MULTICOINTEGRATION AND FISCAL SUSTAINABILITY IN INDIA: EVIDENCE FROM STANDARD AND REGIME SHIFTS MODELS*

1. Introduction

The issue of fiscal sustainability attracted considerable attention in the empirical literature in recent years, mainly as a result of severe fiscal imbalances in industrial and in emerging market economies.

Fiscal policy is defined as sustainable when the government intertemporal budget constraint is satisfied: namely when the current market value of debt equals the discounted sum of expected future surpluses. The fulfillment of the intertemporal budget constraint relies on a transversality condition ("no Ponzi games" condition), according to which the expected value of public debt must converge to zero as time goes to infinity. Solvency, in other words, implies that the government cannot leave a debt with positive expected value asymptotically.

Standard empirical analyses assume that the discount rate is positive and constant, and derive some testable hypotheses consistent with the government intertemporal budget constraint. Applied work in this area may be traced back to two parallel strands of literature. The former concentrates on the univariate properties of public debt (see, among others, Hamilton and Flavin, 1986; Wilcox, 1989 and Davig, 2005). The latter explores the cointegration properties of revenues and expenditures (see, among others, Trehan and Walsh, 1988; Hakkio and Rush, 1991; Quintos, 1995 and Martin, 2000).

Focusing on emerging market economies, the case of India deserves undoubtedly a particular attention because, notwithstanding a high rate of real growth during the last three decades, the fiscal

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situation exhibits clear signs of distress since the second half of the 90's.

Table 1 compares the fiscal outlook of BRIC countries (Brazil, Russia, India, and China) from 1993 to the first half of the last decade.

TABLE 1 - *Fiscal Outlook in BRIC Countries*

	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
BRAZIL													
Fiscal Balance	59.6	24.7	6.6	5.3	5.5	7.0	5.3	3.4	3.3	4.2	4.7	2.4	3.0
Public Debt	32.6	30.0	27.9	30.7	31.8	38.9	44.5	45.5	48.4	50.5	52.4	46.7	46.5
RUSSIA													
Fiscal Balance	-	-	-4.9	-7.4	-6.4	-4.8	-1.2	2.4	3.1	1.7	2.4	4.9	7.5
Public Debt	65.1	45.7	40.7	32.8	55.0	79.4	88.8	56.8	42.9	36.6	31.9	24.5	16.6
INDIA													
Fiscal Balance	-7.0	-5.6	-5.0	-4.9	-4.9	-5.3	-5.4	-5.2	-4.7	-5.9	-4.5	-4.0	-4.1
Public Debt	83.9	77.9	71.4	67.8	68.3	69.5	69.9	71.1	82.7	85.5	82.1	86.0	83.1
CHINA													
Fiscal Balance	-1.9	-1.9	-1.4	-1.2	-1.1	-1.5	-2.2	-2.8	-4.0	-2.6	-2.1	-1.3	-1.2
Public Debt	-	-	-	-	-	-	-	-	17.7	18.9	19.2	18.5	17.9

Source: De Paula (2008), Tables 4.1 - 4.4.
Notes: Fiscal Balance and Public Debt are expressed as a percentage to GDP.

A careful inspection of this table reveals that the overall fiscal position of India is significantly worse than all other BRIC countries. In terms of fiscal balance, India exhibits a persistent deficit on relatively high levels, whereas other countries display a permanent surplus (Brazil) or an impressive reversal from deficit to surplus (Russia). A similar pattern emerges focusing on the evolution of the debt/GDP ratio, where the values for India are by far the largest in this group without any significant tendency to debt reduction. This compares unfavorably with the strikingly low debt/GDP levels recorded by China, or with the strong debt reduction observed for Russia since the beginning of the last decade.

In a recent paper, I have reassessed the sustainability of the fiscal process in India applying standard cointegration techniques (Tronzano, 2013). The main result is that Indian fiscal policy is "weakly" sustainable according to Quintos (1995) original definition. This means that, although the intertemporal budget constraint is formally satisfied, the slow convergence speed of the bubble term may create problems for the government's ability to market its debt in the long run.

The econometric approach implemented in Tronzano (2013) suffers however from some limitations. As emphasized in Bohn (1995), the assumption of a positive and constant discount rate is not adequate in a stochastic economy. In this framework, uncertainty implies that the discount rate becomes subjective and time variant, as risk-averse agents attach a higher value to sure claims on future consumption. The transversality condition ensuring that the government will not engage in "Ponzi schemes" (issuing new debt to pay interest on the outstanding debt) will therefore incorporate a discount factor depending on the probability distribution of future debt and the marginal rate of substitution between present and future consumption.

The above considerations have far-reaching implications for applied research. In a stochastic environment, a more encompassing set of econometric criteria is needed in order to verify whether a given fiscal process is truly sustainable. More specifically, besides the standard cointegration relationship between revenues and expenditures, it is necessary to test for the existence of an additional equilibrium link between government debt and revenues (or expenditures). If these equilibrium conditions jointly hold, revenues and expenditures are multicointegrated, according to the definition originally proposed in Granger and Lee (1989). Multicointegration implies a deeper equilibrium relationship in a system of integrated variables, since these variables are tied together by two equilibrating forces rather than by the single equilibrium relationship characterizing standard cointegrated systems.

This paper extends my previous work on fiscal policy in India assuming a more realistic framework, namely a stochastic environment where the discount rate is time variant. In line with the above discussion, I rely on a multicointegration approach in order to assess whether the fiscal process is truly sustainable.

The outline of the paper is as follows. Section 2 spells out the concept of multicointegration, presents the original I(1) series, and outlines the evolution of their cumulative sums over time. The same section discusses some preliminary results obtained from the seminal approach of Granger and Lee (1989). Section 3 extends the analysis using more powerful econometric techniques. I first present some estimates obtained through the standard "single-step" approach pioneered in Haldrup (1994) and Engsted *et al*. (1997). This evidence is then complemented with results derived from several types of multicointegration models allowing for regime shifts. Section 4 concludes, summarizing the main results and drawing some general policy implications.

2. Econometric framework, data and preliminary results

2.1 Multicointegration and Fiscal Sustainability

Consider two economic time series y_t, x_t, and assume that they are non-stationary and integrated of order one [I(1)]. Linear combinations of y_t and x_t will generally be I(1). However, cointegration between y_t and x_t implies the existence of a stationary linear combination between them, that is: $z_t = y_t - bx_t$, where b is the cointegrating parameter and $z_t \sim I(0)$ is the stationary equilibrium error.

Consider now the cumulated error series $s_t = \sum_{j=0}^{p} z_{t-j}$ which, by construction, will be I(1). If s_t and y_t are cointegrated, the bivariate system of I(1) variables (y_t , x_t) is said to be multicointegrated[1]. Since s_t depends on y_t, x_t, and their lags, multicointegration implies the existence of long-run relationships at two different levels between two series. Multicointegration typically arises in the context of an intertemporal optimization problem (see Granger and Lee, 1990 and Engsted and Haldrup, 1999, section 2). In this case, both the levels and the rate of change of the variables entering the system are determinants of the optimal policy response[2].

This formal definition of multicointegration can be easily related to the issue of fiscal sustainability.

Assume that the bivariate system of I(1) variables (y_t, x_t) corresponds, respectively, to government tax revenues (y_t) and government total expenditures inclusive of interest payments on the stock of outstanding debt (x_t). The stationary linear combination between these variables defined above ($z_t = y_t - bx_t$) corresponds, in this case, to the cointegration relationship between revenues and expenditures analyzed in the empirical literature[3]. The short-run deviation from this equilibrium relationship, namely z_t, represents

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[1] Note that, in this case, s_t and x_t will also be cointegrated.

[2] One classical example of optimal behavior is the case of inventory determination, as shown in Granger and Lee (1989) seminal paper documenting multicointegration of production and sales in many US industries and industrial aggregates. Other influential applied contributions in this area include, among others, the Hendry *et al.* (1981) study of the UK consumption function obtained from a proportional integral and derivative control model, and Lee (1996) analysis of the US housing sector.

[3] In this set up, the existence of cointegration [i.e. $z_t \sim I(0)$] supports the sustainability of the fiscal process. Moreover, fiscal policy is defined as "strongly" sustainable if the cointegrating scalar equals one ($b = 1$), and "weakly" sustainable if the same parameter is positive but lower than one ($0 < b < 1$) (see Quintos, 1995).

now a proxy for the current period deficit (or surplus). As a consequence, the cumulated error series (s_t) is a proxy for the stock of outstanding debt (or savings).

The focus of the standard cointegration approach on the "flow" equilibrium relationship is motivated by the assumption of a constant discount rate. This assumption, however, is highly restrictive, as underlined in Bohn (1995) where theoretical issues related to the sustainability of budget deficits are reconsidered in the framework of a stochastic intertemporal general equilibrium model.

As discussed in the recent literature, the more realistic assumption of a stochastic discount factor implies that standard cointegration tests do not provide sufficient criteria to assess the sustainability of fiscal policy (see, e.g. Leachman *et al.*, 2005; Kia, 2008; Kia and Gardner, 2009; Kiran, 2011). More specifically, besides the usual "flow" cointegration relationship between revenues (y_t) and expenditures (x_t), it is necessary to investigate the existence of a "stock-flow" cointegration relationship between government debt (s_t) and revenues (y_t) [or between government debt (s_t) and expenditures (x_t)]. The assessment of sustainability in a stochastic environment requires therefore to explore the existence of multicointegration between revenues and expenditures.

The economic intuition about the role of multicointegration in this context can be better grasped focusing on "bad" states of nature, i.e. on circumstances in which the growth rate of the economy is lower than the real interest rate on sovereign debt. In these cases, although low or negative income growth destabilizes the debt/income ratio, multicointegration guarantees an endogenous policy response supporting fiscal sustainability.

Public debt and government revenues (and expenditures) cannot drift too far apart in the long run if multicointegration holds. This means that, in "bad" states of nature, when debt tends to become too large and its rate of increase accelerates, taxes are increased and/or expenditures are decreased in order to counteract these tendencies.

Multicointegration involves therefore an optimal policy response of the government, reacting both to the level and to the rate of change of relevant economic variables, and therefore satisfying its intertemporal budget constraint in a stochastic environment.

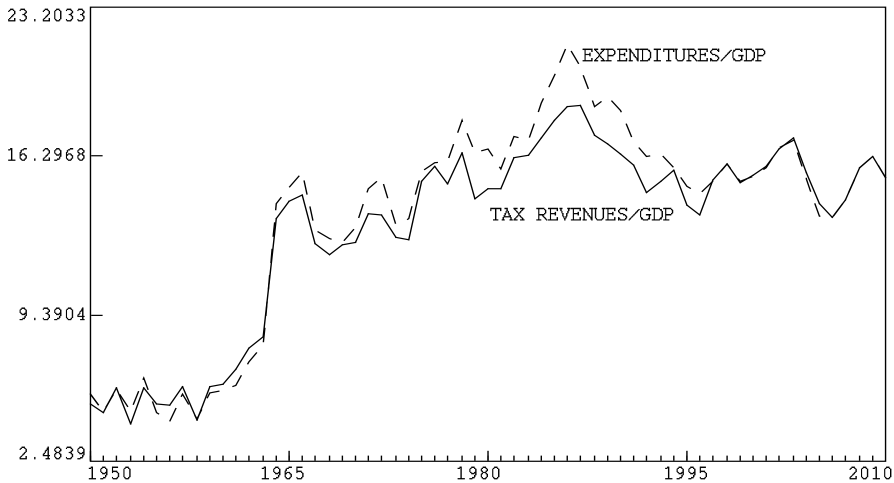
2.2 Data

I adopt the same set of basic variables employed in Tronzano (2013). More specifically, I focus on the ratios of Central Government

Revenues and Central Government Expenditures to nominal GDP. As discussed in Tronzano (2013), the use of fiscal variables expressed as ratios to GDP is mainly motivated, in the present context, by the explosive behavior of real revenues and real expenditures series. These ratios, moreover, are quite commonly employed both in the literature relying on standard cointegration methodologies (Goyal *et al.*, 2004; Payne *et al.*, 2008) and in research applying multicointegration techniques (Leachman and Francis, 2000; Leachman *et al.*, 2005).

The empirical investigation is based on annual data on Central Government Receipts and Expenditures (inclusive of interest payments) spanning from 1950 to 2010. Revenues and Expenditures data from 1950 to 1970 are taken form Jha and Sharma (2004). Since 1971 onwards, these series were updated using data on Central Government Receipts and Expenditures published by the Bank of India ("Handbook of Statistics on Indian Economy", Part 1, Annual Series – Public Finances; Tables 101 and 103, pp. 190-192). All data on nominal GDP are from the IMF "International Financial Statistics"[4].

FIGURE 1 - *India: Shares of Tax Revenues and Government Expenditures on Nominal GDP*



[4] Since Indian GDP is expressed in billions of Rupees, whereas Central Government Receipts and Expenditures are expressed in Rupees crores, the former series was multiplied by 100 before computing the percentage shares of tax revenues and government expenditures on nominal GDP.

Figure 1 plots the evolution of the Revenue/GDP and Expenditure/GDP ratios. Both series exhibit a strong increasing trend up to the first half of the 1980's, while subsequently oscillating inside a rather small range. A notable feature is that the expenditure share exceeds, on average, the revenue share.

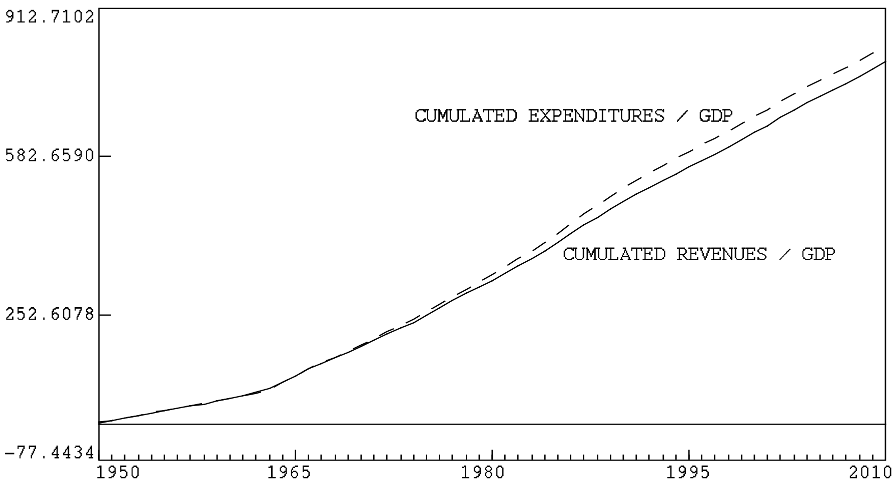
Tronzano (2013) performs a wide battery of unit root tests on these fiscal variables (including standard methodologies and the testing procedures outlined in Perron, 1990; Kwiatkowski *et al.*, 1992; Zivot and Andrews, 1992 and Elliott *et al.*, 1996) and finds overwhelming evidence that these series are integrated of order one [I(1)][5].

These fiscal variables correspond to the bivariate system of I(1) series (y_t, x_t) defined in section 2.1.

Most econometric methodologies applied in the present paper rely, however, on the use of cumulated values of (y_t) and (x_t), defined respectively as $Y_t = \sum_{j=0}^{p} y_{t-j}$, and $X_t = \sum_{j=0}^{p} x_{t-j}$.

Figure 2 outlines the evolution of these cumulated fiscal variables.

FIGURE 2 - *India: Cumulated Series of Tax Revenue/GDP and Government Expenditures/GDP*



Note that, since the original series (y_t, x_t) are I(1), these cumulated series (Y_t , X_t) are I(2) by construction. Both series exhibit the typical

[5] See Tronzano (2013), section 2, for a detailed discussion of these results.

smooth pattern characterizing I(2) variables. Moreover, while these variables appear closely tied together during the first part of the sample, since the early 1970's a significant divergence is apparent, with cumulated expenditures rising at a higher pace than cumulated revenues.

2.3 Some Preliminary Results

This section presents some preliminary results based on the seminal two-step approach put forward in Granger and Lee (1989).

The first step of this methodology involves estimating a standard cointegrating regression between y_t and x_t in order to obtain a super-consistent estimate of (b). If cointegration exists [i.e. if $z_t = y_t - bx_t$ is I(0)], the error series from this first-step regression is cumulated, thus generating the following I(1) variable: $s_t = \sum_{j=0}^{p} z_{t-j}$. In the second step, the cumulated variable s_t is regressed on y_t (and/or on x_t) in order to obtain a super-consistent estimate of the latter cointegrating parameter. Multicointegration between y_t and x_t holds if residuals from this second-stage regression yield a stationary equilibrium error.

Focusing on the topic addressed in this paper, y_t and x_t correspond, respectively, to the Revenue/GDP and Expenditure/GDP ratios. The first-stage regression reads:

$$y_t = a + bx_t + \varepsilon_t \tag{1}$$

where (a) is a constant term, (b) is the cointegrating parameter, and (ε_t) is an error term. If cointegration between y_t and x_t holds, the second step involves assessing whether a further cointegrating relationship exists between the cumulated error term from equation (1) (i.e. $s_t = \sum_{j=0}^{p} \varepsilon_{t-j}$) and y_t (or x_t)[6].

The implementation of two-step multicointegration tests draws on the empirical investigation carried out in Tronzano (2013).

In the first-stage regression step, I use four alternative specifications of equation (1) which, as documented in Tronzano

[6] As discussed in section 2.1, the cumulated error series (s_t) represents, in this case, a proxy for the stock of outstanding debt (or savings), and the existence of a second cointegrating vector captures the long run stock-flow equilibrium relationship necessary to ensure a sustainable fiscal policy in all states of nature.

(2013), yield a significant cointegrating relationship between y_t and x_t. These specifications include a standard cointegrating regression incorporating a constant and a linear trend, and three further specifications allowing for a structural break based on the approach originally outlined in Gregory and Hansen (1996). These structural break specifications assume, respectively, a shift in the constant term (model [C] in Gregory and Hansen, 1996), a shift in the constant term in the presence of a linear time trend (model [C/T]), and a shift in the constant and in the slope terms (model [C/S])[7].

In the second step, the (stationary) residuals from the four different specifications described above are cumulated, and the existence of cointegration between (s_t) and y_t (or x_t) is explored through a standard ADF test.

Table 2 summarizes the results from these second-step regressions. The left-hand side column indicates the alternative specifications of equation (1) estimated in the first step. The remaining columns report the results of stationarity tests on residuals from second-stage regressions involving, respectively, Tax Revenues and Government Expenditures.

The results are remarkably stable across alternative specifications of first-stage regressions. Moreover, these results are not significantly affected by the specific fiscal variable included in the second-stage regressions.

Multicointegration between the Revenue/GDP ratio and the Expenditure/GDP ratio is strongly rejected by this empirical evidence. The ADF statistics never exceed the critical values for cointegration tests reported in Engle and Yoo (1987), at conventional significance levels. These statistics, therefore, do not allow to reject the null hypothesis of non-stationary residuals for all estimated second-stage regressions.

Overall, these multicointegration tests suggest that the fiscal process in India is not sustainable in a stochastic environment. In

[7] Since the variables used and the sample period considered are identical to Tronzano (2013), the results from these first-stage regressions are not reported here. The reader is referred to Tronzano (2013), sections 3.1 - 3.2, for a discussion of this empirical evidence. More details about the specification of structural break models can be found in Tronzano (2013), section 3.2. Recursive empirical estimates carried out in Tronzano (2013) identify one structural break in 1994 for models [C] and [C/S], and one regime shift in 1977 for model [C/T].

other words, under particularly adverse economic conditions ("bad" states of nature), the government might be tempted to inflate away the debt or to resort to Ponzi financing schemes. This result contrasts with most applied literature on India relying on standard cointegration techniques, according to which fiscal policy in this country is at least "weakly" sustainable (Jha and Sharma, 2004; Goyal *et al*., 2004; Tronzano, 2013).

The findings reported in Table 2 have a preliminary character, since the underlying econometric approach raises some statistical problems. The robustness of these results is extensively explored in the next section, applying more powerful econometric techniques.

TABLE 2 - *Two-Step Multicointegration Tests*

Estimated First-Stage Regression	ADF Test on Residuals from Second-Stage Regression: Tax Revenues (y_t) and (s_t)	ADF Test on Residuals from Second-Stage Regression: Government Expenditures (x_t) and (s_t)
Standard Model (Constant and Time Trend)	-2.17	-2.13
Break in Constant Term	-2.56	-2.66
Break in Constant Term (Time Trend included)	-1.80	-1.80
Break in Constant and Slope Terms	-2.60	-2.71

Critical values for the ADF statistic are as follows: -2.90 (10%); -3.29 (5%); -4.12 (1%). See Engle-Yoo (1987), Table 3.
The optimal lag length in auxiliary equations used to compute the ADF statistic was selected through standard AIC and SBC information criteria.
The estimated first-stage regressions allowing for structural breaks correspond to the three alternative models outlined in Gregory-Hansen (1996):
Break in Constant Term: "Level-shift" model (Gregory-Hansen, 1996: model [C]);
Break in Constant Term (Time Trend included): "Level-shift with trend" model (Gregory-Hansen, 1996: model [C/T]);
Break in Constant and Slope Terms: "Regime-shift" model (Gregory-Hansen, 1996: model [C/S]).

3. EMPIRICAL EVIDENCE

3.1 Single-Step Multicointegration Tests

The Granger and Lee (1989) methodology suffers from a generated regressor problem, which arises when the cointegrating vector in the system of I(1) variables is not known in advance and requires to be estimated. In this case, the auxiliary regression run in the second step is based on cumulated regression residuals from another regression, and this complicates the limiting theory to test for multicointegration. Conventional cointegration tests become inappropriate, under these circumstances, since the asymptotics are no more expressed in terms of functionals of a Brownian motion process as is usually the case[8].

Engsted *et al*. (1997) outline an alternative methodology based on an integral equation incorporating both I(1) and I(2) variables among the regressors. The main advantage of this new approach is that, differently from the two-step procedure, the distributions of test statistics are well known.

In the present context, this methodology requires the estimation of the following regression:

$$Y_t = \delta_0 + \delta_1(t) + \delta_2(t)^2 + K_0\,(X_t) + K_1\,(\Delta X_t) + \varepsilon_t \qquad (2)$$

where (t) and $(t)^2$ are respectively a linear and a quadratic time trend; $Y_t = \sum_{j=0}^{p} y_{t-j}$ represents cumulated government revenues [I(2)]; $X_t = \sum_{j=0}^{p} x_{t-j}$ represents cumulated government expenditures [I(2)]; and $\Delta X_t = x_t$ is a flow variable corresponding to government expenditures [I(1)][9].

The single-step methodology relying on equation (2) simultaneously tests both levels of cointegration by exploiting the fact that multicointegration implies a particular form of I(2) cointegration. This procedure is a natural generalization of the residual based test for cointegration developed in Engle and Granger (1987) for the case of I(1) variables. More specifically, the idea is to test whether the least squares regression residuals (ε_t) from equation (2) follow an I(0)

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[8] See Engsted *et al*. (1997), section 2, and the references quoted therein for a more technical discussion of this topic. Note, moreover, that the Granger and Lee (1989) methodology is perfectly valid if the cointegrating vector in the system of I(1) variables is known in advance and therefore does not require a preliminary estimation. This happens, for instance, in the seminal work of Granger and Lee (1989) on production, sales and inventories, where the (b) parameter is *a priori* restricted to be equal to one.

[9] The introduction of deterministic trend components in equation (2) is motivated by the fact that the generated I(2) variables in this cointegrating regression (Y_t, X_t) may potentially have a drift (see Engsted *et al*., 1997 and Engsted and Haldrup, 1999).

process (the case of multicointegration), an I(1) process (the case of first level cointegration but no multicointegration), or an I(2) process (the case of no cointegration among fiscal variables).

Since in most cases I(2) variables cointegrate at the I(1) level at least, the null hypothesis is that (ε_t) is I(1) (first level cointegration); the alternative hypothesis is instead that (ε_t) is I(0), namely the existence of multicointegration among fiscal variables (y_t, x_t). Critical values for the ADF t-test on (ε_t) depend on the number of I(1) variables (m_1) and I(2) variables (m_2) on the right-hand side of the cointegrating regression, as well as on the specific deterministic components included. Note finally that, if multicointegration holds, the least squares estimation of equation (2) will yield a super consistent estimate of the parameter associated with the I(1) variable (K_1), and a super-super consistent estimate of that associated with the I(2) variable (K_0) (Haldrup, 1994, Theorem 1; Engsted *et al.*, 1997, section 2.2).

It may be useful to briefly discuss the meaning of key parameters in equation (2) and the restrictions on these parameters consistent with sustainability criteria. Fiscal sustainability does not impose any *a priori* restriction on K_0. If $K_0 > 1$, cumulated revenues are on average higher than cumulated expenditures, leading to the accumulation of government surpluses. If $K_0 = 1$, a balanced budget prevails on average. Conversely, if $K_0 < 1$ expenditures exceed revenues on average, and the possibility of a default on public debt arises. If $K_0 < 1$, fiscal sustainability in a stochastic environment requires government expenditures and the present value of debt to be negatively related, namely $K_1 < 0$. This negative relationship ensures that expenditures fall to accommodate rising levels of debt, thus preventing the government or the private sector to engage in a Ponzi scheme or gamble (Leachman *et al.*, 2005, sections 3-4)[10].

Table 3 summarizes the empirical evidence from single-step multicointegration tests.

Consistently with the above discussion, I focus on three alternative specifications of equation (2). The former includes only a constant term among deterministic regressors ($\delta_1 = \delta_2 = 0$). The remaining specifications allow, respectively, for a linear trend ($\delta_1 \neq 0$;

[10] When the government accumulates a positive saving ($K_0 > 1$), the constraint on the coefficient relative to the flow variable is clearly reversed ($K_1 > 0$). In this case, $K_1 > 0$ ensures that government expenditures rise to accommodate increasing levels of savings.

$\delta_2 = 0$), and for a linear and a quadratic trend ($\delta_1 \neq 0$; $\delta_2 \neq 0$) among deterministic regressors.

TABLE 3 - *Single-Step Multicointegration Tests: Standard Models*

Models	δ_0	δ_1	δ_2	K_0	K_1	ADF
Constant	2.44 (1.73)	-	-	0.94 (415.4)	0.13 (1.05)	- 0.846
Constant + Trend	- 1.97 (- 1.47)	1.39 (5.89)	-	0.86 (59.1)	- 0.45 (- 3.19)	- 1.269
Constant + Trend + Quadratic Trend	- 4.15 (- 7.97)	1.26 (14.1)	0.021 (18.6)	0.77 (104.5)	0.44 (6.18)	- 2.831

t-statistics in parentheses below parameters values.
Critical values for this single-step multicointegration test are as follows:
Model with constant term: -4.65 (1%); -3.93 (5%); -3.60 (10%) ; see Haldrup (1994), Table 1, p. 168.
Model with constant term and linear trend: -5.11 (1%); -4.42 (5%); -4.08 (10%); see Engsted *et al*. (1997), Table 1, p. 263.
Model with constant term, linear trend and quadratic trend: -5.56 (1%); -4.83 (5%); -4.47 (10%); see Engsted *et al*. (1997), Table 2, p. 264.

As shown in Table 3, the most general specification yields highly significant parameters estimates for both deterministic and stochastic regressors; this specification, moreover, is clearly the one providing the best fit of the data[11].

The selection of the optimal lag order for the ADF test on $\{\varepsilon_t\}$ is based on the sequential t-test approach outlined in Ng and Perron (1995): I start with a maximum lag of 4 and set a cut-off level of 0.05 for the p-value of the last lagged coefficient.

Overall, the empirical evidence strongly support the absence of multicointegration among fiscal variables. Focusing on the last column of Table 3, the ADF test never rejects the null hypothesis of I(1) residuals across alternative specifications of equation (2). Even the preferred specification, including linear and quadratic trends,

[11] The standard error of the regression incorporating a linear and a quadratic trend amounts to 1.05. Notably higher values are recorded for the model including only a constant term (3.52) and for that allowing for a constant and a linear trend component (2.79). More details about the results from alternative estimates of equation (2) are available by the author upon request.

yields a negative test statistic (-2.83) largely above critical values at conventional significance levels, thus not supporting the existence of stationary I(0) residuals.

As regards parameters estimates, the absence of multicointegration precludes to draw robust statistical inferences. The estimates of K_0 appear anyway fairly stable and consistently lower than one. This points out that government expenditures exceed on average government revenues, in line with the evidence discussed in Tronzano (2013) and with the patterns of cumulated fiscal variables shown in Figure 2. The estimates of K_1 reveal instead a substantial degree of instability. Although the sign of K_1 deserves no particular interest in this context (given the absence of multicointegration), it is interesting to observe that the preferred specification of equation (2) delivers a positive value of this parameter, thus suggesting an inappropriate government response to the accumulation of public debt.

This empirical evidence was submitted to two different robustness checks. It is well known that an integral cointegration regression can be estimated using ΔY_t instead of ΔX_t as a regressor (Engsted *et al.,* 1997, section 2.2). Equation (2) was therefore re-estimated substituting the flow of government expenditures with that of government revenues on the right-hand side. In a latter empirical exercise, equation (2) was normalized on cumulated government expenditures, namely Y_t was replaced with X_t as a left-hand variable. Both robustness checks were implemented using all the deterministic specifications outlined above. Overall, these additional estimates never allowed to reject the null hypothesis of I(1) residuals, thus strongly reiterating the absence of multicointegration[12].

To sum up, the results from the single-step tests are closely in line with previous findings, pointing out the existence of first-level cointegration, but rejecting the existence of a deeper long-run equilibrium relationship between fiscal variables. The absence of multicointegration means that the government does not implement adequate corrective measures on expenditures and revenues flows in the presence of a growing amount of public debt. In short, India's fiscal policy does not appear to be sustainable in a stochastic environment.

[12] These additional results are not included in the paper in order to save space and are available from the author upon request.

3.2 Multicointegration Tests with Regime Shifts

The auxiliary equation underlying standard single-step multicointegration tests assumes that the long run stock-flow equilibrium relationship is stable over time. This assumption might actually be too restrictive in the present context.

Tronzano (2013) documents various structural breaks in India's fiscal variables. Focusing on the levels of these variables, this paper identifies a break in their intercepts in 1964, one break in the slope of the Revenue/GDP ratio in 1977, and one break in the slope of the Expenditure/GDP ratio in 1979. Moreover, standard cointegration tests on flow variables allowing for regime shifts reveal a highly significant shift in the level of the equilibrium relationship in the mid-90's and other possible breaks under alternative model specifications.

Overall, this evidence suggests that the possibility of further structural breaks in the multicointegration relationship between Indian fiscal variables should be explored in the present research. This requires an extension of the approach implemented so far since, in the presence of regime shifts, standard multicointegration tests are inappropriate.

In line with the above remarks, this section applies a new multicointegration test allowing for regime shifts. This analysis innovates on the existing literature since no recent contribution on emerging market economies employs this class of models (see, e.g. Kia, 2008; Kia and Gardner, 2009; Kiran, 2011).

I implement the approach outlined in Berenguer-Rico and Carrion-I-Silvestre (2011), where a new statistic to test various types of I(2) cointegration and multicointegration models with regime shifts is proposed. This approach extends standard multicointegration tests to incorporate structural breaks affecting both the deterministic and the stochastic part of the stock-flow equilibrium relationship[13].

Consistently with the above framework, equation (2) may be reformulated as follows:

[13] Berenguer-Rico and Carrion-I-Silvestre (2011), section 3, provide a detailed technical discussion of this approach. Theorem 1 in this section defines the limiting distribution of the new test statistic (t^*_δ) (see also Appendix A1, pp. 316-318). Theorem 2, in the same section, establishes the methodology to obtain a highly consistent estimate of the break fraction over all possible values, through the minimization of the sum of squared residuals (see also Appendix A2, pp. 318-320). Berenguer-Rico and Carrion-I-Silvestre (2011), section 4, investigate the finite sample performance of the t^*_δ statistic through Monte Carlo simulation, documenting its good properties in terms of empirical size and power.

$$Y_t = \beta_{00} + \beta_{01}\, t + \beta_{02}\, 1(t > T_b) + \beta_{03}\, (t - T_b)\, 1(t > T_b) + \alpha\, t^2 + \qquad (3)$$
$$+ \beta_{11}\, \Delta X_t + \beta_{12}\, \Delta X_t\, 1(t > T_b) + \beta_{21}\, X_t + \beta_{22}\, X_t\, 1(t > T_b) + u_t$$

where all symbols retain their previous meanings, $1(\cdot)$ is the usual indicator function, T_b denotes the break point, and u_t represents the residuals from the OLS estimation of equation (3).

Although the basic structure of this equation is similar to equation (2), some distinctive features are evident. As regards the deterministic variables, the parameters β_{02} and β_{03} capture, respectively, a shift in the level and in the trend of the equilibrium condition. As regards the stochastic components, the parameters β_{12} and β_{22} express, respectively, a change in the effect of the "flow" variable (ΔX_t) and of the "stock" variable (X_t) on cumulated government revenues (Y_t).

The stock-flow system summarized by equation (3) admits various cointegration possibilities.

If u_t is I(2) no long-run equilibrium relationship exists. If u_t is I(1), the I(2) variables (Y_t, X_t) cointegrate into an I(1) process, namely $Y_{t,}\, X_t \sim$ CI (2,1). From the economic standpoint, this implies that the government keeps revenue and expenditure flows under control, without paying attention to the stock of debt [note that in this case the stock of debt is I(1)]. Finally, if u_t is I(0) multicointegration holds. The I(2) variables now cointegrate into a stationary process, namely $Y_{t,}\, X_t \sim$ CI (2,2). In this case, not only does the government equilibrate revenue and expenditure flows, but it also intervenes on flow variables whenever some exogenous shock tends to drive debt out of control. The stock of debt is now I(0); moreover, two cointegration relationships exist: the former between two flows, and the latter between one stock and one flow variable. This means that fiscal sustainability is achieved controlling both the flow variables and the stock of debt.

Berenguer-Rico and Carrion-I-Silvestre (2011) observe that in most situations it is likely that two I(2) variables will at least cointegrate at the I(1) level. On this basis, this approach assumes first-level cointegration [i.e. $u_t \sim$ I(1) and Y_t, $X_t \sim$ CI(2,1)] as the null hypothesis. This null is tested against the alternative hypothesis of multicointegration with structural breaks [i.e. $u_t \sim$ I(0) and Y_t, $X_t \sim$ CI(2,2), in the presence of significant structural breaks in equation (3)].

In order to assess the stochastic properties of u_t, the following ADF-type regression is estimated for each possible break point in the sample:

$$\Delta u_t = \delta u_{t-1} + \sum_{j=0}^{p} \phi_j\, \Delta u_{t-j} + \eta_t \qquad (4)$$

Denoting with $t_\delta(\lambda)$ the t-ratio ADF statistic testing the null hypothesis that $\delta = 0$ in equation (4), this procedure yields a sequence of ADF-statistics for each possible break point $\lambda \in \Lambda$[14].

The infimum value of this sequence, denoted as:

$$t^*_\delta = \inf_{\lambda \in \Lambda} t_\delta(\lambda)$$

represents the test statistic to assess the existence of multicointegration with structural breaks. The limiting distribution of t^*_δ depends on the particular deterministic specification assumed in equation (3), as well as on the number of I(1) and I(2) stochastic regressors (denoted, respectively, as m_1 and m_2). Critical values of t^*_δ , for various specifications of the underlying model and different values of m_1 and m_2, are tabulated in Berenguer-Rico and Carrion-I-Silvestre, 2011 (see Table II, p. 305).

As regards the detection of T_b, the date associated with t^*_δ does not provide, in general, a consistent estimate of the break point. If multicointegration holds, Berenguer-Rico and Carrion-I-Silvestre (2011) prove that a consistent estimate of the break date can be obtained from the minimization of the sum of squared residuals (SSR) in equation (3) over all possible values of the break fraction (see ibid., Theorem 2, pp. 318-320).

A nice feature of this econometric approach is that the regime-shift model expressed by equation (3) includes many alternative specifications, which can be derived imposing suitable restrictions on model's parameters. I exploit this feature in the present section, focusing both on the most general specification and on some relevant nested alternatives for which specific critical values are provided in Berenguer-Rico and Carrion-I-Silvestre (2011).

Table 4 summarizes the alternative regime-shift models estimated in the present section.

The former two specifications (models [A], [B]) do not assume any structural break in the stochastic regressors ($\beta_{12} = \beta_{22} = 0$). Model [A] includes a constant term and a linear trend as deterministic regressors ($\alpha = 0$), whereas model [B] allows also for a quadratic trend ($\alpha \neq 0$)[15]. The remaining specifications (models [C], [D], [E])

[14] In line with the previous literature on regime-shifts cointegration models (see Gregory and Hansen, 1996), the break point date is treated as unknown. The break point date may be expressed as $T_b = [\lambda n]$, where n is the sample size, and λ is the break fraction parameter. More specifically, $\lambda \in \Lambda$ where Λ is a closed subset of (0,1), with $[\cdot]$ being the integer part.

[15] Note that model [B] is the only one incorporating a quadratic trend, whereas for all other models the restriction $\alpha = 0$ holds.

allow for structural breaks both in the deterministic and in the stochastic regressors. Focusing on stochastic regressors, the regime-shift involves the I(2) variable (X_t) in the case of model [C] ($\beta_{22} \neq 0$), and the I(1) variable (ΔX_t) in the case of model [D] ($\beta_{12} \neq 0$). Model [E], finally, represents the most general specification, with no restrictions on parameters of stochastic regressors ($\beta_{12} \neq 0$; $\beta_{22} \neq 0$) and structural breaks affecting both I(1) and I(2) variables.

TABLE 4 - *Multicointegration Models with Regime Shifts: Alternative Specifications*

Models	Deterministic Part	Stochastic Part
Model [A]	$\alpha = 0$	$\beta_{12} = \beta_{22} = 0$
Model [B]	$\alpha \neq 0$	$\beta_{12} = \beta_{22} = 0$
Model [C]	$\alpha = 0$	$\beta_{12} = 0$; $\beta_{22} \neq 0$
Model [D]	$\alpha = 0$	$\beta_{12} \neq 0$; $\beta_{22} = 0$
Model [E]	$\alpha = 0$	$\beta_{12} \neq 0$; $\beta_{22} \neq 0$

The coefficients reported in this table refer to the regime-shift multicointegration relationship [equation (3)] analyzed in the present section, which is reproduced below for convenience:

$Y_t = \beta_{00} + \beta_{01}\, t + \beta_{02}\, 1(t > T_b) + \beta_{03}\, (t - T_b)\, 1(t > T_b) + \alpha\, t^2 + \beta_{11}\, \Delta X_t + \beta_{12}\, \Delta X_t\, 1(t > T_b) + \beta_{21}\, X_t + \beta_{22}\, X_t\, 1(t > T_b) + u_t$

See the main text for the meaning of all symbols appearing in this equation.

The regime-shift models outlined above provide the basis for the multicointegration tests implemented in the present section. Since the break point date (T_b) is treated as unknown, equation (4) was sequentially estimated for each possible break point along the sample. At each iteration, the order of the parametric correction (p) was selected using the t-sig criterion of Ng and Perron (1995), setting $p_{max} = 4$ as the maximum number of lags.

Table 5 reports the values of the t^*_δ statistic for each model.

As shown in this table, the values of t^*_δ do not exceed their corresponding critical values at standard significance levels. This result is robust across all models, starting from the simpler specification allowing for a structural break in deterministic regressors (model [A]), up to the most general specification allowing for structural breaks in all stochastic regressors (model [E]).

This empirical evidence points out that the null hypothesis of I(1) residuals can never be rejected for Indian fiscal variables. This implies that the stocks of I(2) fiscal variables exhibit first-order cointegration [i.e. Y_t, X_t ~ CI(2,1)], but do not display full cointegration [i.e. Y_t, X_t are not CI(2,2)]. These results support therefore the existence of a simple cointegration relationship between flow variables, but exclude the existence of a deeper equilibrium relationship between stock and flow variables (multicointegration) in the presence of structural breaks.

TABLE 5 - *Multicointegration Tests with Regime Shifts*

Models	$t^*_\delta = \inf t_\delta (\lambda)$	5% C.V.	10% C.V.
Model [A]	- 4.717	- 6.50	- 6.23
Model [B]	- 5.267	- 6.96	- 6.64
Model [C]	- 4.770	- 6.81	- 6.49
Model [D]	- 4.99	- 6.71	- 6.43
Model [E]	- 4.858	- 6.97	- 6.65

$t^*_\delta = \inf t_\delta (\lambda)$ is the multicointegration test statistic developed in Berenguer-Rico and Carrion-I-Silvestre (2011).
Critical values displayed in this table are from Berenguer-Rico and Carrion-I-Silvestre, 2011 (Table II, p. 305) and refer to a sample of 50 observations.
These critical values assume $m_1 = 1$ and $m_2 = 1$, where m_1 and m_2 correspond, respectively, to the number of I(1) and I(2) variables included in the multicointegration regression.

The robustness of these findings was assessed in a slightly different framework where the flow fiscal variable (ΔX_t) is removed from equation (3), thus imposing *a priori* the restriction $\beta_{11} = \beta_{12} = 0$. In formal terms, this means working inside a simple I(2) cointegration framework allowing for structural breaks, instead of the equivalent multicointegration set up considered above. These additional estimates reiterate the existence of non-stationary I(1) residuals for all model's specifications[16].

The parameters of stochastic regressors can be estimated super-consistently (β_{11}, β_{12}) and super-super-consistently (β_{21}, β_{22}) only in the presence of multicointegration (Haldrup, 1994; Engsted *et al.*, 1997). The absence of multicointegration precludes therefore to draw

[16] Various I(2) cointegration models were estimated, using different specifications of the deterministic component and imposing alternative forms of structural breaks. These additional results are not reported in order to save space, and are available from the author upon request.

accurate inferences on these parameters, and to use more efficient estimators displaying better small sample properties (such as DOLS, see Stock and Watson, 1993).

Notwithstanding the above caveats, it may be interesting to inspect coefficients estimates from alternative regime-shifts specifications, and to compare some basic statistics associated with these models. Table 6 summarizes this information.

The first two columns report the coefficients estimates of deterministic and stochastic regressors, while the last column includes the minimized sum of squared residuals (SSR), the corresponding period (T_b), and the values of some model selection criteria (AIC, SBC).

One salient feature emerging from Table 6 is that β_{21} estimates are highly significant and stable around 0.86 for all models (the only exception being model [B] where the estimated value is slightly lower). This means that, on average, the growth of cumulated revenues has been significantly lower than that of cumulated expenditures, an empirical regularity which is apparent from Figure 2, particularly since the early 70's. The point estimates of β_{21} corroborate therefore the large support for "weak" fiscal sustainability documented in Tronzano (2013), and are closely in line with the results of the previous section.

Focusing on model selection criteria, model [B], incorporating a quadratic trend ($\alpha \neq 0$), and model [D], allowing for a regime-shift in the I(1) stochastic variable ($\beta_{12} \neq 0$), are among the best performing specifications.

As regards model [B], all coefficients are highly significant, and the estimates point out a noticeable increase in the parameters related to the linear trend in 1964 (β_{01} = 1.61 ; $\beta_{01} + \beta_{03}$ = 2.82). Although the t^*_δ statistic does not support the existence of significant structural breaks and of multicointegration, this shift in the linear trend is clearly reflected in the increasing slopes of cumulated series since the mid-60's (see Figure 2).

Turning to model [D], the coefficients of the flow fiscal variable (ΔX_t) exhibit a potential regime-shift from positive to negative values in 1994 (β_{11} = 0.265 ; β_{12} = -0.39). These coefficients are the equivalent of the K_1 parameter in equation (2) and, as discussed before, if expenditures exceed revenues on average, only a negative value of K_1 satisfies fiscal sustainability criteria in a stochastic environment. Therefore, although the results exclude the occurrence of significant structural breaks, the negative estimate of ($\beta_{11} + \beta_{12}$) after 1994

TABLE 6 - *Multicointegration Models With Regime Shifts: Parameters Estimates and Basic Statistics*

	Deterministic Regressors					Stochastic Regressors							
Models	α	β_{00}	β_{01}	β_{02}	β_{03}	β_{11}	β_{12}	β_{21}	β_{22}	SSR	T_b	AIC	SBC
Model [A]	-	-3.42 (-10.3)	1.04 (18.0)	-0.48 (-1.03)	1.20 (27.7)	0.22 (5.05)	-	0.86 (244.6)	-	25.05	1995	-65.4	-71.7
Model [B]	0.02 (29.5)	-2.74 (-7.4)	1.61 (22.0)	1.42 (3.01)	1.21 (9.48)	0.36 (7.53)	-	0.68 (62.9)	-	20.39	1964	-60.1	-67.5
Model [C]	-	-3.42 (-10.2)	1.05 (17.8)	-9.99 (-0.13)	0.96 (0.48)	0.22 (4.97)	-	0.86 (241.8)	0.016 (0.12)	25.04	1995	-66.4	-73.8
Model [D]	-	-3.53 (-10.7)	1.02 (17.6)	4.94 (1.74)	1.19 (30.04)	0.265 (5.71)	-0.39 (-2.18)	0.86 (248.3)	-	23.47	1994	-64.4	-71.8
Model [E]	-	-3.53 (-10.7)	1.02 (17.6)	-63.6 (-0.82)	-0.65 (-0.31)	0.265 (5.69)	-0.45 (-2.35)	0.86 (247.7)	0.11 (0.89)	23.12	1994	-64.9	-73.4

See footnote to Table 4 for the general specification of multicointegration models allowing for regime shifts. T_b : Break Point Date.
SSR: Minimized Sum of Squared Residuals.
AIC and SBC indicate, respectively, the Akaike and Schwarz Model Selection Criteria. t-statistics in parentheses below parameters estimates.

captures an improvement in India's fiscal outlook since the mid-90's. Quite interestingly, this evidence displays close similarities with the results obtained in Tronzano (2013) inside a standard cointegration framework (see ibid. section 3.2).

To sum up, this section extends the analysis to a new class of multicointegration models allowing for structural breaks in their deterministic and stochastic components. The main result is that the null hypothesis of non-stationary I(1) residuals is never rejected under all regime-shifts specifications. This result supports first-order cointegration between cumulated revenues and expenditures in the case of India, while rejecting the existence of multicointegration in the presence of structural breaks.

This evidence is fully in line with previous findings, pointing out the absence of multicointegration both through the seminal approach of Granger and Lee (1989), and through standard single-step methodologies. Moreover, since these tests exclude the existence of structural breaks, the simpler econometric methodologies implemented before do not suffer from model mis-specification. These estimates, finally, corroborate the empirical evidence obtained in Tronzano (2013) documenting the existence of "weak" fiscal sustainability in India.

The absence of multicointegration, strongly supported in the present paper, raises some specific policy implications which go beyond those associated with a "weakly" sustainable fiscal policy. These policy implications are addressed in full detail in the next section.

4. Concluding remarks and policy implications

A large strand of literature exploring the sustainability of fiscal budgeting processes employs standard cointegration techniques to assess the existence of an equilibrium relationship between government revenues and expenditures. In this context, empirical analyses usually discriminate between "strong" and "weak" fiscal sustainability, depending on whether the cointegrating scalar is equal to one, or strictly bounded between zero and one (see, e.g. Trehan and Walsh, 1988; Quintos, 1995).

This standard approach relies on the assumption of a constant (or stationary) interest rate, and hence of a constant (or stationary) discount rate. This assumption allows to express the government intertemporal budget constraint as a testable cointegrating relationship between revenues and expenditures (after imposing the usual transversality condition ruling out Ponzi schemes on government debt).

Recent theoretical work suggests however that the above approach does not provide adequate criteria to establish whether fiscal policy is truly sustainable. Bohn (1995) outlines a model of a stochastic economy where the discount factor depends on the probability distribution of revenues, spending and debt across states of nature. In this set up, a positive probability of a large income decline (i.e. a "bad" state of nature) can invalidate fiscal sustainability conditions even in the presence of a relatively low debt growth, since the discount factor depends on state contingent prices that can become arbitrarily large in bad states of nature.

These theoretical results call for a more encompassing set of criteria to investigate fiscal sustainability. Besides standard cointegration tests, it becomes necessary to assess whether government debt is cointegrated with revenues and expenditures, namely to implement a multicointegration analysis on fiscal variables. As formally discussed in Leachman *et al*. (2005), the existence of multicointegration identifies a sophisticated policy response, in which the government takes into account both the change and the rate of change of economic conditions in formulating its fiscal policy.

This paper extends the analysis contained in Tronzano (2013), where the sustainability of fiscal policy in India is tested using standard cointegration techniques. In line with the above discussion, I rely on a multicointegration approach to assess the sustainability of the fiscal process in a stochastic environment. Moreover, for ease of comparison with previous findings, I focus on the same period examined in Tronzano (2013) (annual data: 1950-2010) and use the same basic data set (fiscal variables expressed as ratios to nominal GDP).

The empirical investigation proceeds along three main steps. I first provide some preliminary results implementing the seminal approach outlined in Granger and Lee (1989). The analysis is subsequently extended using a more robust methodology proposed in Engsted *et al*. (1997). This approach is finally extended inside a regime-shift framework, allowing for various specifications of deterministic and stochastic components. To the best of my knowledge, no available contribution on emerging market economies applies this particular class of multicointegration models (see, e.g. Kia, 2008; Kia and Gardner, 2009; Kiran, 2011).

Overall, the empirical evidence points out that fiscal variables are not multicointegrated, and therefore that India's fiscal policy is not sustainable in a stochastic environment. This result is supported by all econometric methodologies, which do not reject the null hypothesis of non-stationary I(1) residuals.

In line with Tronzano (2013), I document the existence of first-order cointegration between revenues and expenditures flows and, more specifically, an average growth of cumulated revenues lower than that of cumulated expenditures (a result which is consistent with "weak" fiscal sustainability inside a standard cointegration approach). However, the empirical evidence excludes the existence of a deeper long-run equilibrium between stock and flow fiscal variables. Various robustness checks, using both standard and regime-shift multicointegration models, confirm this result.

The main policy implication of this research is that the Indian government should ensure a more robust and systemic link between tax and expenditures policies and the evolution of public debt. In the absence of multicointegration, debt dynamics is out of control because, notwithstanding an equilibrium relationship between revenues and expenditures, an adverse shock on the real economy may destabilize the debt pattern. Indian policy-makers should therefore establish a closer connection between fiscal policy guidelines and debt dynamics. This implies that taxes should be increased and/or spending decreased whenever debt becomes too large and/or its rate of increase is accelerating.

It is interesting to observe that this policy prescription displays substantial analogies with the approach implemented in Bohn (1998) to assess US fiscal sustainability, focusing on the response of the primary budget surplus to changes in the debt-income ratio.

Tronzano (2013) maintains that India should strengthen its commitment to fiscal consolidation because, in the presence of "weak" fiscal sustainability, speculative pressures might significantly increase the risk of default on public debt. The main components of a prudent fiscal policy stance outlined in Tronzano (2013) are a widening of the tax base in a non-distortionary fashion, and a structural reduction in inefficient government spending.

The empirical evidence obtained in the present paper is clearly consistent with these policy prescriptions. The absence of multicointegration suggests, however, that this fiscal adjustment needs to be implemented in the context of explicit targets for medium-term debt dynamics. In other words, a systematic reaction of tax and expenditures flows to contingent disequilibria in public debt is crucial in order to ensure the sustainability of India's fiscal process.

International comparisons with other emerging economies reveal that the fiscal outlook for India raises serious concerns. Focusing on a large set of emerging countries, Kochhar (2004) documents that India's ranking in terms of general government finances is particularly low, both in terms of deficit and in terms of debt ratios to GDP, with only two countries (Turkey, Argentina) displaying more unbalanced fiscal positions. A more recent comparison, based on 2009 data for 20 emerging economies, reports a gross general government debt to GDP ratio of over 80% for India, against an average value of about 40% for the whole sample (Citigroup, 2010).

During the first half of 2012, finally, two major ratings agencies (Fitch, Standard & Poor's) cut their credit outlook for India from stable to negative as a result of a significant deterioration of real GDP growth since 2011, signs of fiscal slippage, and an uncertain timing in the implementation of some key fiscal reforms.

The relatively weak fiscal position of India and the latest events quoted above provide further support to the policy implications of this paper, namely the need to restore public debt on a sustainable path in the context of the ongoing process of fiscal consolidation.

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REFERENCES

Berenguer-Rico, V. and J.L. Carrion-I-Silvestre (2011), "Regime Shifts in Stock-Flow I(2)-I(1) Systems: The Case of US Fiscal Sustainability", *Journal of Applied Econometrics*, 26(2), 298-321.

Bohn, H. (1995), "The Sustainability of Budget Deficits in a Stochastic Economy", *Journal of Money, Credit, and Banking*, 27(1), 257-271.

Bohn, H. (1998), "The Behavior of US Public Debt and Deficits", *Quarterly Journal of Economics*, 113(3), 949-963.

Citigroup Global Markets (2010), "Sovereign Debt Problems in Advanced Industrial Countries", *Global Economics View*, April 2010.

Davig, T. (2005), "Periodically Expanding Discounted Debt: A Threat to Fiscal Policy Sustainability", *Journal of Applied Econometrics*, 20(7), 829-840.

De Paula, L.F. (2008), *Financial Liberalization, Exchange Rate Regime and Economic Performance in BRICs Countries*, in: P. Arestis, L.F. De Paula (Eds), "Financial Liberalization and Economic Performance in Emerging Countries", Palgrave MacMillan: New York.

Elliott, G., T.J. Rothenberg and J.H. Stock (1996), "Efficient Tests for an Autoregressive Unit Root", *Econometrica*, 64(4), 813-836.

Engle, R.F. and C.W.J. Granger (1987), "Co-integration and Error Correction: Representation, Estimation, and Testing", *Econometrica*, 55(2), 251-276.

Engle, R.F. and B.S. Yoo (1987), "Forecasting and Testing in Cointegrated Systems", *Journal of Econometrics*, 35(1), 149-159.

Engsted, T., J. Gonzalo and N. Haldrup (1997), "Testing for Multicointegration", *Economics Letters*, 56(3), 259-266.

Engsted, T. and N. Haldrup (1999), "Multicointegration in Stock-Flow Models", *Oxford Bulletin of Economics and Statistics*, 61(2), 237-254.

Goyal, R., J.K. Khundrakpam and P. Ray (2004), "Is India's Public Finance Sustainable? Or, Are The Claims Exaggerated?", *Journal of Policy Modeling*, 26(3), 401-420.

Granger, C.W.J. and T.H. Lee (1989), "Investigation of Production, Sales and Inventory Relationships Using Multicointegration and Non-Symmetric Error Correction Models", *Journal of Applied Econometrics*, S145-S159.

Granger, C.W.J. and T.H. Lee (1990), *Multicointegration*, in: G.F. Rhodes, T.B. Fomby, (Eds), "Advances in Econometrics", Vol. 8, 71-84, JAI Press: New York.

Gregory, A.W. and B.E. Hansen (1996), "Residual-Based Tests for Cointegration in Models with Regime Shifts", *Journal of Econometrics*, 70(1), 99-126.

Hakkio, C.S. and M. Rush (1991), "Is the Budget Deficit 'Too Large'?", *Economic Inquiry*, 29(3), 429-445.

Haldrup, N. (1994), "The Asymptotics of Single-Equation Cointegration Regressions With I(1) and I(2) Variables", *Journal of Econometrics*, 63(1), 153-181.

Hamilton, J.D. and M.A. Flavin (1986), "On the Limitations of Government Borrowing: A Framework for Empirical Testing", *American Economic Review*, 76(4), 808-819.

Hendry, D. and T. von-Ungern Sternberg (1981), *Liquidity and Inflation Effects of Consumer Expenditure*, in: A.S. Deaton (Ed.), "Essays in the Theory and Measurement of Consumers' Behaviour", Cambridge University Press: Cambridge.

Jha, R. and A. Sharma (2004), "Structural Breaks, Unit Roots, and Cointegration: A Further Test of the Sustainability of the Indian Fiscal Deficit", *Public Finance Review*, 32(2), 196-219.

Kia, A. (2008), "Fiscal Sustainability in Emerging Countries: Evidence from Iran and Turkey", *Journal of Policy Modeling*, 30(6), 957-972.

Kia, A. and N. Gardner (2009), "Analyzing the Fiscal Process Under a Stochastic Environment", Economic Research Forum Working Paper No. 475, Utah Valley University.

Kiran, B. (2011), "Sustainability of the Fiscal Deficit in Turkey: Evidence from Cointegration and Multicointegration Tests", *International Journal of Sustainable Economy*, 3(1), 63-76.

Kocchar, K. (2004), "Macroeconomic Implications of the Fiscal Imbalances", Paper presented at the IMF/NIPFP Conference on Fiscal Policy in India, New Delhi, 16-17 January 2004.

Kwiatkowski, D., P.C. Phillips, P. Schmidt and Y. Shin (1992), "Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root", *Journal of Econometrics*, 54(1-3), 159-178.

Leachman, L. and B. Francis (2000), "Multicointegration Analysis of the Sustainability of Foreign Debt", *Journal of Macroeconomics*, 22(2), 207-227.

Leachman L., A. Bester, G. Rosas and P. Lange (2005). "Multicointegration and Sustainability of Fiscal Practices", *Economic Inquiry*, 43(2), 454-466.

Lee, T. (1996), "Stock Adjustment for Multicointegrated Series", *Empirical Economics*, 21(4), 633-639.

Martin, G.M. (2000), "US Deficit Sustainability: A New Approach Based on Multiple Endogenous Breaks", *Journal of Applied Econometrics*, 15, 83-105.

Ng, S. and P. Perron (1995), "Unit Root Tests in ARMA Models with Data Dependent Methods for Selection of the Truncation Lag", *Journal of the American Statistical Association*, 90(429), 268-281.

Payne, J.E., H. Mohammadi and M. Cak (2008), "Turkish Budget Deficit Sustainability and the Revenue-Expenditure Nexus", *Applied Economics*, 40(7-9), 823-830.

Perron, P. (1990), "Testing for a Unit Root in a Time Series with a Changing Mean", *Journal of Business & Economic Statistics*, 8(2), 153-162.

Quintos, C.E. (1995), "Sustainability of the Deficit Process with Structural Shifts", *Journal of Business & Economic Statistics*, 13(4), 409-417.

Stock, J.H. and M.W. Watson (1993), "A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems", *Econometrica*, 61(4), 783-820.

Trehan, B. and C.E. Walsh (1988), "Common Trends, the Government's Budget Constraint, and Revenue Smoothing", *Journal of Economic Dynamics and Control*, 12(2), 425-444.

Tronzano, M. (2013), "The Sustainability of Indian Fiscal Policy: A Reassessment of the Empirical Evidence", *Emerging Markets Finance and Trade*, 49(S1) 63-76.

Wilcox, D.W. (1989), "The Sustainability of Government Deficits: Implications of the Present-Value Borrowing Constraint", *Journal of Money, Credit, and Banking*, 21(3), 291-306.

Zivot, E. and D.W.K. Andrews (1992), "Further Evidence on the Great Crash, the Oil Price Shock, and the Unit-Root Hypothesis", *Journal of Business & Economic Statistics*, 10(3), 251-270.

ABSTRACT

This paper assesses the sustainability of fiscal policy in India from 1950 to 2010 using multicointegration techniques. Starting from the seminal approach of Granger and Lee (1989), the analysis is extended through more powerful econometric methodologies, including several types of multicointegration models allowing for regime-shifts. Overall, the evidence rejects multicointegration between revenues and expenditures, thus pointing out that India's fiscal policy is not sustainable in a stochastic environment. The main policy implication is that, in the context of the ongoing fiscal consolidation process, the Indian government should establish a more systematic connection between fiscal policy guidelines and the evolution of public debt.

Keywords: Fiscal Sustainability, India, Cointegration, Multicointegration, Structural Breaks

JEL Classification: C22, E62, H60

RIASSUNTO

Multicointegrazione e sostenibilità fiscale in India: evidenze empiriche da modelli standard e modelli incorporanti cambiamenti di regime

L'articolo esamina la sostenibilità della politica fiscale in India dal 1950 al 2010 attraverso l'approccio di multicointegrazione. Partendo dalla metodologia inizialmente proposta in Granger-Lee (1989), l'analisi viene estesa a tecniche econometriche più efficienti basate su modelli di multicointegrazione che incorporano la possibilità di cambiamenti di regime. Nel complesso, i risultati respingono l'esistenza di multicointegrazione tra entrate e spese, evidenziando che la politica fiscale in India non è sostenibile in un contesto stocastico. La principale implicazione di politica economica è che, nel corso dell'attuale processo di consolidamento fiscale, il governo indiano dovrebbe stabilire una connessione più stretta tra obiettivi di politica fiscale e dinamica del debito pubblico.