where $\widehat{y_t}$ is predicted y_t in non-logarithmic form). Hence, e_t is exponentially detrended y_t. This method is applied to detrend both variables – government expenditure and GDP.

Standard tests for stationarity may be misleading for non-linearly trended data (for instance quadratic or exponential, both of which are rising at a rising rate over time) as because standard tests of stationarity such as Augmented Dickey-Fuller and Philips-Perron tests include linear trend terms only (i.e., some 'constant' times 'time'). For an exponentially growing variable, stationarity may not be attained even at second difference, although for de-trended series it may be attained either at level (if trend stationary) or at first difference. Moreover, the autocorrelation function (ACF) helps us to select the lag lengths p (order of AR) and q (order of MA) and the ACF of the residuals is an important diagnostic tool. Unfortunately ACF as used in linear models may be misleading for non-linear models. The reason is that autocorrelation coefficients measure the degree of linear association between Y_t and Y_{t-i} (Y is the time series variable in question). As such ACF may fail to detect important non-linear relationships in the data. It is thus desirable to work with detrended data.

2.2 Testing Stationarity in the Presence of Structural Breaks

In the long run macroeconomic variables are expected to experience structural breaks, some of which may be the result of macroeconomic policy shifts, regime changes, or random shocks (droughts, warfare, socio-political instability and violence, etc.) at the domestic level or due to similar factors at the international level. The present paper applies the Bai-Perron (1998 and 2003) multiple unknown structural break point test to original as well as the detrended series and compares the periods of break for each of the three variables. Instead of going into the mathematical details, the method of break date determination as performed using EVIEWS 9 is as follows. First the time series variable in question is regressed (using OLS) on a constant only allowing for serial correlation that varies across break dates (regimes) through the use of HAC covariance estimation. Three break dates are considered along with a trimming percentage of 20, which implies around 12 observations per regime (since the period 1954-2913 implies 60 observations). Since the errors are assumed to be serially correlated, quadratic spectral kernel based HAC covariance estimation is specified using prewhitened residuals. The kernel bandwidth is determined automatically using the Andrews AR(1) method. The default **method** setting in EVIEWS 9 **(sequential L+1 breaks vs. L)** instructs the software to perform sequential testing of *l+1* versus *l* breaks using the methods outlined in Bai (1997) and Bai and Perron (1998). The error distribution is allowed to differ across breaks to allow for heterogeneity. This test employs the same HAC covariance settings as used in the original equation but assumes regime specific error distributions. The break dates along with the respective F-statistic values are presented in the results empirical section. Stationarity related issues are discussed next. The Bai-Perron 'Global break point vs. none' test is not carried out in this study.

Perhaps the most widely used unit root test to examine the stationarity of a time series (order of its integration) is the Augmented Dickey-Fuller test (ADF test) which makes use of equation (2.3.1). This generalised form includes both trend and intercept in the model.

$$\Delta y_t = a_0 + \gamma . y_{t-1} + a_1 . t + \sum_{i=1}^{p} \beta_i . \Delta\, y_{t-i} + \varepsilon_t \qquad (2.3.1)$$

Equation (2.3.1) tests the null hypothesis of a unit root against a trend stationary alternative. The optimum number of lagged Δy_t terms (introduced to tackle serial correlations in the errors) may be determined by the optimum value of some information criterion such as Schwartz's Information Criterion (SIC). Phillips and Perron (1988) proposed a nonparametric method of controlling serial correlation while testing for unit root. They estimate the unaugmented Dickey-Fuller test equation [Equation (2.3.1) without the term ($\sum_{i=1}^{p} \beta_i . \Delta\, y_{t-i}$) on the right hand side], and modifies the t-ratio of the γ coefficient so that serial correlation does not affect the asymptotic distribution of the test statistic.

Kwiatkowski, Phillips, Schmidt and Shin (1992) propose a test of the null hypothesis that the observed series is stationary around a deterministic trend. The series is expressed as the sum of deterministic trend, random walk and stationary error and the test is the LM test of the null hypothesis that the random walk has zero variance. The asymptotic distribution of the statistic is derived under the null and under the alternative that the series is difference stationary. KPSS test is quite contrary to the ADF and PP tests which consider the null hypothesis of unit root (i.e. a non-stationary series) as opposed to the former (KPSS) which considers a null hypothesis of stationary series.

The ADF and other traditional stationarity tests do not normally include a structural break term. But one can insert structural break dummies (say, seasonal dummies, for example) in equation (2.3.1) that may include both slope and intercept dummies. The point of break may be exogenously determined (approximately) by a visual scrutiny of the time series line plots. Importantly, the ADF test fails to perform well in the presence of structural breaks especially when the breaks are ignored. In such situations unit root tests with structural breaks are more suitable [see Perron (1989); Zivot and Andrews (1992)]. Perron (1989) demonstrated, assuming an exogenously fixed break date, that the power to reject the null hypothesis of unit root decreases (given that the alternative hypothesis of stationarity is actually true) when the structural break is ignored.

Zivot and Andrews (1992) suggest an improvement over the Perron (1989) test where they presume that the exact break point is unknown and endogenise the break date determination. A data dependent algorithm is used to proxy Perron's subjective procedure to determine the break points endogenously. Following Perron's characterization of the form of structural break, they adopt the following three models to test for unit roots.

$$\Delta y_t = a_0 + \gamma . y_{t-1} + a_1 . t + \delta . DU_t + \sum_{i=1}^{p} \beta_i . \Delta\, y_{t-i} + \varepsilon_t \qquad \text{(Model A)}$$

$$\Delta y_t = a_0 + \gamma . y_{t-1} + a_1 . t + \theta . DT_t + \sum_{i=1}^{p} \beta_i . \Delta\, y_{t-i} + \varepsilon_t \qquad \text{(Model B)}$$

$$\Delta y_t = a_0 + \gamma . y_{t-1} + a_1 . t + \theta . DU_t + \delta . DT_t + \sum_{i=1}^{p} \beta_i . \Delta\, y_{t-i} + \varepsilon_t \quad \text{(Model C)}$$

Here DU_t captures mean shift occurring at each possible break-date (TB) while DT_t is corresponding trend shift variable. Formally the values assigned to DU_t and DT_t may be summarised as follows. $DU_t = 1\ for\ t > TB\ ,\ and\ = 0\ otherwise$. On the other hand $DT_t = t - TB\ for\ t > TB\ ,\ and\ = 0\ otherwise$.

The null hypothesis in all three models is that $\gamma = 0$, which implies that $\{y_t\}$ has a unit root with drift without any structural break. The alternative hypothesis if $\gamma < 0$, implies that the series is a trend-stationary with a single break occurring at some unknown time point. Zivot and Andrews regard every point as a potential break-date (*TB*) and run a regression for every possible break-date sequentially. From all possible break-points (*TB*), the procedure selects as its choice of break-date (*TB*) the date which minimizes the one-sided t-statistic for testing $\gamma = 0$ against $\gamma < 0$ [or $\gamma = (\varphi - 1) < 0$]. According to Zivot and Andrews, the presence of the end points cause the asymptotic distribution of the statistics to diverges towards infinity. Therefore, some region must be chosen such that the end points of the sample are not included. More recently, Sen (2003) showed that if one uses model A and if the break occurs according to model C then there would be a sizeable loss in power of the test. However, if break is characterized according to model A, but model C is used then the loss in power is negligible, suggesting the superiority of model C over model A. While Zivot and Andrews (1992) and Perron (1997) determined the point of break 'endogenously' from the data, Lumsdaine and Papell (1997) suggested an improvement over the Zivot and Andrews (1992) model by incorporating a couple of structural breaks. However, such endogenous tests have been subject to criticism for their treatment of breaks under the null hypothesis. If the breaks are absent under the null hypothesis of unit root these tests may suggest evidence of stationarity with breaks (Lee and Strazicich, 2003). Lee and Strazicich (2003) on the other hand propose a two break minimum Lagrange Multiplier (LM) unit root test in which the alternative hypothesis unambiguously implies that the series is trend stationary.

2.3 Toda – Yamamoto Modified Granger Causality under VAR Environment

A simple definition of Granger Causality, in the case of two time-series variables, *X* and *Y* is as follows. "*X* is said to Granger-cause *Y* if *Y* can be better predicted using the histories of both *X* and *Y* than it can by using the history of *Y* alone." The absence of Granger causality can be tested by estimating the following VAR model (equations 2.4.1 and 2.4.2).

$$y_t = \alpha + \sum_{i=1}^{p} \alpha_i y_{t-i} + \sum_{i=1}^{p} \beta_i x_{t-i} + u_{1t} \qquad (2.4.1)$$

$$x_t = \beta + \sum_{i=1}^{p} \lambda_i y_{t-i} + \sum_{i=1}^{p} \delta_i x_{t-i} + u_{2t} \qquad (2.4.2)$$

For the present study y_t represents detrended real GDP for India and x_t represents government expenditure or G. *X* does not Granger cause *Y* is tested by H_{01}: $\beta_1 = \beta_2 = \cdots = \beta_p = 0$ against the alternative that $\beta_1 \neq \beta_2 \neq \cdots \neq \beta_p \neq 0$. On the other hand *Y* does not Granger cause *X* is tested by H_{02}: $\lambda_1 = \lambda_2 = \cdots = \lambda_p = 0$ against the alternative the $\lambda_1 \neq \lambda_2 \neq \cdots \neq \lambda_p \neq 0$. In each case rejection of null hypothesis implies the presence of Granger causality. The modified Wald test for testing Granger causality as proposed by Toda and Yamamoto (1995) avoids the problems associated with the usual Granger causality testing (which ignores non-stationarity and cointegrations between series while testing for causality). If the Wald test is being used to test linear restrictions on the parameters of a VAR model, and the data are non-stationary (which is most likely), then the Wald test statistic *does not follow its usual asymptotic chi-square distribution under the null hypothesis (Toda and Yamamoto, 1995).*

The approach to modified Granger causality as adopted in this paper is outlined as follows. ***First****, each time series variable is tested for stationarity (or for its order of integration) using standard tests such as ADF, PP and KPSS.* The maximum order of integration (*m*) for the group of time-series is determined. Structural breaks if any are identified and a structural break dummy variable is created. **Second**, a VAR model is set up in level, regardless of the orders of integration of the various time-series. None of the variables are differenced.

Third, the optimum lag length for each variable in the VAR, say *p*, is determined using AIC, SIC, HQ, or other usual statistics. Care is taken so that there is no serial correlation in the residuals. The length *p* may be increased slightly until autocorrelation issues are resolved. Normality of the VAR residuals is highly desirable. **Fourth**, if both the time-series have the same order of integration, then Johansen Co-integration test is applied to test for co-integration (based on the selected VAR model). It provides some cross-check on the validity of the Causality results. **Fifth**, the favoured VAR model is constructed and additional *m* lags of each variable are inserted into each equation. In EVIEWS 9 these new *m* variables are to be treated as exogenous to the VAR system. The structural break dummy is also added (not shown) as an exogenous variable. It is thus ensured that the additional *m* lags and the structural break dummy would not be dropped while testing for Granger non-causality (via the Wald tests). The new VAR is presented in equations 2.4.1(a) and 2.4.1 (b).

$$y_t = \alpha + \sum_{i=1}^{p} \alpha_i y_{t-i} + \sum_{j=p+1}^{p+m} \alpha_j y_{t-j} + \sum_{i=1}^{p} \beta_i x_{t-i} + \sum_{j=p+1}^{p+m} \beta_j x_{t-j} + u_{1t} \qquad [2.4.1(a)]$$

$$x_t = \beta + \sum_{i=1}^{p} \alpha_i y_{t-i} + \sum_{j=p+1}^{p+m} \alpha_j y_{t-j} + \sum_{i=1}^{p} \beta_i x_{t-i} + \sum_{j=p+1}^{p+m} \beta_j x_{t-j} + u_{2t} \qquad [2.4.2\ (b)]$$

Finally, the hypothesis that the coefficients of only the first *p* lagged values of *x* are restricted to zero in the first equation (i.e. 2.4.1(a)), is tested using the standard Wald test (to test H_{01}: *x* does not Granger cause *y*). Analogously, a similar procedure is followed (for equation 2.4.2(b)) to test that *y* does not Granger cause *x*. The Wald statistic under the null hypothesis will be asymptotically distributed as chi-square with *p* degrees of freedom. Importantly enough, if two or more time-series are cointegrated, then there must be Granger causality between them (either uni-directional or both ways). The converse however is not true. The next section presents empirical results of the study along with necessary discussions.

3. Empirical Results and Analysis

This section presents the empirical results of the present study. For the purpose of choosing the appropriate detrended series for each time series variable both exponential and parabolic curves are first fitted to the data and the goodness of fit statistics of both are presented in the appendix (see table A1). The EVIEWS reported values of R-square, adjusted R-square, AIC, SIC and HQ are presented for both models. It is evident that the exponential trend fit is a statistically better compared to the parabolic fit for each variable – real GDP and government expenditure. This is by virtue of obtaining higher R-square and lower AIC, SIC and HQ values in case of exponential trend fit. Thus the results in table A1 justify exponential detrending rather than quadratic or parabolic detrending. Hence the present study makes use of exponentially detrended data on each of the three variables.

Identification of structural breaks is of utmost importance. Table 1 presents the results of Bai-Perron test for unknown multiple structural break points of original *vis-a-vis* de-trended annual time series of selected variables. Interestingly, the detrended series exhibit single break points only.

Table 1. Bai-Perron Test for Unknown Multiple Structural Break Points of Original *vis-a-vis* De-trended Annual Time Series

Variables	Break dates in Original Series	Break Dates in De-trended Series
GDP	**1985, 1993, 2004**	**1999**
F-Statistic	35.22, 49.97, 39.55	46.28
G	**1996, 2005**	**1998**
F-Statistic	86.09, 29.88	66.53

Source: Computed on the basis of original and exponentially detrended time series data for major macroeconomic indicators of India (1951-2013) taken from *RBI: Handbook of Statistics on the Indian Economy,* 2014. **Notes**: F-statistic values corresponding to each repatriation are presented below the break date series.

For detrended variables the break points are either in 1998 or in 1999. In other words there is a consistency in the time series behaviour of the detrended series of both real GDP and government expenditure. The original or non-detrended series on the other hand exhibits different break dates. Real GDP exhibits significant breaks in 1985, 1993 and 2004. Interestingly no breaks in the original real GDP series are observed during the plan holidays of the 1960s or just after nationalisation of banks. The first statistically significant break is found to occur at 1985, the first year of the period of weak liberalisation in India. The second break date in the original GDP series is 1993, two years after the first wave of major economic reforms of 1991. Finally the third break date in real GDP is found at 2004. Government expenditure exhibits two points of break, one at around 1996 and the other at 2005.

Table 2. Structural Break Point Unit Root Test of De-trended Series

Variables	**ADF**		**Zivot-Andrews**	
	Level	1st Diff.	Level	1st Diff.
GDP	-4.02 (0.144, 4)	-7.98 (<0.01, 4)	-3.29 (<0.84, 4)	-6.53 (<0.01, 4)
Break Date	1999	2001	2000	2001
G	-9.67 (<0.01,5)	NA	-9.32 (<0.01,4)	NA
Break Date	1999	NA	1999	NA

Source: Estimated on the basis of secondary time series data on relevant variables (RBI: Handbook of Statistics on the Indian Economy, 2014) using EVIEWS 9 for Windows.

Notes: (i) Figures free of parenthesis in each cell are computed test statistic values. The first figure in parenthesis indicates p-value. For very small p-values (0.001, etc, exact p-values are not presented, instead <0.01 is used. (ii) The second figures in parenthesis indicate optimum lag length as selected by Schwartz's Criterion (automatic selection by the EVIEWS 9). (iii) A single unknown break date is selected by minimising the Dickey-Fuller t-Statistic automatically set in EVIEWS 9.

Stationarity testing is important from the point of view of knowing the order of integration of each time series variable. For example, if a time series is stationary not at level but at first difference then it follows an *I(1)* process. If a time series has a structural break the usual unit root test results (without incorporation of a break dummy) would be not only different, but would be misleading. Structural break point unit root tests are most appropriate under such circumstances. The structural break point unit root test results for all detrended variables are shown in table 2. GDP is found to be non-stationary at level but stationary at first difference according to both the ADF and Zivot-Andrews tests and the break dates are identical (at 2001), with software determined optimum lag length at 4. EVIEWS 9 automatic optimum lag length selection option on the basis of Schwartz's Information

Criterion was chosen. However, the detrended government expenditure is found to be stationary at level according to both tests. Both tests suggest a structural break date of 1999 for government expenditure.

Table 3. Stationarity Tests of Original Time Series (non-detrended) Ignoring Structural Breaks in the Series

Variable	ADF			PP			KPSS		
	Level	1st Dif.	2nd Dif.	Level	1st Dif.	2nd Dif.	Level	1st Dif.	2nd Dif.
G	6.911 (0.999,6)	1.424 (0.999,6)	-4.268 (<0.01,7)	10.634 (0.999)	-0.902 (0.799)	-18.513 (<0.01)	0.263	0.278	0.223
GDP	16.202 (0.999,5)	0.639 (0.990,5)	-7.01 (<0.01,8)	17.288 (0.999)	-1.284 (0.632)	-16.392 (<0.01)	0.350	0.371	0.268

Source: Estimated on the basis of secondary time series data on relevant variables taken from RBI: Handbook of Statistics on the Indian Economy, 2014.

Notes: (i) First figures in each cell are computed test statistic values. The first figures in parenthesis in each cell indicate p-value. For very small p-values (smaller than 0.001, exact p-values are not presented, instead <0.01 is used. (ii) The second figures in parenthesis indicate optimum lag length as selected by Schwartz's Information Criteria (automatic selection by the EVIEWS-9). (iii) Asymptotic critical values of KPSS Test Statistic with trend and intercept: 1% = 0.216; 5% = 0.146; 10%=0.119. **'**Null hypothesis' for KPSS test is that the time series variable is stationary (or does not have unit root).

The picture however is very different in table 3 which presents the stationarity test results of original time series (non-detrended) ignoring structural breaks in each series. The ADF, PP and the KPSS test results are presented at level, first difference and second difference for each variable. In sharp contrast to the results in table 2, none of the time series variables are stationary at level or at first difference. The KPSS test shows no stationarity at level, first difference, or second difference in case of all three variables. Thus contrasting outcomes observed in tables 2 and 3 justifies 'detrending the long run time series data' on the one hand and 'incorporation of structural breaks while testing for unit root' on the other.

After testing the structural break points and stationarity (i.e. unit roots), the vector auto regression (VAR) between real GDP and government expenditure and consequently the modified Granger–Causality results are presented and discussed. But first the optimum lag length for the VAR (i.e., the number of lagged regressors to be incorporated in the VAR – both GDP and government expenditure terms) needs to be determined. The EVIEWS 9 reported optimum lag length selection criteria results are presented in table A2 of the appendix. Most criteria suggest that 4 endogenous lags must be chosen in the VAR system. According to Toda-Yamamoto (1995) however (m+d) lags have to be incorporated in the VAR model where d is the order of integration of each variable. The d additional lagged terms cannot be restricted to zero while testing for Granger –Causality.

The estimated results of the VAR between real GDP and government expenditure are presented in table A3 of the appendix. The terms year and period are synonymous here. When GDP is the dependent variable, the 1 year lagged GDP significantly explains current year GDP. Rest of the lagged GDP coefficients are insignificant. More importantly, 1, 2 and 4 years lagged government expenditure terms are statistically significant in explaining current year GDP. When government expenditure is the dependent variable only the 1 year lagged GDP term is significant. R-square and adjusted R-square are both close to 99 percent implying that the VAR in table A3 is in fact very well fit. Since both G and GDP are integrated of order 1, additional 5^{th} period lagged terms of both variables are introduced in the VAR as exogenous variables as per Toda-Yamamoto requirement. Apart from the intercept or constant, a structural break dummy variable is also included, the break date being taken as 1999 (the structural break dummy D_1999 assumes score 0 for pre 1999 observations and assumes score 1 for observations pertaining to 1999 onwards). The constant is insignificant in both models but the structural break dummy is statistically significant.

Before conducting Wald test for Granger Causality the statistical robustness of the VAR must be ensured. First, serial correlation if any must be eliminated from the VAR residuals. That is, VAR residuals must not be serially correlated and to this end the number of lagged endogenous regressors may have to be adjusted. Second, it is desirable that the VAR residuals be normal. Statistical testing and estimation based on non-normal disturbances may be problematic. The residual serial correlation LM tests for the GDP-G VAR were conducted in EVIEWS and the results are presented in table 4. The LM statistic is significant at 5.56 percent (so insignificant at 5 percent) only for the 6^{th} period lagged residual and the rest are statistically insignificant. The results of White's heteroscedasticity tests (not presented in tabulated form) reveal that the VAR residuals are jointly heteroscedastic at 6.2 percent level of significance as the computed Chi-square value of 81.09 for 63 degrees of freedom has a p-value of 0.062. Thus the homoscedasticity hypothesis may be accepted at 5 percent level but not at 10 percent level. The normality test results for the VAR residuals are shown in table 5. The joint hypothesis of zero skewness is accepted. Similarly the joint hypothesis of a kurtosis of 3 is also accepted. Finally the p-value corresponding to the Jarque-Bera test statistic is high implying that the joint null hypothesis of normality of residuals is accepted.

Table 4. The Residual Serial Correlation LM Tests
For the GDP – Government Expenditure VAR

Lags	LM-Stat	Prob
1	8.428	0.0774
2	8.392	0.0605
3	2.424	0.6257
4	1.395	0.8265
5	7.988	0.0722
6	6.343	0.1457
7	4.121	0.3508
8	7.523	0.0883
9	7.418	0.0924
10	8.275	0.0637

Source: Estimated from secondary data compiled from RBI: Handbook of Statistics on the Indian Economy, 2014. The results as generated under the post VAR option of Residual Tests in EVIEWS 9 are exactly presented without rounding-off.

The Wald tests for Granger Non-Causality, tests for zero parameter restrictions on the coefficients of the lagged endogenous variables of the VAR model. However the exogenous variables are not dropped. The Wald test results of Granger non-causality between real GDP and G are presented in table 6. The first null hypothesis that G does not Granger-cause real GDP is rejected at less than 0.1 percent. Thus the alternative that G causes GDP is accepted. The second null hypothesis that GDP does not Granger-cause G is accepted at 17.11 percent. Thus Wagner's Law is found to be invalid for the Indian economy during the study period. Hence real GDP Granger causes G but the converse is not true. In other words there is uni-directional causality between real GDP and government expenditure and runs from G to GDP. So fiscal expansion in India is found to have a positive influence on real GDP, but whether this expansionary fiscal policy is independent or triggered due to monetary factors is beyond the scope of the present paper.

Table 5. The Real GDP–Government Expenditure VAR Model: Normality Test of Residuals
Method of Orthogonalization: Cholesky (Lutkepohl)
Null Hypothesis: Residuals Are Multivariate Normal

Component	Skewness	Chi-sq	df	P-value
1	-0.338781	1.091123	1	0.2962
2	0.420140	1.652338	1	0.1986
Joint		2.743461	2	0.2537
Component	**Kurtosis**	**Chi-sq**	**df**	**P-value**
1	2.108114	1.988652	1	0.1585
2	1.964485	2.680702	1	0.1212
Joint		4.669354	2	0.0968
Component	**Jarque-Bera Test Statistic**		**df**	**P-value**
1	3.136601		2	0.2084
2	4.443382		2	0.1084
Joint	7.579983		4	0.1082

Source: Estimated from secondary data compiled from RBI: Handbook of Statistics on the Indian Economy, 2014. **Notes:** The results are EVIEWS 9 generated under the post VAR option of Residual Tests. The figures as reported in EVIEWS output sheet are exactly reproduced without rounding off.

Table 6. Wald Tests for Granger Causality between GDP and G (Included observations: 55)

Null Hypothesis	Chi-sq	df	P-value	Inference
(i) G does not Granger Cause GDP (absence of Keynesian mechanism)	28.53	5	<0.001	Reject Null Hypothesis
(ii) GDP does not Granger Cause G (absence of Wagner's Law)	7.74	5	0.1711	Accept Null Hypothesis

Source: Estimated from secondary data compiled from RBI: Handbook of Statistics on the Indian Economy, 2014. **Notes:** The results are EVIEWS 9 generated under the post VAR option of Lag Structure. G represents government expenditure. The 2nd null hypothesis implies absence of Wagner's Law.

The Johansen Co-integration test between GDP and government expenditure are presented in table 7. Clearly the trace test and maximum eigen value test indicates 1 co-integrating vector each between real GDP and G implying thereby that there is a long run equilibrium relationship between real GDP and government expenditure in India over the period 1951-2013. The long – run co-integrating or equilibrium relationship justifies the causality results obtained earlier.

Table 7. Johansen Co-integration Test between GDP and G

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	P-value **
None *	0.622082	56.71246	17.99296	0
At most 1	0.006361	0.363783	3.756954	0.5503

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	P-value **
None *	0.622082	56.34867	16.77044	0
At most 1	0.006361	0.363783	3.756954	0.5503

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

Source: Estimated from secondary data compiled from RBI: Handbook of Statistics on the Indian Economy, 2014. **Notes:** The results are EVIEWS 9 generated and are not rounded off. G implies government expenditure.

4. Summary and Conclusions

The present study has tested for short run causality between government expenditure and real GDP in India during 1951-2014 adopting the Toda-Yamamoto (1995) modified Granger causality approach under a VAR setup. Exponentially detrended annual time series data on constant price GDP and government expenditure are used for this purpose. Bai-Perron tests for structural breaks of the detrended data series reveal significant breaks in the variables around the period 1999-2001. The findings are suggestive of a uni-directional causality from government expenditure to GDP which supports the Keynesian prescription and Wagner's law is found to be invalid. Further both real GDP and government expenditure have a long-run co-integrating relationship. Hence short run causal relations may be expected. But to a certain extent fiscal expansion in India may not be economically independent of monetary expansion. To validate the results of the present study and refute Wagner's Law, both VECM and ARDL approaches need to be separately undertaken. Furthermore it has to be investigated whether fiscal deficits trigger broad money supply over the same period in India, but the issue of fiscal stimulus – money supply inter-linkage is beyond the scope of the present paper.

Appendix

Selected Tables

Table A1. Comparing Goodness of Fit Statistics of Parabolic Trend Fitting vis-s-vis Exponential Trend Fitting for each Time Series Variable for the period 1951-2013

Variables	Parabolic Trend Fitting		Exponential Trend Fitting	
	R^2;Adj.R^2	AIC;SIC;HQ	R^2;Adj.R^2	AIC;SIC;HQ
GDP	0.961; 0.956	28.19; 28.28; 28.23	0.983; 0.974	-1.35; -1.28; -1.32
G	0.949; 0.939	34.21; 34.29; 34.24	0.989; 0.981	-2.18; -2.09; -2.13

Source: Computed on the basis of secondary time series data compiled from *RBI: Handbook of Statistics on the Indian Economy,* 2014.

Notes: Parabolic trend is fitted by estimating the model $y_t = \beta_0 + \beta_1.t + \beta_2.t^2$.

Table A2. Optimum Lag Length Selection in VAR for the GDP and G Model

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-1445.74	NA	5.221E+20	53.6951	53.8434	53.7523
1	-1323.23	226.8864	6.660E+18	49.3065	49.6029	49.4208
2	-1296.25	47.9665	2.867E+18	48.4562	48.9008	48.6277
3	-1287.67	14.6050	2.435E+18	48.2877	48.8806	48.5164
4	-1278.42	12.4659*	1.750E+18*	47.9464*	48.8353*	48.3800
5	-1270.41	15.0736	2.012E+18	48.0942	48.8357	48.2894*

Source: Estimated on the basis of Secondary Data compiled from RBI: Handbook of Statistics on the Indian Economy, 2014. Results are EVIEWS 9 generated.

Notes: * 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

Table A3. VAR Model Estimates between GDP and Government Expenditure (detrended) for India during 1951-2014

Endogenous Variables	Dependent Variables	
	GDP	G
GDP(-1)	1.103349	0.329476
	(0.169121)	(0.101528)
	[6.52420]	[3.24521]
GDP(-2)	-0.359945	0.009899
	(0.209091)	(0.125524)
	[-1.72148]	[0.07886]
GDP(-3)	0.293079	0.076447
	(0.213756)	(0.128322)
	[1.37109]	[0.59573]
GDP(-4)	-0.019958	-0.171119
	(0.235844)	(0.141588)
	[-0.08462]	[-1.20858]
G(-1)	0.499152	1.106018
	(0.202977)	(0.178831)
	[2.45916]	[6.18474]
G(-2)	1.489192	-0.028759
	(0.407112)	(0.244405)
	[3.65794]	[-0.11767]
G(-3)	-0.413816	-0.295683
	(0.441518)	(0.265055)
	[-0.93727]	[-1.11554]
G(-4)	1.193324	-0.160814
	(0.452657)	(0.271748)
	[2.63625]	[-0.59177]
Exogenous Variables		
C	-5937.78	8873.661
	(10091.5)	(6058.3)
	[-0.58840]	[1.46471]
D_1999	77121.97	24779.46
	(23960.3)	(13385.3)
	[3.21873]	[1.85132]
GDP(-5)	-0.04639	-0.18993
	(0.20952)	(0.12578)
	[-0.22143]	[-1.50995]
G(-5)	-0.84891	0.033951
	(0.36927)	(0.22169)
	[-2.29890]	[0.15315]
R-squared	0.987677	0.993068
Adj. R-squared	0.984780	0.991550
F-statistic	340.6348	653.7953
Log likelihood	-662.2313	-634.1949
Akaike AIC	24.51707	23.49756
Schwarz SC	24.95459	23.93508

Source: Estimated on the basis of Secondary data on relevant variables compiled from RBI: Handbook of Statistics of the Indian Economy, 2014. Estimations are done using EVIEWS 9 for Windows.
Notes: The figures as reported in EVIEWS output sheet are exactly reproduced in table 6 without rounding off. The structural break dummy (for 2004) and the two 5th period lagged terms of both variables

are exogenous to the VAR system. The Granger Causality or Block Exogeneity Wald Tests imply zero parameter restrictions only the endogenous lagged terms i.e., on lag 1 to 4 only.

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