The sustainability of Swedish fiscal policy: a re-examination

Arcade Ndoricimpa
Faculty of Economics and Management, University of Burundi, Bujumbura, Burundi

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Abstract

Purpose – This study reexamines the sustainability of fiscal policy in Sweden.

Design/methodology/approach – To test the sustainability of fiscal policy, two approaches are used; the methodology of Kejriwal and Perron (2010), testing for multiple structural changes in a cointegrated regression model and time-varying cointegration test of Bierens and Martins (2010), and Martins (2015).

Findings – Using the first approach of testing for multiple structural changes in a cointegrated regression model, the results indicate that government spending and revenue are cointegrated with two breaks. An estimation of a two-break long-run model shows that the slope coefficient increases from 0.678 to 0.892 from the first to the second regime, implying that fiscal deficits were weakly sustainable in the first two regimes, from 1800 to 1943, and from 1944 to 1974. Further, results from time-varying cointegration test indicate that cointegration between spending and revenue in Sweden is time-varying. Fiscal deficits were found to be unsustainable for the periods 1801–1811, 1831–1838, 1853–1860 , 1872–1882, 1897–1902, 1929–1940 and 1976–1982 and weakly sustainable over the rest of the study period.

Research limitations/implications – A number of implications arise from this study: (1) Accounting for breaks in cointegration analysis and in the estimation of the level relationship between spending and revenue is very important because ignoring breaks may lead to an overestimated slope coefficient and hence a bias on the magnitude of fiscal deficit sustainability. (2) In testing for cointegration between spending and revenue, assuming a constant cointegrating slope when it is actually time-varying can also be misleading because deficits can be sustainable for a period of time and unsustainable over another period.

Originality/value – The contribution of this study is three-fold; first, the study uses a long series of annual data spanning over a period of two centuries, from 1800 to 2011. Second, because of the importance of structural change in economics, to examine the existence of a level relationship between spending and revenue, the study uses the methodology of Kejriwal and Perron (2010) to test for multiple structural changes in a cointegrated regression model, as well as time-varying cointegration of Bierens and Martins (2010) and Martins (2015).

Keywords Fiscal deficits, Multiple breaks, Cointegration, Time-varying cointegration, Sweden

Paper type Research paper

1. Introduction

Achieving fiscal sustainability for countries in the European Union zone has been paramount given the Maastricht requirement for membership in the European Monetary Union. Indeed, fiscal sustainability has been an explicit criterion with a budget deficit threshold set at 3% of GDP. There is a consensus that fiscal policy has three main functions; i.e. economic stabilization, allocation and redistribution. The fiscal policy landscape of Sweden has evolved over time. According to Irandoust (2018), before the modern tax system which came into the picture in 1903, Sweden had a state income tax system based on “appropriations”. Some lumpsum taxes such as armament fees and personal protection fees were also levied. With the modern tax system, a number of tax reforms were undertaken, and the function of the tax

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system has changed over time. Irandoust (2018) gives an overview of how the Swedish tax system has evolved with regard to its functions. Before the modern tax system, tax system in Sweden had a pure fiscal function. In the 19th century, the allocative function was the main function of the Swedish tax system with a limited government intervention mainly for infrastructure development. In the 20th century, the distribution function received increased attention with more importance put on social welfare benefits, although in the 1950s and 1960s, stabilization became central. In the 1980s and 1990s, emphasis was once again put on allocative issues, while stabilization and distributive functions became less significant.

Fredrik and Lars (2019) highlight very well the fiscal history of Sweden from 1750 to 2017, which is summarized hereafter. Before the industrialization process, the public debt level remained relatively low and stable; it was around 10% from 1750 to 1788 and increased to 30% during the war against Russia (1788–1790). In the 1820s, public debt ratio reduced and was close to zero. This situation will prevail until the start of industrialization in the 1850s which increased public debt to 20% of GDP because of government's investments in infrastructure. For almost a century (from the 1880s until 1970), the debt ratio remained stable varying between 15% and 25% except during the Second World War where the debt ratio reached 50% but reduced quickly to the prewar level (20%) by 1950. The first and second world war had a brief effect of the government's borrowing since Sweden was not active in those two wars. With the introduction of the Bretton Woods system in the 1970s, the fiscal history of Sweden became volatile. From 12.5% in 1970, public debt ratio reached 62% in 1985 then fell to 40% in 1990 but rapidly increased to 74%i n 1995 because of the financial crisis of 1991–1993. The budget deficit in 1993 reached as high as 15% of GDP. The fiscal reforms that followed the financial crisis of the 1990s helped public debt to fall quickly and to be managed. Central government debt fell from 74% in 1995 to 33% in 2008 and 29% in 2017. Fiscal deficits have also been managed; from a fiscal surplus of 3.3% in 2007, Sweden experienced fiscal deficits but very low, 0.7%, 1.4%and 1.9% respectively for 2009, 2013 and 2014.

In response to the fiscal crisis in the 1990s, the Swedish fiscal framework was reformed but with the same goals of keeping public spending under control and ensuring that the national debt ratio declines over time (Fredrik and Lars, 2019). As part of the new fiscal framework, multiyear expenditure ceilings were introduced since 1996 to control long-term spending; a surplus target of 2% of GDP was also introduced in 1997 but was reduced to 1% in 2007 and further to 1/3 of a percent of GDP in 2016. The objective of the surplus target was to reduce government debt and prepare for an elderly population. (Fredrik and Lars, 2019). The surplus target is taken into account in setting the expenditure ceiling (Merrifield and Poulson, 2016). An independent fiscal institution (a fiscal policy council) was also created in 2007 with 3 functions, namely, monitoring compliance with fiscal rules or targets; macrofiscal evaluation and forecasting; assessment of long-term debt sustainability and formulating policy advice (Kopits, 2011). The fiscal policy council acts as a fiscal watchdog, to report deviations from the set expenditure ceilings and fiscal balance targets, to the Parliament. From the reform of 2016, a debt anchor was introduced, set at 35% of GDP ± 5 percentage points. Consequently, with the new fiscal framework, Sweden has been one of the best fiscal performers in OECD (Fredrik and Lars, 2019).

Several studies have examined the sustainability of fiscal deficits in the euro area countries. These include Bravo and Silvestre (2003), Afonso (2005), Claeys (2007), Mercan (2014), Afonso and Jalles (2015), and Brady and Magazzino (2019). Bravo and Silvestre (2003) examined the intertemporal sustainability of fiscal policy in 11 European countries. By means of Johansen cointegration tests, the results show that fiscal deficits were weakly sustainable only in Austria, France, The Netherlands, the UK and in Germany. Afonso (2005) studied the sustainability of fiscal deficits for the 15 EU countries for the 1970–2003 period by testing for cointegration between government spending and revenue and concluded that Swedish fiscal deficits were not sustainable. Claeys (2007) examined the sustainability of European fiscal policies over the period 1970–2001 by testing stationarity of total net lending (% GDP) and then cointegration among government expenditures, revenues and net interest payments. He concluded that Swedish fiscal deficits were sustainable. Arghyrou and Luintel (2007) examined government solvency for four Euro zone countries, namely Greece, Ireland, Italy and the Netherlands. Using cointegration tests accounting for breaks, they found that government finances of all four countries satisfied the intertemporal budget constraint across different time horizons. Afonso and Jalles (2015) analyzed fiscal sustainability for 19 countries over the period 1880–2009 by examining stationarity of the public debt ratio and found that fiscal deficits were sustainable in most countries including Sweden. Mercan (2014) investigated the sustainability of fiscal deficits for OECD countries using panel cointegration test with multiple structural breaks and concluded that budget deficits of OECD countries are sustainable in weak form. Brady and Magazzino (2019) examine the sustainability of Italian fiscal policy over the period 1862 to 2013. The study conducted the analysis over the entire period and on two subperiods 1862–1913 and 1947–2013 and concluded that deficits were only sustainable over the subperiod 1862–1913. Miyazaki (2014) examined the influence of fiscal rules adopted since the mid-1990s by Australia and Sweden, on the sustainability of their fiscal policy. Based on the estimation results of the long-run relationship between government spending and revenue, he concluded that the fiscal reform in Sweden has been beneficial for running a budget surplus, while for Australia, the reform has not been useful for ensuring the sustainability of fiscal policy.

Few studies have focused solely on Sweden; among them are Hatemi (2002a, b) and Irandoust (2018). Hatemi (2002a) investigated the sustainability of Swedish fiscal policy during the period 1963–2000 by paying special attention to the effect of European Monetary Union (EMU) criteria convergence. He divided the sample into two with the chosen date break being January 1990, corresponding to the period when the EMU criteria convergence came into effect. Using Johansen cointegration test, he found that spending and revenue were cointegrated for the period 1963:01–1989:04 and not for 1990:01–2000:01. However, for the whole sample, he found a cointegration relationship between the variables. Hatemi (2002a) concluded that Sweden is not in violation of its intertemporal budget constraint. Hatemi (2002b) analyzed whether Swedish government complied with its budget constraint for the period 1963–2000 using quarterly data. He found that spending and revenue were cointegrated. An estimation using state-space model revealed that the estimated time-varying coefficient remains close to one, providing evidence that the government fulfilled its budget constraint during the sample period. In a more recent paper, Irandoust (2018) examined the relationship between government spending and revenue for Sweden using hidden cointegration test and other causality tests. He found that there is cointegration between both positive and negative components of the variables. Based on the estimated slope coefficient, he concluded that the Swedish government follows a hard budget constraint strategy (strong deficit sustainability) for negative components of the variables and a soft budget constraint strategy (weak deficit sustainability) in the case of positive components of the variables.

However, most of these above reviewed studies do not account for breaks in the approaches used, and the same goes for the few studies focusing on Sweden. Yet as Fredrik and Lars (2019) point out, Sweden has experienced a number of structural breaks from various sources, internal and external, including the crisis of 1877/1878, the international financial crisis of 1907, the depression of the early 1920s, the Great Depression in the 1930s, the two world wars, oils shocks, the financial crisis in the early 1990s and the international crisis of 2008/2009. Sweden undertook also several tax reforms between 1980 and 1992 and other macroeconomic shocks occurred during that period (Hatemi, 2002b).

As Martins (2015) points out, structural change is of key importance in economics and econometrics, especially for cointegration analysis, involving long-term historical trends which are likely to display breaks in their equilibrium relationship. Similarly, according to Kejriwal (2008), the presence of unaccounted shifts in the long-run relationship biases the usual cointegration tests in favor of nonrejection of the null hypothesis of no cointegration. Accounting for breaks in the estimation of the level relationship between spending and revenue is also important because ignoring it may lead to an overestimation or underestimation of the cointegrating slope and hence bias the magnitude of fiscal deficit sustainability. In addition, assuming parameter constancy in a level relationship among variables might be inefficient because rational agents are expected to react to new conditions that are caused by policy changes Hatemi (2002a). In relation to that, in testing for cointegration between spending and revenue, assuming a constant cointegrating slope when it is actually time-varying can also be misleading because deficits can be sustainable for a period of time and unsustainable over another period.

The contribution of this study is therefore threefold. This study assesses the sustainability of the Swedish fiscal policy by using a long data series spanning for more than two centuries (1800–2011); by applying the methodology of Kejriwal and Perron (2010) to test for multiple structural changes in a cointegrated regression model of government spending and revenue and by applying time-varying cointegration test of Bierens and Martins (2010) and Martins (2015). In this study, these two approaches attempt to highlight subperiods in which fiscal policy has been sustainable in Sweden and in which it has been unsustainable. It should be noted that the approaches of Kejriwal and Perron (2010) and Bierens and Martins (2010) have been applied extensively in the empirical literature (see for example, Bajo-Rubio *et al.*, 2010; Gabriel and Martins, 2011; Dulger, 2016; Gogolin *et al.*, 2018; Esteve *et al.*, 2020).

The rest of this paper is structured as follows. Section 2 gives the conceptual framework for testing fiscal deficit sustainability. Section 3 presents the methodology. Section 4 gives the presentation and interpretation of results, and chapter 5 concludes the study.

2. Conceptual framework for testing fiscal deficit sustainability

The starting point of the framework for testing fiscal deficit sustainability as found in Quintos (1995) is the following one-period government's budget constraint, where B_t is government debt, R_t is government revenue and $G_t^r = G_t + r_t B_{t-1}$ is government spending inclusive of interest payments, with G_t being primary government expenditure and r_t is the real interest rate assumed to follow a stationary process with mean r.

$$\Delta B_t = G_t^r - R_t \tag{1}$$

With the above assumption on the interest rate, Quintos (1995) rewrites Eqn (1) as follows:

$$B_t - (1+r)B_{t-1} = E_t - R_t \tag{2}$$

where $E_t = G_t + (r_t - r)B_{t-1}$ is G_t^r when interest rates are around a zero mean.

Using forward substitution, he writes the present value of the government's borrowing constraint as:

$$B_t = \sum_{j=0}^{\infty} \frac{1}{(1+r_t)^{j+1}}(R_{t+j} - E_{t+j}) + \lim_{j\to\infty} \frac{B_{t+j}}{(1+r_t)^{j+1}} \tag{3}$$

He then transforms Eqn (3) in terms of first difference as follows:

$$\Delta B_t = \sum_{j=0}^{\infty} \frac{1}{(1+r_t)^{j+1}}(\Delta R_{t+j} - \Delta E_{t+j}) + \lim_{j\to\infty} \frac{\Delta B_{t+j}}{(1+r_t)^{j+1}} \tag{4}$$

Eqn (4) is also equivalent to

$$G_t^r - R_t = \sum_{j=0}^{\infty} \frac{1}{(1+r_t)^{j+1}} (\Delta R_{t+j} - \Delta E_{t+j}) + \lim_{j\to\infty} \frac{\Delta B_{t+j}}{(1+r_t)^{j+1}} \tag{5}$$

as $\Delta B_t = G_t^r - R_t$.

According to Quintos (1995), fiscal deficits are sustainable if the present value of the stock of public debt goes to zero in infinity, that is, $\lim_{j\to\infty} \frac{\Delta B_{t+j}}{(1+r_t)^{j+1}} = 0$, in this case, the public debt $\Delta B_t = G_t^r - R_t$, does not grow without limit.

If $\lim_{j\to\infty} \frac{\Delta B_{t+j}}{(1+r_t)^{j+1}} = 0$, testing for fiscal sustainability from Eqn (5), implies testing for the stationarity of the first difference of the government debt, ΔB_t or alternatively testing for the stationarity of $G_t^r - R_t$. This is the approach followed by Trehan and Walsh (1988, 1991). According to Hakkio and Rush (1991), this is equivalent to testing for cointegration between government spending (including interest payments) G_t^r and government revenue R_t, assuming that G_t^r and R_t are both nonstationary processes integrated of order one, with a [1, −1] cointegration vector, using the following long-run relationship:

$$R_t = \alpha + \beta G_t^r + \varepsilon_t$$

However, Quintos (1995) differentiates between strong sustainability and weak sustainability. If G_t^r and R_t are cointegrated with $\beta = 1$, fiscal deficits are strongly sustainable, and the government is said to follow a hard budget constraint strategy. But if $0 < \beta < 1$, fiscal deficits are only weakly sustainable, and the government is said to follow a soft budget constraint strategy. Fiscal deficits are said to be unsustainable if $\beta \leq 0$. However, it should be noted that with the weak form of fiscal deficits sustainability, government spending increases more than revenue. Therefore the government is not able to pay its debt in the long-run.

In the empirical literature, testing fiscal policy sustainability is also achieved by estimating a fiscal reaction function (Bohn, 1998, 2007). This study follows Hakkio and Rush (1991) and Quintos (1995) and test for cointegration between government spending and revenue for Sweden because of a long series data which is available, spanning over two centuries.

3. Methodology

Economies are often exposed to endogenous breaks from different sources (oil shocks, financial crisis, wars, etc.), affecting the path of time series variables or the relationship among them. Accounting for breaks when analyzing the long-run relationship between spending and revenue for Sweden is important given several breaks that the country experienced through the years (Fredrik and Lars, 2019). According to Hatemi (2008), the ADF and PP test statistics on the residuals series, suggested by Engle and Granger (1987) are misspecified in the presence of structural breaks. Hence the importance of cointegration tests that account for breaks. In this study, we first use Gregory and Hansen (1996a, b) cointegration testing approach accounting for one break, but as Kejriwal (2008) notes, the test of Gregory and Hansen (1996a, b) may have low power in case of multiple breaks. Therefore, we also apply the methodology of Kejriwal and Perron (2010) to test for multiple structural changes in a cointegrated regression model. Following Kejriwal (2008), we proceed first by testing for the number of breaks in the level relationship between spending and revenue, then test for cointegration using the number of breaks obtained and finally estimating the long-run relationship between the variables with the number of breaks selected.

Because assuming a constant cointegrating slope could be misleading in case of structural breaks, we also apply time-varying cointegration test of Bierens and Martins (2010), and Martins (2015). Martins (2015) extends Bierens and Martins (2010) to develop bootstrap tests for time-varying cointegration. According to Martins (2015), the test proposed by Bierens and Martins (2010) falsely indicates the existence of time-varying cointegration too often.

3.1 Testing approach for multiple structural changes in a cointegrated regression model

3.1.1 Stability test and selecting the number of breaks in the long-run relationship. To test for stability in a long-run relationship, Kejriwal and Perron (2010) suggest the two following tests, where SSR_0 is the sum of squared residuals under the null hypothesis of no breaks and SSR_k, the sum of squared residuals under the alternative hypothesis of k breaks.

$$\sup F_T^*(k) = \sup_{\lambda\epsilon\Lambda_\epsilon} \frac{\mathrm{SSR}_0 - \mathrm{SSR}_k}{\hat{\sigma}^2}$$

$$\mathrm{UD}\ \max F_T^*(M) = \max_{1\le k\le m} F_T^*(k)$$

where $\lambda = \{\lambda_1, \lambda_2, \ldots, \lambda_m\}$ is the vector of break fractions defined by $\lambda_i = T_i/T$, with T_i the break date and $\hat{\sigma}^2 = T^{-1}\sum_{t=1}^{T}\tilde{u}_t^2 + 2T^{-1}\sum_{j=1}^{T-1} w(j/\hat{h})\sum_{t=j+1}^{T}\tilde{u}_t\tilde{u}_{t-j}\tilde{u}_t(t = 1, \ldots, T)$ are the residuals from the model estimated under the null hypothesis of no structural change.

To determine the number of structural breaks in a cointegrating relationship, Kejriwal and Perron (2010) consider a sequential test with the null hypothesis of k breaks against the alternative of $k + 1$ breaks, given by:

$$\mathrm{SEQ}_T(k+1|k) = \max_{1\le j\le k+1} \mathrm{sub}_{\tau\epsilon\Lambda_{j,\epsilon}} T\{\mathrm{SSR}_T(\hat{T}_1, \ldots, \hat{T}_k) - \mathrm{SSR}_T(\hat{T}_1, \ldots, \hat{T}_{j-1}, \tau, \hat{T}_k)\}/\mathrm{SSR}_{k+1},$$

where

$$\Lambda_{j,\epsilon} = \{\tau; \hat{T}_{j-1} + (\hat{T}_j - \hat{T}_{j-1})\varepsilon \le \tau \le \hat{T}_j - (\hat{T}_j - \hat{T}_{j-1})\varepsilon\}$$

The procedure goes like this: in the first step, the hypothesis of zero versus one break is tested; if rejected, the hypothesis of one versus two breaks is then tested and so on. The procedure stops at the point where the null hypothesis is accepted. The number of breaks is then given by the number of rejections. In addition to the sequential procedure, Kejriwal and Perron (2010) use also two information criteria, Bayesian information criterion (BIC) and another criterion suggested by Liu *et al.* (1997), denoted by LWZ.

$$\mathrm{BIC}(m) = \ln\hat{\sigma}^2(m) + p^*\ln(T)/T,$$

where $p^* = (m+1)q + m + p$, and $\hat{\sigma}^2(m) = T^{-1}S_T(\hat{T}_1, \ldots, \hat{T}_m)$ $\hat{T}_1, \ldots, \hat{T}_m$ are the estimated break dates, $S_T(\hat{T}_1, \ldots, \hat{T}_m)$, is the sum of squared residuals under m breaks, q is the number of coefficients that are allowed to change and p is the number of coefficients that are held fixed.

$$\mathrm{LWZ}(m) = \ln\left(\frac{S_T(\hat{T}_1, \ldots, \hat{T}_m)}{T - p^*}\right) + \left(\frac{p^*}{T}\right)c_0(\ln(T))^{2+\delta_0}$$

Liu *et al.* (1997) suggest using $\delta_0 = 0.1$ and $c_0 = 0.299$.

3.1.2 Cointegration test with multiple breaks. Kejriwal (2008) points out that the test of Gregory and Hansen (1996a), which tests the null hypothesis of no cointegration against the alternative of cointegration in the presence of a possible regime shift, may have low power when the alternative hypothesis involves multiple breaks. To overcome that weakness, Arai and Kurozumi (2007) developed a test with the null hypothesis of cointegration but with only a single break. Kejriwal (2008) built on Arai and Kurozumi (2007) to develop a cointegration test with multiple breaks under the null hypothesis as the latter test may tend to reject the null hypothesis of cointegration in case of multiple breaks (Kejriwal, 2008).

The test statistic is obtained by first estimating the following dynamic OLS model including the leads and lags of the first differences of the *I(1)* regressors.

$$y_t = c_i + z_t'\beta_i + \sum_{j=-l_T}^{l_T} \Delta z_{t-j}'\Pi_j + u_t^*$$

The test statistic is then given by $\tilde{V}_1(\hat{\lambda}) = (T^{-2}\sum_{t=1}^{T} S_t(\hat{\lambda})^2)/\hat{\Omega}_{11}$, where $\hat{\Omega}_{11}$ is a consistent estimate of the long-run variance of u_t^*, and $\hat{\lambda} = \hat{T}_1/T, \ldots, \hat{T}_k/T$ is the vector of the break fractions, with break dates $(\hat{T}_1, \ldots, \hat{T}_k)$ are obtained by minimizing the sum of squared residuals.

Following Kejriwal (2008), to avoid the endogeneity problem in the estimation of the cointegrating equation, dynamic OLS (DOLS) estimation approach is used with 2 lags and 2 leads.

3.2 Time-varying cointegration test

The justification given by Martins (2015) for time-varying cointegration testing is to be able to account for structural change in cointegration analysis as variables are likely to display breaks in their equilibrium relationship especially for long-term historical trends. Bierens and Martins (2010) proposed the following time-varying *VECM (p)* model with a drift, in which the cointegration vectors change smoothly over time.

$$\Delta Y_t = \mu + \alpha\beta_t' Y_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta Y_{t-j} + \varepsilon_t,\ t = 1, \ldots, T,$$

where the β_t's are time-varying $k \times r$ matrices of cointegrating vectors, while the rest of the coefficients are fixed. The null hypothesis is the standard time-invariant (TI) cointegration, that is, $Ho:\ \beta_t = \beta$ for all t, against time-varying cointegration (TVC). It is assumed that the time-varying cointegrating vector varies smoothly over time, following this function:

$$\beta_t = m_t\left(\frac{t}{T}\right) = \sum_{i=0}^{m} \xi_{i,T} P_{i,T}(t),$$

where $P_{i,T}(t)$ are the orthonormal Chebyshev time polynomials, defined as $P_{i,T}(t) = \sqrt{2}\cos\left(\frac{i\pi(t-0.5)}{T}\right)$, with $P_{0,T}(t) = 1,\ \ t = 1, 2, \ldots, T,\ \ i = 1, 2, 3, \ldots, m,$ and the Fourier coefficients, $\xi_{i,T} = \frac{1}{T}\sum_{t=1}^{T} \beta_t P_{i,T}(t)$ are unknown $k \times r$ matrices and m is the order of the Chebyshev time polynomials.

Eqn (1) can be rewritten as $\Delta Y_t = \alpha\xi'\ Y_{t-1}^{(m)} + YX_t + \varepsilon_t$, with $Y_{t-1}^{(m)} = (Y_{t-1}',\ P_{1,T}(t)\ Y_{t-1}', P_{2,T}(t)\ Y_{t-1}', \ldots,\ P_{m,T}(t)\ Y_{t-1}')',\ Y = (\mu,\ \Gamma_1 \Gamma_2, \ldots \Gamma_{p-1}), X_t = (1,\ \Delta Y_{t-1}', \Delta Y_{t-2}', \ldots, \Delta Y_{t-p+1}')$.

The LR test statistics is given as: $LR_{m,T}^{tvc} = T\sum_{j=1}^{r} \ln\left(\frac{1-\hat{\lambda}_{0,j}}{1-\hat{\lambda}_{m,j}}\right)$, where $1 > \hat{\lambda}_{m,1} \geq \hat{\lambda}_{m,2} \geq \ldots \geq \hat{\lambda}_{m,r} \geq \ldots \geq \hat{\lambda}_{m,(m+1)k}$ are the ordered solutions of the following generalized eigenvalue problem:$\det[\lambda S_{11,T}^{(m)} - \lambda S_{10,T}^{(m)} \lambda S_{00,T}^{-1} \lambda S_{01,T}^{(m)}] = 0$, with

$$S_{00,T} = \frac{1}{T}\sum_{t=1}^{T}\Delta Y_t \Delta Y_t' - \left(\frac{1}{T}\sum_{t=1}^{T}\Delta Y_t X_t'\right)\left(\frac{1}{T}\sum_{t=1}^{T}X_t X_t'\right)^{-1}\left(\frac{1}{T}\sum_{t=1}^{T}X_t \Delta Y_t'\right)$$

$$S_{11,T}^{(m)} = \frac{1}{T}\sum_{t=1}^{T}Y_{t-1}^{(m)} Y_{t-1}^{(m)'} - \left(\frac{1}{T}\sum_{t=1}^{T}Y_{t-1}^{(m)} X_t'\right)\left(\frac{1}{T}\sum_{t=1}^{T}X_t X_t'\right)^{-1}\left(\frac{1}{T}\sum_{t=1}^{T}X_t Y_{t-1}^{(m)'}\right)$$

$$S_{01,T}^{(m)} = \frac{1}{T}\sum_{t=1}^{T}\Delta Y_t Y_{t-1}^{(m)'} - \left(\frac{1}{T}\sum_{t=1}^{T}\Delta Y_t X_t'\right)\left(\frac{1}{T}\sum_{t=1}^{T}X_t X_t'\right)^{-1}\left(\frac{1}{T}\sum_{t=1}^{T}X_t Y_{t-1}^{(m)'}\right)$$

$$S_{10,T}^{(m)} = \left(S_{01,T}^{(m)}\right)'$$

For $m \geq 1$, and $r \geq 1$, Bierens and Martins (2010) show that the test statistic $LR_{m,T}^{tvc}$ under the null hypothesis of standard cointegration, follows a χ^2_{mkr} distribution.

For small T and large m, Bierens and Martins (2010) show that the $LR_{m,T}^{tvc}$ test suffers from size distortions and tends to overreject the correct null hypothesis of standard cointegration. Therefore, Martins (2015) introduced two bootstrap versions [1] of $LR_{m,T}^{tvc}$ test statistic, namely the wild bootstrap and the *i.i.d.* bootstrap.

4. Presentation and interpretation of results

This study uses annual data on government spending and revenue for Sweden spanning from 1800 to 2011. The data is from the historical public finance dataset of the International Monetary Fund (IMF) (see, Mauro *et al.*, 2013). Due to internal / external political and economic factors, structural breaks do occur and affect the time path of variables. We first therefore test for the presence of structural breaks in the series using Bai and Perron (2003) approach which suggests three tests, namely, *SupF* test, *AveF* test and *ExpF* test, with a null hypothesis of no structural change against an alternative hypothesis of arbitrary number of changes. The test results reported in Table 1 shows that the null hypothesis of no structural break in the series, government spending (% GDP) and government revenue (% GDP), is strongly rejected, at 1% level, by all three tests (*SupF*, *AveF* and *ExpF*). We next test the null hypothesis of l changes against the alternative of $l+1$ changes. The results indicate the presence of four structural breaks in both series; 1914, 1939, 1965 and 1990 for government revenue and 1914, 1939, 1969 and 1990 for government spending. These breaks can be explained by a number of events which occurred during the study period which affected the public finances of Sweden, including the First and Second World War for 1914 and 1939 and other events.

To examine the characteristics of the variables, a standard unit root test suggested by Ng and Perron (2001) is used, as well as unit root tests with breaks accounting for one break suggested by Zivot and Andrews (1992), and Lee and Strazicich (2003) test and a test

Table 1. Bai and Perron (2003) Test for the presence of structural breaks in the series

	Tests statistics			
	Sup-*F*	Ave-*F*	Exp-*F*	Detected breaks
Revenue (% GDP)	715.6***	254.69***	353.08***	1914, 1939, 1965, 1990
Spending (% GDP)	776.88 ***	277.17***	384.37***	1914, 1939, 1969, 1990

Note(s): The tests are done in *R* software using an *R* package called "strucchange". *** denotes rejection of the null hypothesis at 1% level of significance accounting for two breaks suggested by Clemente Lopez *et al.* (1998). Indeed, accounting for breaks in unit root testing is important; as Baum (2001) points out, if breaks are not accounted for in testing for unit root, there might be a confusion of structural breaks in the series as evidence of nonstationarity. And as Kejriwal (2008) says, the presence of unaccounted shifts in the long-run relationship biases the usual cointegration tests in favor of nonrejection of the null hypothesis of no cointegration. In this study, the standard unit root test used was suggested by Ng and Perron (2001). It is an extension of the *M* tests developed in Perron and Ng (1996) to allow for GLS detrending of the data. It suggests also a modified information criterion (MIC) for lag selection in the ADF regression that reduces considerably the size and power distortions. Four test statistics are developed by Ng and Perron (2001), namely, MZ_a, MZ_t, MSB and MPT.

The results for unit root tests are presented in Table 2. All the tests used, standard and those accounting for breaks, show that government spending and revenue in Sweden are nonstationary processes becoming stationary after one differentiation. A level relationship between them can therefore be tested. For both variables, government spending and revenue, Zivot and Andrews (1992) test detects the break in 1975, while Lee and Strazicich (2003) test detects the break in 1993. For government spending, the breaks dates suggested by Clemente Lopez *et al.* (1998) test are 1938 and 1975, as well as 1937 and 1978, respectively for innovational outlier (IO) and additive outlier (AO) models. For government revenue, the breaks are given respectively by the two models as 1950 and 1991 and 1945 and 1990. Again these breaks can be justified by a number of events that occurred and disturbed the normal time path of the variables, including the great depression of the 1930s, the Second World War,

Tests	Spending (% GDP) Level	Spending (% GDP) 1st difference	Revenue (% GDP) Level	Revenue (% GDP) 1st difference
Standard Unit root tests				
Ng and Perron (2001)				
MZ_a	−6.549	−102.46***	−4.297	−104.58***
MZ_t	−1.746	−7.138 ***	−1.346	−7.197***
MSB	0.266	0.069***	0.313	0.069 ***
MPT	13.949	0.275***	20.147	0.296***
Unit root tests with one break				
Zivot and Andrews (1992)				
Innovational outlier	−4.064 (0) [1975]	−18.875*** (0)	−3.495 (0) [1975]	−16.916*** (0)
Additive outlier	−4.087 (0) [1975]	−18.695*** (0)	−3.513 (0) [1975]	−16.916*** (0)
Lee and Strazicich (2003)				
Impulse dummy	−1.977 (0) [1993]	−14.044*** (1)	−1.560 (0) [1993]	−18.767*** (0)
Shift dummy	−1.912 (0) [1993]	−2.631* (0)	−1.810 (0) [1993]	−2.983** (0)
Exponential shift	−1.907 (0) [1993]	−2.388 (0)	−1.815 (0) [1993]	−2.724*(0)
Rational shift	0.790 (0) [1993]	−8.273***(0)	0.790 (0) [1993]	−8.490***(0)
Unit root tests with two breaks				
Clemente Lopez et al. (1998)				
Innovational outlier	−4.691 (7) [1938, 1975]	−10.621*** (6)	−5.557** (0) [1950, 1991]	–
Additive outlier	−3.979 (7) [1937, 1978]	−4.856* (12)	−4.486 (5) [1945, 1990]	−2.273 (2)

Note(s): In parenthesis, are lags used for the tests, selected automatically using Schwarz information criterion from a maximum lags of 14. In brackets, are the years of break. *, ** and *** denote rejection of the null hypothesis at 10%, 5% and 1% level, respectively

Table 2.
Unit root tests

the international economic crisis of the 1970s and a banking crisis in the early 1990s in Sweden. Indeed, there was in Sweden a banking crisis from 1991 to 1993, which caused a deep economic recession and a massive increase in unemployment as well as a rapidly growing budget deficit. The crisis halted a welfare-state expansion that had been going on for decades (Bergmark and Palme, 2003).

Because we have confirmed that our variables, government spending and revenue are nonstationary, *I(1)* processes, we can test for cointegration relationship between the variables. We first present the Gregory-Hansen (1996a, b) cointegration test allowing only one break. The results in Table 3 show that for all the models, the null hypothesis of no cointegration is strongly rejected, in favor of the presence of cointegration with one break. However, as discussed in the methodology, because multiple breaks is a possibility which could bias the cointegration test, we proceed next by following Kejriwal and Perron (2010) and test for multiple structural changes in a cointegrated regression model. A 15% trimming is used, and a maximum number of breaks equal to 5.

The results in Table 4 (panel A) indicate that none of the tests is significant, suggesting a stable cointegrating relationship between government spending and revenue in Sweden. The sequential procedure selects no break while the information criteria, BIC and LWZ, both select 2 breaks, in 1944 and 1975.

Next, we test for cointegration test between government spending and revenue using the tests corresponding to the number of breaks detected by sequential procedure and information criteria. Because the sequential procedure selects no break, we use the usual no-break cointegration test of Engle-Granger (1987) and Phillips and Ouliaris (1988). The results reported in Table 4 (panel B) indicate that both tests point to a cointegration relationship between spending and revenue in Sweden irrespective of which variable is endogenized between the two. However, because the information criteria (BIC and LWZ) select instead two breaks, we also use the cointegration test with multiple breaks of Kejriwal (2008). The results in Table 4 (panel C) indicate that the null hypothesis of no cointegration is rejected at 10% level, in favor of a cointegration with two breaks.

Next, we compare the cointegration coefficients obtained from a model assuming no break and a model with two breaks affecting both the intercept and the slope. In the model with no break, we estimate a cointegrated regression equation with DOLS, allowing two leads and two lags; the estimated slope coefficient is found to be 0.939, statistically significant at 1% level (see, Table 4, panel B) and the intercept estimate is not significant. The Wald test rejects the hypothesis that the slope coefficient is equal to one. The cointegrated two-break model is also estimated using DOLS approach with two leads and two lags. Estimation results reported in Table 4 (panel C) indicate that the slope coefficient is statistically significant (at 5% level) in the first two regimes. The slope coefficient increases from 0.678 to 0.892 from the first to the second regime. The results imply that fiscal deficits were weakly sustainable in the first two regimes, from 1800 to 1943 and from 1944 to 1974. The estimated coefficient in the two regimes is smaller compared to that obtained from the model without break. This shows that ignoring breaks in the estimation of the long-run relationship between spending and revenue may lead to overestimating the slope coefficient. In the third regime, the slope coefficient is equal to 1.26 but is not statistically significant. It should be noted that from the 1990s until now, with the new fiscal framework and the introduction of a surplus target, fiscal performance has been quite good in Sweden, running fiscal surpluses for most of the years.

4.1 Time-varying cointegration test

In the time-varying cointegration test, the cointegration rank, r, is fixed and assumed to be known, obtained using the Johansen procedure (Martins, 2015). For this study, the cointegration rank is found to be one ($r = 1$) [3], from a VAR model of order 1 as determined by Hannan–Quinn (HQ) information criterion.

Table 3.
Gregory-Hansen (1996a, b)
Cointegration test with one break

Model	ADF	Break	z_t^*	z_a^*	Break	Lag
Panel A: test statistics						
C	−5.33***	1912	−5.36***	−50.13***	1915	4
C/T	−5.26**	1912	−5.34**	−49.86**	1937	4
C/S	−4.99**	1912	−5.62***	−52.65**	1977	4
C/S/T	−6.05***	1977	−6.37***	−70.40***	1978	4

	Model C			Model C/T			Model C/S			Model C/S/T		
	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%
Panel B: asymptotic critical values												
ADF*	−5.13	−4.61	−4.34	−5.45	−4.99	−4.72	−5.47	−4.95	−4.68	−6.02	−5.50	−5.24
z_t^*	−5.13	−4.61	−4.34	−5.45	−4.99	−4.72	−5.47	−4.95	−4.68	−6.02	−5.50	−5.24
z_a^*	−50.0	−40.48	−36.19	−57.2	−47.9	−43.22	−57.1	−47.0	−41.8	−69.3	−58.5	−53.3

Note(s): Gregory-Hansen (1996a, b) test was performed using “ghansen”, a STATA module available in the statistical software components archive. The lag length was selected using Bayesian information criterion out of a maximum lag of 12. Break dates are selected automatically by the software. *, ** and *** denote rejection of the null hypothesis at 10%, 5% and 1% level, respectively

	Sup F^* (1)	Sup F^* (2)	Sup F^* (3)	Sup F^* (4)	Sup F^* (5)	UDMax	(S)	(B)	(L)
Panel A: structural break tests in the long-run relationship									
Value	8.628	7.142	5.050	3.958	3.168	8.628	0	2	2
10% CV	10.34	8.85	7.66	6.66	5.30	10.53			
5% CV	12.11	9.96	8.60	7.36	5.90	12.25			
1% CV	17.03	12.41	10.40	8.71	7.08	17.40			

	Engle–Granger Test		Phillips and Ouliaris Test	
Dependent	$\tau -$ Statistic	$z -$ Statistic	$\tau -$ Statistic	$z -$ Statistic
Panel B: cointegration tests without breaks				
Revenue	−4.963*** (0.000)	−45.250*** (0.000)	−5.206*** (0.000)	−49.996*** (0.000)
Spending	−5.053*** (0.000)	−46.220*** (0.000)	−5.313*** (0.000)	−51.425*** (0.000)

DOLS Estimated long-run cointegrating equation without breaks (dependent: revenue)

Intercept	Slope (β)	$H0 : \beta = 1$
−0.267 (0.672)	0.939*** (0.000)	F-Stat = 8.789 (0.003)

Test statistic: 0.081*

Panel C: Kejriwal (2008) cointegration test with multiple breaks

Critical values

10%	5 %	1%
0.076	0.096	0.158

Estimated regression under Breaks

	c_1	c_2	c_3	δ_1	δ_2	δ_3	$\widehat{T}_1$	$\widehat{T}_2$
Parameter estimates	1.821**	1.423	−16.758***	0.678**	0.892**	1.260	1944	1975
Std errors	0.471	1.018	1.018	0.106	0.192	0.111		

Note(s): Results for structural break tests are obtained using a GAUSS code [2] written by Kejriwal (2008). S, B, and L are number of breaks selected by the sequential procedure, BIC and LWZ criteria, respectively. The critical values (CV) for the break tests are from Kejriwal and Perron (2010). The cointegrating equation is estimated using DOLS approach with two leads and two lags. Between brackets are the p-values. *, **, *** denote rejection of the null hypothesis at 10%, 5% and 1%, respectively. Critical values for Kejriwal (2008) test are obtained by simulation using the GAUSS code provided by Kejriwal (2008). $c_1, c_2,$ and c_3 are the estimated intercepts in the three regimes, while. δ_1, δ_2, and δ_3 are the estimated slope coefficients in the three regimes

Table 4. Testing for multiple structural changes in a cointegrated regression model

According to Martins (2015), the value of m to consider is given by T/10. Because the number of observations, T is 212 in this study, m ranges therefore from 1 to 21. The tests results for the time-varying cointegration are reported in Tables 5 and 6, for m ranging from 1 to 21. The results show that both the asymptotic test and bootstrap tests reject the null hypothesis of standard time-invariant (TI) cointegration, for all values of m. The asymptotic test rejects the null hypothesis at 1% level for all values of m. The bootstrap tests (Wild Boostrap [4] and Swensen's i.i.d) reject also the null hypothesis of time-invariant cointegration for all values of m, except for m equal to 3 for the Wild test. Indeed, the optimal value of m is equal to 7 where the value of HQ criterion is minimized.

The results support time-varying cointegration between government spending and revenue in Sweden. The cointegrating vector is therefore also time-varying.

Given the vector of government spending and revenue used in the study, $Y_t = (R_t, G_t)'$, the short-run budgetary disequilibrium is given by $\beta_t' Y_t = \varepsilon_t$, that is, $\beta_{1t}\ R_t + \beta_{2t}\ G_t = \varepsilon_t$. The plot of the estimated cointegrating vector $(\beta_{1t}, \beta_{2t})'$ and the corresponding normalized vector $(1, \ -\delta_{2t})'$, where $\delta_{2t} = -\beta_{2t}/\beta_{1t}$, is in Figure 1. For the strong sustainability hypothesis, the cointegrating vector is expected to be $(1, -1)'$. Figure 1 shows that the

m	HQ	LR TVC	p-value (asympt)	5% CV (asympt)	p-value (boots)	5% CV (boots)
1	8.7402	16.765	0.000	5.991	0.061	18.284
2	8.7531	20.751	0.000	9.487	0.082	26.906
3	8.7817	21.459	0.001	12.591	0.152	36.503
4	8.7306	38.881	0.000	15.507	0.081	47.162
5	8.7557	40.326	0.000	18.307	0.049	40.139
6	8.7556	47.056	0.000	21.026	0.000	26.808
7	8.6898	67.587	0.000	23.684	0.000	31.165
8	8.7083	70.400	0.000	26.296	0.000	35.601
9	8.7275	73.067	0.000	28.869	0.000	40.878
10	8.7535	74.315	0.000	31.410	0.000	44.386
11	8.7271	86.577	0.000	33.924	0.000	47.429
12	8.7581	86.775	0.000	36.415	0.000	51.761
13	8.7804	88.787	0.000	38.885	0.000	56.113
14	8.7953	92.370	0.000	41.337	0.001	60.986
15	8.8226	93.335	0.000	43.772	0.001	64.559
16	8.8484	94.621	0.000	46.194	0.003	72.450
17	8.8717	96.443	0.000	48.602	0.002	75.434
18	8.8945	98.349	0.000	50.998	0.002	79.029
19	8.9175	100.241	0.000	53.383	0.007	86.228
20	8.9250	105.359	0.000	55.758	0.013	92.770
21	8.9437	108.147	0.000	58.124	0.029	101.386

Note(s): A GAUSS code written by Martins (2015) is used to obtain the results. 1999 bootstraps are used to generate the bootstrap 5% critical values and p-values

Table 5.
The Wild bootstrap test of time-varying cointegration

m	HQ	LR TVC	p-value (asympt)	5% CV (asympt)	p-value (boots)	5% CV (boots)
1	8.7402	16.765	0.000	5.991	0.016	8.960
2	8.7531	20.751	0.000	9.487	0.013	11.179
3	8.7817	21.459	0.000	12.591	0.022	15.432
4	8.7306	38.881	0.000	15.507	0.007	20.320
5	8.7557	40.326	0.000	18.307	0.011	22.313
6	8.7556	47.056	0.000	21.026	0.005	25.840
7	8.6898	67.587	0.000	23.684	0.004	32.931
8	8.7083	70.400	0.000	26.296	0.004	30.802
9	8.7275	73.067	0.000	28.869	0.007	39.766
10	8.7535	74.315	0.000	31.410	0.009	45.956
11	8.7271	86.577	0.000	33.924	0.005	44.176
12	8.7581	86.775	0.000	36.415	0.011	53.201
13	8.7804	88.787	0.000	38.885	0.018	61.881
14	8.7953	92.370	0.000	41.337	0.011	61.442
15	8.8226	93.335	0.000	43.772	0.021	72.972
16	8.8484	94.621	0.000	46.194	0.030	82.577
17	8.8717	96.443	0.000	48.602	0.022	77.397
18	8.8945	98.349	0.000	50.998	0.045	95.068
19	8.9175	100.241	0.000	53.383	0.060	106.244
20	8.9250	105.359	0.000	55.758	0.029	95.430
21	8.9437	108.147	0.000	58.124	0.056	110.379

Note(s): A GAUSS code written by Martins (2015) is used to obtain the results. 1999 boostraps are used to generate the bootstrap 5% critical values and p-values

Table 6.
The Swensen's i.i.d. bootstrap test of time-varying cointegration

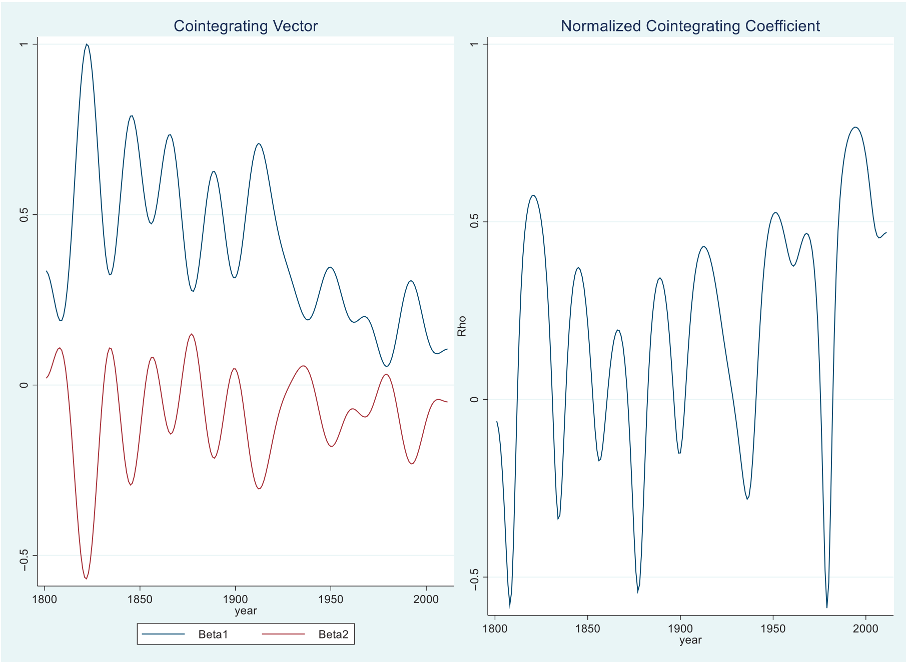

Figure 1. Time-varying cointegrating coefficients

Note(s): Time-varying cointegrating coefficients are estimated from EasyReg software. Beta1 and Beta2 is the cointegrating vector, for government revenue and spending, respectively

normalized cointegrating slope coefficient is negative for the periods 1801–1811, 1831–1838, 1853–1860, 1872–1882, 1897–1902, 1929–1940 and 1976–1982, implying that fiscal deficits in Sweden were unsustainable over those periods of time. Some events can explain the unsustainability of deficits during those periods, including the great depression of the 1930s, oil shocks in the 1970s, etc.

Fiscal deficits were weakly sustainable over the rest of time periods, where the normalized cointegrating slope coefficient is positive but less than 1, that is, 1812–1830 with an average cointegrating slope coefficient of 0.412; 1839–1852 with an average of 0.263; 1861–1871 with an average of 0.133; 1883–1896 with an average of 0.235; 1903–1928 with an average of 0.269; 1941–1975 with an average of 0.394 and 1983–2011 with an average of 0.597. The average cointegrating slope coefficient is bigger for the period 1983–2011, in which a new fiscal framework was introduced since 1997.

5. Conclusion

This study sought to re-examine the sustainability of Swedish fiscal policy using annual data spanning over a period of two centuries. Two approaches were used; the methodology of Kejriwal and Perron (2010) to test for multiple structural changes in a cointegrated regression model and time-varying cointegration test of Bierens and Martins (2010) and Martins (2015). Using the first approach, the results indicated that government spending and revenue are cointegrated with two breaks. An estimation of a two-break long-run model showed that the slope coefficient increases from 0.678 to 0.892 from the first to the second regime, implying that fiscal deficits were weakly sustainable in the first two regimes, from 1800 to 1943 and from 1944 to 1974. Further, results from time-varying cointegration test indicate that cointegration between spending and revenue is time-varying. Fiscal deficits were found to be unsustainable for the periods 1801–1811, 1831–1838, 1853–1860 , 1872–1882, 1897–1902, 1929–1940 and 1976–1982 and weakly sustainable over the rest of the study period. The findings in this study imply that accounting for breaks in cointegration analysis and in the estimation of the level relationship between spending and revenue is very important because ignoring breaks may lead to an overestimated slope coefficient and hence a bias on the magnitude of fiscal deficit sustainability. In addition, in testing for cointegration between spending and revenue, assuming a constant cointegrating slope when it is actually time-varying may also be misleading because deficits can be sustainable for a period of time and unsustainable over another period. The approaches used in this study allowed to highlight subperiods in which fiscal policy has been sustainable in Sweden and in which it has been unsustainable. In that, this study differs from Hatemi (2002a, b), and Irandoust (2018).

Notes

1. The details of these procedures can be found in Martins (2015).
2. The author would like to thank Kejriwal (2008) for making available the GAUSS code for testing multiple breaks in cointegrated regression models.
3. For space requirement, the results are not presented but are available upon request.
4. For space requirement, the results from the Wild bootstrap test are not presented but are available upon request.

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Corresponding author
Arcade Ndoricimpa can be contacted at: arcade_ndoricimpa@yahoo.com