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Chronic Fiscal Deficits and Chronic Inflation in the Burundian Economy: An Empirical Test of the Validity of the Olivera–Tanzi and Patinkin Effects Using ARDL Models

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Abstract

For nearly two decades, the Burundian economy has faced challenges in public finance management. On the one hand, public spending has increased dramatically, and revenues are no longer sufficient to cover it. As a result, the economy is experiencing recurring budget deficits. On the other hand, inflation continues its upward trend, becoming increas-ingly chronic and pervasive, steadily eroding the purchasing power of the average citizen. Although this situation is underappreciated, the coexistence of twin problems—chronic deficits and chronic inflation—raises fundamental ques-tions about the impact of inflation on public finances, particularly with respect to the theoretical "Olivera-Tanzi" and "Patinkin" effects. Therefore, this study aims to empirically analyze the Olivera-Tanzi and Patinkin effects in the Burundian economy. To achieve this objective, the study uses real tax revenues and real public expenditures as de-pendent variables. Inflation is the main independent variable, whereas real GDP, the money supply, and the real inte-rest rate are control variables that may influence the relationship between the main independent variable and the de-pendent variable. The data are quarterly, spanning 2004:Q1 to 2024:Q2. The ADF and PP unit root tests showed that the models' series were integrated of order I(0) and I(1), respectively. This indicates that the ARDL approach and the bounds test for cointegration should be employed. Additionally, after applying the CUSUM and CUSUMSQ stability tests to both models, it was found that the real tax revenues (the Olivera-Tanzi effect) model exhibited potential breaks, particularly from the second quarter of 2015 onward. This break corresponds to the political violence that plagued the country from the beginning of that quarter. Accordingly, this paper introduced a dummy variable (PV = Political Violence) into the ARDL model of real tax revenues. The Olivera–Tanzi effect model results indicate that, in the short run, a one–unit increase in the contemporaneous inflation rate leads to a 0.68% decrease in real tax revenues, and a one–unit increase in the past inflation rate leads to a 0.54% decrease in real tax revenues. Hence, in the short run in Burundi, the Olivera–Tanzi effect holds. In the long run, a 1% change in real GDP leads to a 2.80% change in real tax revenues. In the long run, the presence of political violence resulted in a 23.4% decrease in real tax revenues rela-tive to the period without political violence. According to the Patinkin effect model, contemporaneous inflation does not significantly affect real public expenditures. In contrast, a 1% increase in the money supply is associated with a 1.91% decrease in real public expenditures, as observed in the second quarter. In the short run, a 1% increase in real GDP leads to a 1.33% increase in real public expenditures. However, this effect appears to change over subsequent quarters. Furthermore, in the long run, a 1% increase in real GDP is associated with a 2.39% increase in real public expenditures. Conversely, a 1% point increase in inflation leads to a 1% increase in real public expenditures. This result, therefore, reinforces the Olivera–Tanzi effect rather than the Patinkin effect. Thus, in Burundi, the Patinkin effect postulate does not hold.

Keywords: Chronic Fiscal Deficits, Chronic Inflation, Burundian Economy, Olivera–Tanzi effect, Patinkin Effect, ARDL Models

1. Introduction

Over time, high inflation has become a chronic problem for Burundi's economy (Kimolo et al., 2024). By associating it with the erosion of the local currency's purchasing power (UNICEF, 2024), public opinion has already begun to characterize it as an economic necrosis that is paralyzing the economy. Persistent high inflation puts the country into turmoil by eroding the public's confidence in the national currency. This situation demonstrates the extent to which the price increase has significantly and negatively affected households' daily lives (Kimolo et al., 2024; UNICEF, 2023).

The persistence of chronic budget deficits in Burundi has also led to macroeconomic vulnerability, increasing depen-dence on monetary financing, and fueling inflationary pressures (Mvuyekure et al., 2025). These trends create an inflationary spiral in which the state lacks sufficient resources to finance recurring deficits and therefore resorts to creating money (Luis and Marco, 2003). This situation fuels doubts about the credibility of fiscal and monetary poli-cies and undermines the institutional capacity. Therefore, limited fiscal space and severe inflationary pressures leave institutions with only a few instruments for stabilizing the economy (Kimolo et al., 2024).

Empirically, studies have shown that inflation reduces the real value of fiscal revenues, a phenomenon known as the "Olivera–Tanzi effect" (Tanzi et al., 1987; Anušić and Švaljek, 1996; Nzirorera, 2016). They have also shown that, during high-inflation periods, the "Patinkin effect (the inverse of the Olivera-Tanzi effect)," which indicates an in-verse correlation between the inflation rate and real public expenditures, has thereby caused a monetary loss that impacts the state's capacity to meet its fiscal obligations (Cardoso, 1998; Patinkin, 1993). Studies provide evidence that virtual deficits, which are calculated without considering inflation, expose the public finances to the impact of inflation on tax collection (Cardoso, 1998). This structural difficulty is further translated into a decrease in real public expenditures, which, in turn, intensifies the social and economic pressures. Moreover, the Patinkin effect is a strange stabilizing force in that it curbs hyperinflationary explosions but at the price of collective welfare decline.

Finally, the analysis of the two effects of inflation and chronic deficits in developing economies reveals a major cons-traint on the capacity to mobilize tax revenues. Current research indicates the necessity of enhancing the fiscal and institutional capacity to curb the impact of inflation on tax collection (Benitez et al., 2023; Peprah et al., 2025; Ha, Kose, and Ohnsorge, 2019). Burundi, with its institutions weak and monetary financing dependent, which in turn frac-tures the economy to face more inflationary shocks, is an excellent example (Kimolo et al., 2024). The situation calls for fiscal and monetary policy to work in tandem to build back confidence and stabilize the economy. These effects, which are perceived from the empirical studies, add to the existing knowledge of how inflation transmits in weaker economies like Burundi's.

1.2 Research Questions and Hypotheses

Burundi's economy presents some fundamental features, such as persistent inflation and structural budget defi-cits. This article's framework allows us to link to the Olivera–Tanzi effect, defined as the decrease in real tax reve-nues due to inflationary bracket creep and collection lags (Tanzi et al., 1987; Anušić and Švaljek, 1996), and to the Patinkin effect, stated as the erosion of real public spending due to rising prices (Cardoso, 1998; Patinkin, 1993). Therefore, our main issue is to determine whether these two complementary effects are really important or not in the Burundian context. In other words, questions about the government's ability to maintain budgetary discipline and the real value of resources and expenditures arise. Briefly, we aim to provide empirical evidence about the impor-tance of these two effects in this fragile economy with limited financial possibilities. So our research questions and hypotheses are presented below:

Question 1: Does chronic inflation cause real tax revenues to fall significantly in Burundi's economy, and therefore is the Olivera-Tanzi effect valid?

Question 2: Does chronic inflation reduce the real value of public spending, and therefore is the Patinkin effect valid in Burundi's economy?

Hypothesis 1: Chronic inflation causes real tax revenues to fall significantly in Burundi's economy, and therefore, the Olivera-Tanzi effect is valid.

Hypothesis 2: Chronic inflation reduces the real value of public spending; therefore, the Patinkin effect is valid in Burundi's economy.

1.3 Originality of the Study

The uniqueness of this study is found in the empirical application of the Olivera-Tanzi and Patinkin effects to the Burundi economy, a subject that is seldom examined in global literature. Most research on these phenomena has fo-cused on Latin American, European, and Asian countries; however, few published articles have evaluated them in African countries. Consequently, this study plays a vital role in bridging a scientific gap by presenting a novel pers-pective on the transmission of inflation effects within a fragile and structurally constrained African economy, the Burundi economy.

To address this work, the remainder of this paper is organized as follows: Section Two deals with the literature review, Section Three discusses the methodology, model specification, and results, and Section Four presents the research findings. Finally, the last section concludes this work.

2. Literature Review

2.1 Olivera-Tanzi Effect

The Olivera-Tanzi effect is based on the fact that during high-inflation periods, the real value of tax revenues col-lected decreases. It happens because there is a time lag between tax recognition and determination on one side and a tax collection that takes place later, with the interest accrued growing along due to an administrative separate delay in assessment, evaluation, and eventual collection of taxes (Tanzi et al., 1987; Anušić and Švaljek, 1996). Economists call this the Olivera-Tanzi effect, named after the work of Olivera (1967) and Tanzi (1977).

Researchers in Latin America–especially in Argentina and Brazil–have shown that chronic inflation slashes tax reve-nues by more than 20% in real terms (Tanzi, 1977; Blejer and Cheasty, 1991; Anušić and Švaljek, 1996). These days, economists still talk about this link between fiscal deficits and inflation, known as the Olivera–Tanzi effect, and con-tinue testing it. For instance, Serin and Demir (2025) used the Fourier approach to analyze the relationship between total tax revenues and inflation in Turkey over nearly a century, from 1924 to 2023. Their findings show that inflation drags down tax revenues, and they conclude that the Olivera-Tanzi effect is valid. Additionally, Akgül (2022), through an analysis of asymmetric and symmetric causality between inflation and tax revenues in Turkey, utilizing monthly time series data from January 2010 to June 2021, confirmed the validity of the Tanzi effect in the Turkish context. Furthermore, in nations where tax collection is sluggish and institutions are deficient in technological resources, this effect is intensified. Inflation does not just shrink what families can afford—it also throws a wrench into how go-vernments collect taxes (Dornbusch and Fischer, 1993; Sargent, 1982). However, Erdoğan and Erdoğan (2018) analy-zed Turkey's economy from 2006 to 2017. They ran VAR models and Granger causality tests, and what they found contradicted the previous Tanzi effect: inflation actually pushes both government spending and tax revenue up.

In Burundi, there is no empirical study that analyzes the Olivera-Tanzi effect. Only some studies have analyzed the relationship between inflation and public expenditures. Thus, the dynamic between inflation and public expenditures tends to promote an underestimation of budget deficits and raises concerns regarding the sustainability of public fi-nances (Nzirorera, 2016; Mvuyekure, Nimubona et al., 2025; Peprah et al., 2025). The empirical investigation of this mechanism in Burundi thus represents a novel contribution to the existing literature.

2.2 Patinkin Effect

The Patinkin effect postulates that when inflation goes up, the real value of public spending drops, which reduces budget deficits (Cardoso, 1998; Patinkin, 1993). Basically, as prices keep rising, the money the government spends does not go as far, so deficits end up smaller (Pekşen et al., 2023). The Patinkin effect is seen as the flip side of the Olivera-Tanzi effect (Cardoso, 1998; Patinkin, 1993).

Empirically, in 2020, Yalçın dug into the link between inflation and budget deficits, testing both the Olivera-Tanzi and Patinkin effects using SVAR modeling with quarterly data from 2006 to 2019. Their results clearly showed that the Patinkin effect appears in Turkey, but the Olivera–Tanzi effect does not really hold up there. Abdioğlu and Terzi (2009) took a close look at this using the Pesaran, Shin, and Smith (2001) bounds test and Turkish data from 1975 to 2005. They wanted to see if inflation and budget deficits are linked in the long-run. They also checked for both the Olivera–Tanzi and Patinkin effects. They found that, over time, inflation and budget deficits tend to move together, but there is a negative correlation between them. So, as one goes up, the other goes down. And in Turkey, the Patinkin effect clearly outweighs the Tanzi effect. But not everyone obtained the same results. Some studies could not back up the Patinkin effect at all. For instance, Biçen et al. (2015) used quarterly data from 1999 to 2014 and tested the Olive-ra–Tanzi and Patinkin effects with ARDL modeling and the Hsiao causality test. They did not find evidence for either effect in Turkey. Altunöz (2022) set out to see if the Olivera–Tanzi and Patinkin effects showed up in Turkey between 2005 and 2021. He used two different ARDL models with a bounds test and found that the Olivera–Tanzi effect was valid–basically, when inflation ran high, tax revenues dropped, and budget deficits increased. But he did not find any evidence for the Patinkin effect. Furthermore, Tülümce et al. (2021) utilized panel data from 16 Eurozone countries to investigate the validity of the Olivera–Tanzi and Patinkin effects. By applying the cointegration technique proposed by Pesaran (2006), they determined that only the Olivera–Tanzi effect was valid.

Research on Burundi often looks at how inflation links to economic growth, but hardly anyone digs into the Patinkin effect directly. Nzirorera (2012) dug into how inflation shapes public spending and growth, yet skips over the Patinkin effect's theoretical side. Mvuyekure et al. (2025) take a more dynamic approach, analyzing the interplay between public spending and inflation, which opens the door for deeper studies. Looking closely at the Patinkin effect in Bu-rundi sheds new light on budget constraints, especially given the country's fragile economy.

To the best of our knowledge, no one has run an empirical analysis in Africa–certainly not in Burundi–on the Olivera–Tanzi and Patinkin effects. This work breaks new ground. Our aim is clear: to test whether these effects actually play out in Burundi's economy.

3 Research Methodology

3.1 Variables and Data Collection

To evaluate the two hypotheses and address the research questions, this study used data on total real tax revenues (RTR) and data on total real public expenditures (RPE) for the endogenous variables, while it used data on the infla-tion rate (INF) as key independent variable; and data on the money supply (M3), data on real gross domestic product (RGDP) and data on the real interest rate (RIR) as control variables that can influence the direction of relationships between the endogenous and the exogenous variable.

The data used in this analysis are secondary ones, sourced from the Ministry of Finance, Budget and Digital Economy, and the Central Bank (BRB). This dataset is quarterly and spans from the first quarter of 2005 to the second quarter of 2024.

3.2 Models Specification, and Results

This study utilized the bounds test procedure of the Autoregressive Distributed Lag (ARDL) model to analyze the existence of potential long-run relationships (cointegration) between the endogenous and exogenous variables of each model, following the approach of Pesaran et al. (2001). The ARDL approach offers numerous advantages. Namely, it can be applied regardless of whether the variables are purely I(0) and/or I(1), provided that none is I(2)) (Kripfganz and Schneider, 2023). It also performs efficient results even with relatively small sample sizes, permits estimation of the cointegration relationship using OLS, and allows for the simultaneous estimation of short– and long–run elastici-ties, thereby enabling the validation or rejection of the underlying theoretical framework (Pesaran and Shin, 1995).

4 Research Findings

4.1 Graphical evolution of variables

Within the framework of this research, it is crucial to illustrate the evolution of significant variables to empirically contextualize the problems of persistent inflation and budget deficits in Burundi. The aforementionned also allows us to gain insights into their behavior over time.

Overall inflation trend

Graph 1 tracks Burundi's quarterly inflation and interest rates from early 2005 through the start of 2024. Some quar-ters—Q1 and Q3—stand out more clearly. Inflation swings up and down, sometimes dramatically as the interest rate does the opposite.

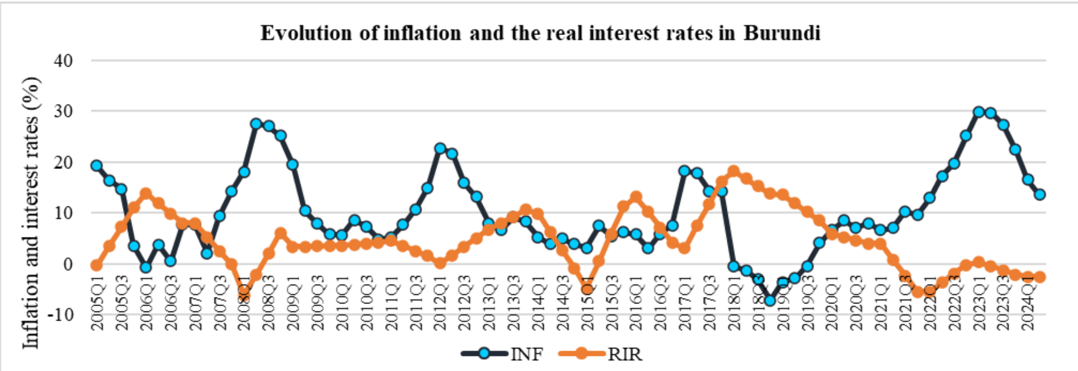

Figure 1: Evolution of quarterly inflation in Burundi (from 2005:Q1 to 2024:Q2)
Source: Plotted by the authors using data from the Ministry of Finance, Budget, and Digital Economy.

In general, Graph 1 above indicates that inflation and the real interest rates in Burundi are highly volatile. Notably, for inflation, peaks were recorded in 2004, the beginning of 2008, 2012, 2017, and 2023; during these periods, the infla-tion rate exceeded even 25%, indicating persistent inflationary pressures. After these inflation spikes, we see periods where it slows down—sometimes inflation drops low or even turns negative, like in 2018 and 2019. This back-and-forth shows just how hard it is to keep prices stable over time.

The real interest rate has moved in exactly the opposite direction to the inflation rate over the study period. The real interest rate peaked in 2006, at the end of 2008, 2013, 2016, and 2018. From the first quarter of 2009 to the first quar-ter of 2011, the real interest rate remained relatively stable. In contrast, the interest rate was very low with their mi-nimum points in 2008, 2015, and then again in 2022.

Trends of total real tax revenues and total real public expenditures in Burundi's economy

Graph 2 below shows the evolution of total real tax revenue (RTR) and total real public expenditures (RPE) over the period 2005Q1–2024Q2. These two variables form the core of the empirical analysis, as they allow for direct testing of the relevance of the Olivera-Tanzi and Patinkin effects in the Burundian context. Simultaneously observing reve-nue and expenditure in real terms provides an integrated view of fiscal dynamics and highlights the constraints faced by the state in an environment marked by chronic inflation.

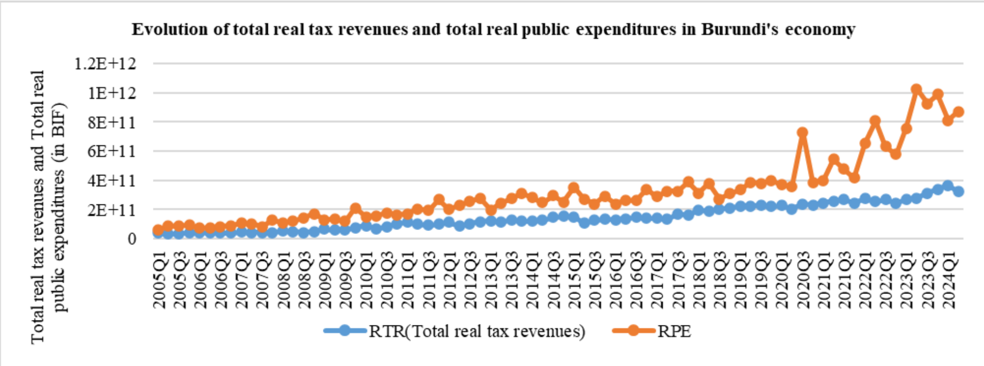

Figure 2: Evolution of total real tax revenues and total real public expenditures in Burundi's economy (from 2005:Q1 to 2024:Q2)
Source: Plotted by the authors using data from the Ministry of Finance, Budget, and Digital Economy.

From Graph 2 above, both series move upward, pointing to larger budget flows. But real public expenditure (RPE) fluctuates much more, especially from 2020. Government spending shot up, showing just how much strain public finances faced during that period. The gap between RTR and RPE keeps getting wider. That gap says a lot about Bu-rundi's fiscal vulnerability and backs up the main arguments of this research.

Trend in quantity of money in Burundi's economy

Graph 3 below shows the evolution of the quantity of money in the economy from the first quarter of 2005 until the second quarter of 2024.

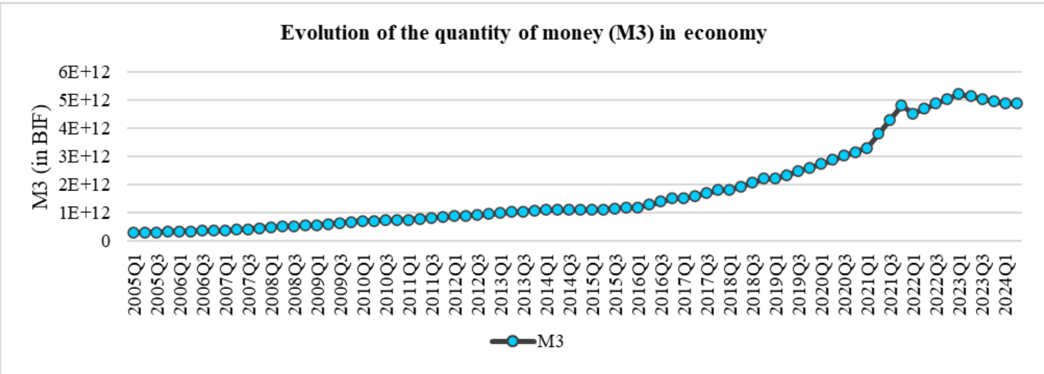

Figure 3: Evolution of the quantity of money in Burundi's economy (from 2005:Q1 to 2024:Q2)
Source: Plotted by the authors using data from the Ministry of Finance, Budget, and Digital Economy and, the Central Bank, BRB.

From Graph 3 above, it is obvious that for more than two decades, the quantity of money continued to grow. As far as the connection with the lessons from the figure 2 on the budgetary balance may be concerned, it is clear that the more the deficit increases, in tandem, the more the quantity of money continues to grow at a likely trend as fiscal imba-lances. Since 2005, the money supply has grown steadily. This suggests, other things remaining constant, the theoreti-cal relationship between the quantity of money and inflation appears to be evident in the Burundi economy.

Evolution of real GDP

Graph 4 below shows the evolution of real GDP from the first quarter of 2005 to the second quarter of 2024.

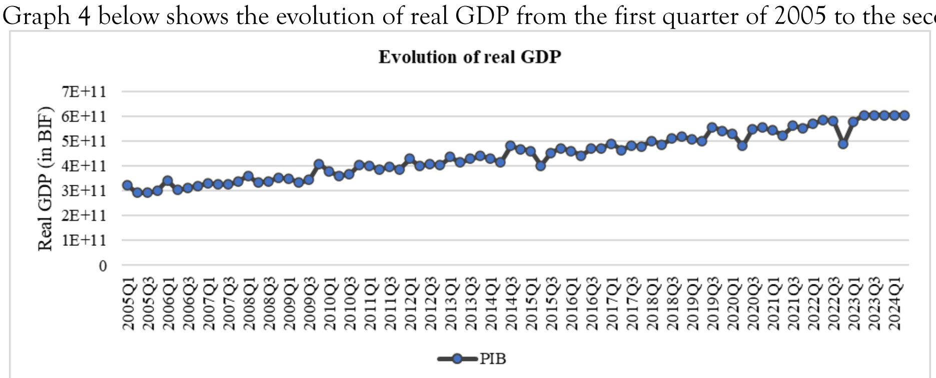

Figure 4: Evolution of real GDP (from 2005:Q1 to 2024:Q2)
Source: Plotted by the authors using data from the Ministry of Finance, Budget, and Digital Economy.

Graph 4 above shows an upward trend throughout the study period. However, there have been some significant de-clines characterized by valleys, particularly in the second quarter of 2015, 2020, and 2022. These shocks have di-srupted the growth of the Burundian economy. These shocks are probably explained by factors such as the political crisis that began at beginning of the second quarter of 2015 (Nkurunziza, 2018), the COVID-19 crisis, whose econo-mic effects began to be felt in the subsequent quarters, and finally, fuel shortages, very high inflation, and external imbalances since 2022.

4.2 Testing the Validity of the Olivera-Tanzi Effect

Based on the Olivera–Tanzi effect theory, this paper studied the equational relationship expressed as follows :

$$\mathrm{RTR} = \mathrm{f(INF, RGDP, M3, RIR)} \quad (1)$$

Where:

RFR= real fiscal revenues as the dependent variable of the model.
INF= inflation rate;
RGDP= real Gross Domestic Product;
M3= money supply; and
RIR= the real interest rate; are the independent variables.

From relation (1), and following Pesaran et al. (2001), cointegrationamong the variables in the Olivera–Tanzi effect model is analyzed using the ARDL specification expressed in its unrestricted error correction form. The unrestricted error correction model (ECM) form of the ARDL is given as follows:

$$\Delta \mathrm{Ln\ RFR_t} = \alpha_0 + \sum_{n-1}^{p} \emptyset_1\, \Delta \mathrm{LnRFR}_{t-i} + \sum_{n-0}^{q} \emptyset_2\, \Delta \mathrm{INF}_{t-i} + \sum_{n-0}^{q} \emptyset_3\, \Delta \mathrm{LnRGDP}_{t-i} + \sum_{n-0}^{q} \emptyset_4\, \Delta \mathrm{LnM3}_{t-i} + \sum_{n-1}^{p} \emptyset_5\, \Delta \mathrm{RIR}_{t-i} + \beta_1 \mathrm{LnRFR}_{t-1} + \beta_2\, \mathrm{INF}_{t-1} + \beta_3\, \mathrm{LnRGDP}_{t-1} + \beta_4 \mathrm{LnM3}_{\ t-1} + \beta_5 \mathrm{RIR}_{t-1} + \varepsilon_t \quad (2)$$

Cointegration is tested according to the following hypothèses :
The null hypothesis is: $\mathrm{H_0}: \emptyset_1 = \emptyset_2 = \emptyset_3 = \emptyset_4 = \emptyset_5 = 0$: There is no cointegration (no long-run relationship).
The alternative hypothesis is: $\mathrm{H_a}: \emptyset_1 \neq \emptyset_2 \neq \emptyset_3 \neq \emptyset_4 \neq \emptyset_5 \neq 0$: There is cointegration (a long-run relationship).

The decision criterion is based on comparing the empirical F–statistic with the lower-bound I (0) and upper-bound I (1) critical values at the 10%, 5%, 2.5% and 1 % significance levels. If the empirical F–statistic is lower than the I(0) bound values, the null hypothesis cannot be rejected, indicating the absence of cointegration (the long-run relationship) among the variables. If the empirical F–statistic lies within the range of the lower-bound I(0) and upper-bound I(10) critical values, the test is inconclusive. Conversely, if the empirical F–statistic is greater than the upper-bound I(1) critical values, the null hypothesis of no cointegration is rejected (Narayan, 2005). In this instance, there is cointegration (the long-run relationship) among the variables (Pesaran et al., 2001; Kripfganz and Schneider, 2023). When the cointegration is confirmed, the long-run relationship is then estimated based on the ARDL model, expressed as:

$$\mathrm{Ln\ RFR_t} = \alpha_0 + \beta_1 \mathrm{LnRFR}_{t-1} + \beta_2\, \mathrm{INF}_{t-1} + \beta_3\, \mathrm{LnRGDP}_{t-1} + \beta_4 \mathrm{LnM3}_{\ t-1} + \beta_5 \mathrm{RIR}_{t-1} + \varepsilon_t \quad (3)$$

The short-run Dynamic Relationship Is estimated based on the ARDL model expressed as follows:

$$\Delta \mathrm{Ln\ RFR_t} = \alpha_0 + \sum_{n-1}^{p} \emptyset_1\, \Delta \mathrm{LnRFR}_{t-i} + \sum_{n-0}^{q} \emptyset_2\, \Delta \mathrm{INF}_{t-i} + \sum_{n-0}^{q} \emptyset_3\, \Delta \mathrm{LnRGDP}_{t-i} + \sum_{n-0}^{q} \emptyset_4\, \Delta \mathrm{LnM3}_{t-i} + \sum_{n-1}^{p} \emptyset_5\, \Delta \mathrm{RIR}_{t-i} + \omega ECM_{t-1} + \varepsilon_t \quad (4)$$

α_0: is the constant term ; $\emptyset_i$: are the short-run elasticities and coefficients ; β_i: are the long-run elasticities and coefficients; ECM_{t-1}: is the error correction term lagged for one period; ω: is the speed of adjustment; Δ: is the first difference operator; Ln: is the natural logarithm; p: is the number of lags for the endogenous variable; q: is the number of lags for the exogenous variables, and ε_t: is the white noise.

Results of the Olivera–Tanzi Effect Test

This section presents the results of the various estimation techniques, following methodological steps mentionned Kripfganz and Schneider (2023) and Pesaran et al. (2001). The empirical analysis was carried out using Eviews 12 version. As first step, we analyzed the order of integration of the variables by studying their time series properties. Time series analysis typically begins with testing the stationarity of the variables, which is essential for identifying their integrating order. To this end, we employed ed two widely used unit root tests–the Augmented Dickey–Fuller (ADF) test and the Phillips–Perron (PP) test–both of which have non–stationarity as their null hypothesis. These tests were applied to each variable to ensure that none were integrated of order two I(2), as the presence of I(2) variables would invalidate the ARDL approach. Determining the integrating order is a crucial step for avoiding spurious regression in ARDL models, even though the bounds–testing does not, in principle, require pre-testing for unit roots (Kripfganz and Schneider, 2023). The results of the ADF and PP unit root tests are reported in Tables 1 and 2.

Table 1. Results of the ADF Unit Root Test

Model		ADF-Statistics				
				At level		
Variables		LnRFR	INF	LnRGDP	LnM3	RIR
With Constant	t-Statistic	-0.90	-1.68	0.73	0.94	-3.91***
	Prob.	0.99	0.44	0.99	0.99	0.00
		NO	NO	NO	NO	
With Constant & Trend	t-Statistic	-3.49*	-1.59	-8.06***	-1.27	-3.95**
	Prob.	0.05	0.78	0.00	0.88	0.01
			NO		NO	
Without Constant &Trend	t-Statistic	-2.98	-0.13	5.56	2.39	-1.74*
	Prob.	0.99	0.64	1.00	0.99	0.07
		NO	NO	NO	NO	
		At first difference				
With Constant	t-Statistic	-11.92***	-5.32***	-7.85***	-5.41***	-6.83***
	Prob.	0.00	0.00	0.00	0.00	0.00
With Constant & Trend	t-Statistic	-12.06***	-5.27***	-7.89***	-5.75***	-6.77***
	Prob.	0.00	0.00	0.00	0.00	0.00
Without Constant &Trend	t-Statistic	-11.02***	-5.31***	-10.74***	-4.65*	-6.85***
	Prob.	0.00	0.00	0.00	0.09	0.00

Note (*): Significant at the 10%; (**): Significant at the 5%; (***): Significant at the 1%; NO: Not significant.
*Mackinon (1996) one sided P-Value.
Source: Authors' computation using Eviews 12.

Table 2. Results of the Phillips-Perron Unit Root Test

Model		Phillips-Perron Statistics				
				At level		
Variables		LnRFR	INF	LnRGDP	LnM3	RIR
With Constant	t-Statistic	1.52	-3.00**	-0.49	1.45	-2.42***
	Prob.	0.99	0.04	0.88	0.99	0.14
		NO		NO	NO	
With Constant & Trend	t-Statistic	-3.41*	-3.01	-12.87***	-1.08	-2.53
	Prob.	0.06	0.13	0.00	0.92	0.31
			NO		NO	NO
Without Constant &Trend	t-Statistic	4.49	-1.97**	3.85	3.54	-1.50
	Prob.	1.00	0.04	1.00	0.99	0.12
		NO		NO	NO	NO
		At first difference				
With Constant	t-Statistic	-12.88***	-6.46***	-34.83***	-5.38***	-4.14***
	Prob.	0.00	0.00	0.00	0.00	0.00
With Constant & Trend	t-Statistic	-15.39***	-6.42***	-35.14***	-5.73***	-4.05**
	Prob.	0.00	0.00	0.00	0.00	0.01
Without Constant &Trend	t-Statistic	-10.98***	-6.50***	-12.88***	-4.59***	-4.19***
	Prob.	0.00	0.00	0.00	0.00	0.00

Note (*): Significant at the 10%; (**): Significant at the 5%; (***): Significant at the 1%; NO: Not significant.
*Mackinon (1996) one-sided P-Value.
Source: Authors' computation using Eviews 12.

The results of the ADF unit root test presented in Table 1 indicate that all variables, with exception of the interest rate–which is stationary at level and therefore classified as I(0)-become stationary after first differencing. Specifically, total real tax revenues revenues, the inflation rate, real GDP, and the money supply are stationary at first difference, implying that they are I(1). Phillips–Perron (PP) unit root results in Table 2 confirm this general pattern, with the exception of the inflation rate, which is found to be stationary in level, i.e., I(0), while the remaining variables become stationary after their first difference and are therefore I(1).
Based on combined findings from both tests, we conclude that the dependent variable is integrated of order one, I (1), while the independent variables are mixed integrating of I (0) and of order I (1). Since none of the variable is integrated of order two, the use of the ARDL model is relevant. Consequently, both the ARDL modelling approach and the OLS

estimation method be used to study the short-run and long-run relationships among the variables (Pesaran and Shin, 1995; Kripfganz and Schneider, 2023).
After establishing the order of integration, the optimal ARDL model specification was determined reffering to the Akaike Information Criterion (AIC). Based on this criterion, the ARDL (4, 2, 0, 0, 1) model was selected as the most suitable model, and the corresponding estimation results are reported in Table 3.

Table 3. Estimates of the optimal ARDL (4, 2, 0, 0, 1) model Selected

Variable	Coefficient	Std. Error	t-Statistic	Prob.*
LnRFR(-1)	0.262559	0.096559	2.719159	0.0085
LnRFR(-2)	0.074193	0.100120	0.741040	0.4615
LnRFR(-3)	0.135687	0.095903	1.414843	0.1621
LnRFR(-4)	0.219484	0.096025	2.285713	0.0257
INF	-0.007332	0.002956	-2.480188	0.0159
INF(-1)	0.001750	0.003905	0.448125	0.6556
INF(-2)	0.006389	0.002658	2.403441	0.0192
LnRGDP	0.973863	0.229743	4.238932	0.0001
LnM3	0.002552	0.062451	0.040861	0.9675
RIR	-0.010378	0.004348	-2.386966	0.0201
RIR(-1)	0.010935	0.004841	2.258638	0.0274
c	-18.27300	4.540566	-4.024388	0.0002
R-squared	0.986274	Mean dependent var.		25.56106
Adjusted R-squared	0.983839	S.D. dependent var.		0.638760
S.E. of regression	0.081203	Akaike info criterion		-2.036343
Sum squared resid	0.408821	Schwarz criterion		-1.662710
Log likehood	87.34468	Hannan-Quinn criterion		-1.887296
F-statistic	405.0065	Durbin-Watson stat.		1.840379
Prob(F-statistic)	0.000000			

Note (*): p-values and any subsequent tests do not account for the model.
Source: Authors' computation using Eviews 12.

After choosing the optimal ARDL (4, 2, 0, 0, 1) model and before submitting it to the cointegration test, we subjected it to the various validation tests in order to check whether there was one or more specification problems, namely: non-normality, autocorrelation of errors, heteroscedasticity, and instability. The results of these robustness tests are given in Table 4 and the two CUSUM and CUSUM-SQ graphs (Figure 5) displayed below:

Table 4. Diagnostic tests on the selected ARDL (4, 2, 0, 0, 1) model.

Test hypothesis	Test	F-statistic	Prob.
Autocorrelation	Breusch-Godfrey Serial Correlation LM test	1.22932	0.3080
Heteroscedasticity	ARCH test	2.010984	0.1038
Normality	Jarque-Berra test	1.213880	0.545016

Source: Authors' computation using Eviews 12.

For all robustness tests, the null hypothesis is accepted; and therefore, the model meets the validity conditions, namely, the absence of autocorrelation, the existence of normality, and homoscedasticity (Table 4). However, the stability tests of the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of squares of recursive residuals (CUSUMSQ) reveal a stability problem: the plots fall outside the 95 % significance level, particularly from the second quarter of 2015 and the fourth quarter of 2016 (Figure 5), showing potential structural breaks in the model of Burundi's real tax revenues.

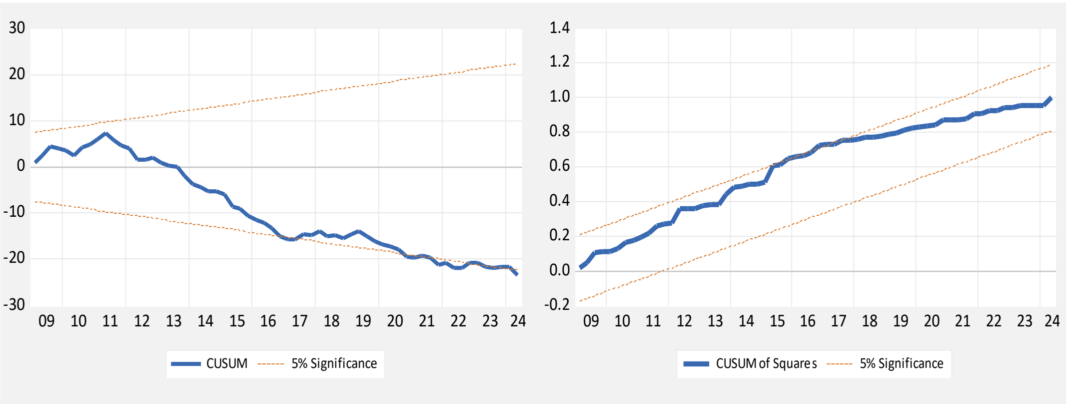

Figure 5. Stability Plots CUSUM stability test on ARDL (4, 2, 0, 0, 1)

Source: Authors' computation using Eviews 12.

To address the issue of model instability, this paper takes these structural breaks into account because they have particular significance for the country. In 2015, especially from its second quarter, Burundi was marked by a profound political crisis, fueled by protests and a failed coup, and the political and economic consequences of this violence were remarkable (Nkurunziza, 2018). This major shakeup in the economy gave us reason to update the model. We added a dummy variable, PV (Political Violence), to capture what happened in Burundi. Hence, here is how we set it up: from the first quarter of 2005 through the first quarter of 2015, PV's value was 0. Starting in the second quarter of 2015 up to the second quarter of 2024, PV's value switched to 1. The unrestricted error–correction model (ECM) from the ARDL model specification of Equation (2) turns into:

$$\Delta \text{Ln RFR}_t = \alpha_0 + \sum_{n-1}^{p} \emptyset_1 \, \Delta \text{LnRFR}_{t-i} + \sum_{n-0}^{q} \emptyset_2 \, \Delta \text{INF}_{t-i} + \sum_{n-0}^{q} \emptyset_3 \, \Delta \text{LnRGDP}_{t-i} + \sum_{n-0}^{q} \emptyset_4 \, \Delta \text{LnM3}_{t-i} + \sum_{n-1}^{p} \emptyset_5 \, \Delta \text{RIR}_{t-i} + \beta_1 \text{LnRFR}_{t-1} + \beta_2 \, \text{INF}_{t-1} + \beta_3 \, \text{LnRGDP}_{t-1} + \beta_4 \text{LnM3}_{\,t-1} + \beta_5 \text{PV}_{t-1} + \beta_6 \text{RIR}_{t-1} + \varepsilon_t \quad (5)$$

All test hypotheses and steps remain the same as in section 7.1. If cointegration is established, the long-run relationship from Equation (5), which is estimated based on the ARDL model, is now expressed as follows:

$$\text{Ln RFR}_t = \alpha_0 + \beta_1 \text{LnRFR}_{t-1} + \beta_2 \, \text{INF}_{t-1} + \beta_3 \, \text{LnRGDP}_{t-1} + \beta_4 \text{LnM3}_{\,t-1} + \beta_5 \text{PV}_{t-1} + \beta_6 \text{RIR}_{t-1} + \varepsilon_t \quad (6)$$

The short-run dynamic relationship in Equation (5), which is estimated based on the ARDL model, is expressed as follows:

$$\Delta \text{Ln RFR}_t = \alpha_0 + \sum_{n-1}^{p} \emptyset_1 \, \Delta \text{LnRFR}_{t-i} + \sum_{n-0}^{q} \emptyset_2 \, \Delta \text{INF}_{t-i} + \sum_{n-0}^{q} \emptyset_3 \, \Delta \text{LnRGDP}_{t-i} + \sum_{n-0}^{q} \emptyset_4 \, \Delta \text{LnM3}_{t-i} + \sum_{n-1}^{p} \emptyset_5 \, \Delta \text{RIR}_{t-i} + \emptyset_6 \text{PV}_{t-1} + \omega ECM_{t-1} + \varepsilon_t \quad (7)$$

Using the Akaike Information Criterion (AIC) as previously, the appropriate ARDL model selected among others is the ARDL (4, 2, 0, 0, 0, 1) model, as shown in Table 5 below:

Table 5. Estimates of the optimal ARDL (4, 2, 0, 0, 0, 1) selected

Variable	Coefficient	Std. Error	t-Statistic	Prob.*
LnRFR(-1)	0.224155	0.094894	2.362173	0.0214
LnRFR(-2)	0.058177	0.097101	0.599143	0.5513
LnRFR(-3)	0.139553	0.092808	1.395921	0.1678
LnRFR(-4)	0.224434	0.092912	2.415542	0.0187
INF	-0.006843	0.002868	-2.386239	0.0201
INF(-1)	0.000930	0.003794	0.245204	0.8071
INF(-2)	0.005395	0.002608	2.068951	0.0428
LnRGDP	1.019064	0.223109	4.567554	0.0000
LnM3	0.088856	0.071175	1.248421	0.2166
PV	-0.097063	0.042327	-2.293171	0.0253
RIR	-0.005464	0.004720	-1.157708	0.2515
RIR(-1)	0.008117	0.004842	1.676467	0.0988
c	-20.42316	4.491191	-4.547380	0.0000
R-squared	0.987364	Mean dependent var.		25.56106
Adjusted R-squared	0.984878	S.D. dependent var.		0.638760
S.E. of regression	0.078550	Akaike info criterion		-2.092008
Sum squared resid	0.376374	Schwarz criterion		-1.687239
Log likehood	90.40428	Hannan-Quinn criterion		-1.930541
F-statistic	397.1949	Durbin-Watson stat.		1.913767
Prob(F-statistic)	0.000000			

Source: Authors' computation using Eviews 12.

After selecting the optimal ARDL (4, 2, 0, 0, 0, 1) model and before submitting it to the cointegration test, we reapplied the robustness tests to check for specification problems and to determine whether the non-stability problem had been corrected. The results of these robustness tests are repported in Table 6, and the results of the stability tests are shown with CUSUM and CUSUMSQ graphs (Figure 7) presented below:

Table 6. Diagnostic tests on the selected ARDL (4, 2, 0, 0, 0, 1) model.

Test hypothesis	Test	F-statistic	Prob.
Autocorrelation	Breusch-Godfrey Serial Correlation LM test	1.06698	0.3813
Heteroscedasticity	ARCH test	2.095969	0.0914
Normality	Jarque-Berra test	0.569474	0.752212

Source: Authors' computation using Eviews 12.

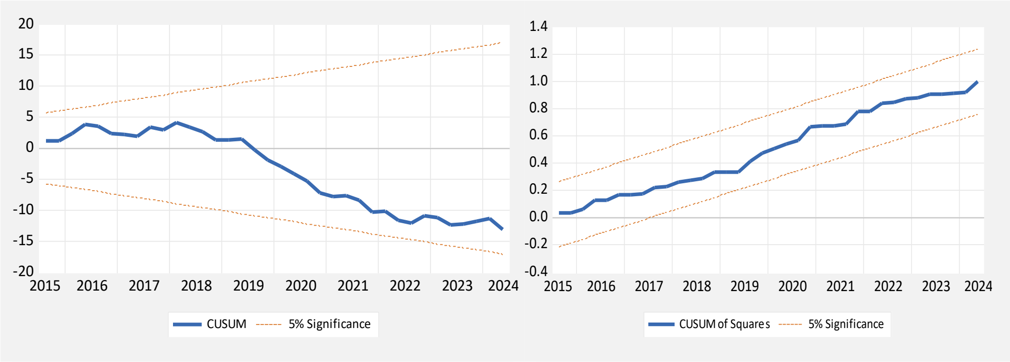

Figure 7. CUSUM stability test on the selected ARDL (4, 2, 0, 0, 0, 1) model.
Source: Authors' computation using Eviews 12.

For all robustness tests, the null hypothesis is accepted; and therefore, the model meets the validity conditions, namely the absence of autocorrelation, the existence of normality, homoscedasticity, and stability. The plots fall within the bounds at 95 % significance level, showing that the model is stable. After the appropriate ARDL (4, 2, 0, 0, 0, 1) model was selected and tested, we applied the Bounds test to check the cointegration among variables. The results are shown in Table 7 below.

Table 7. ARDL (4, 2, 0, 0, 0, 1) Bounds Test for Cointegration

Test Statistic	Value	Signif.	Lower Bounds: I(0)	Upper Bounds: I(1)
F-Statistic	12.08008***	10%	2.08	3
k = 5		5%	2.39	3.38
		2.5%	2.7	3.73
		1%	3.06	4.15

Note (***): Significant at the 1%; k: stands for the number of regressors.
Source: Authors' computation using Eviews 12.

The value of the empirical F-statistic is 12.08008. Since this F-statistic is greater than the values of both the lower and upper bounds at the 1 %, 2.5 %, 5 %, and 10 % thresholds, this work rejected the null hypothesis of no cointegration and accepted the alternative hypothesis of the existence of a long-run relationship. It confirms a cointegrating relationship between the endogenous variable and the exogenous variables in the model. This evidence of the existence of cointegration relationships therefore allows us to proceed with the estimation of the long-run and short-run relationships of our ARDL cointegration model.

The long-run equilibrium relationship in the Olivera-Tanzi Effect model

The results of the long-run estimated coefficients of the selected ARDL (4, 2, 0, 0, 0, 1) model are presented in Table 8 below.

Table 8. Econometric results for the long-run ARDL (4, 2, 0, 0, 0, 1) selected model (real tax revenues is endogenous variable), 2005:Q1-2024:Q2.

Variable	Coefficient	Std. Error	t-Statistic	Prob.
INF	-0.001423	0.005316	-0.267735	0.7898
LnRGDP	2.801083***	0.643293	4.355839	0.0001
LnM3	0.244324	0.163659	1.492886	0.1406
PV	-0.266891**	0.114997	-2.320850	0.0237
RIR	0.007294	0.007057	1.033480	0.3055
c	-56.15681***	13.25191	-4.237640	0.0001
EC = LnRFR - (-0.0014.INF + 2.8021.LnPIB + 0.2443.LnM3 - 0.2669.PV + 0.0073.RIR-56.1568)				

Note (**): Significant at the 5%; (***): Significant at the 1%.
Note (***): Significant at the 1%.
Source: Authors' computation using Eviews 12.

The estimates from Table 8 show a positive long-run equilibrium relationship between real GDP (LnRGDP) and total real tax revenues (LnRFR). Therefore, a 1 % change in real GDP leads to a 2.80 % change in total real tax revenues. As far as the impact of structural breaks on real tax revenues is concerned, since the coefficient of the dummy variable (PV) is negative and statistically significant, this shows that political violence had a negative impact on real tax revenues. Therefore, in the long-run, the presence of political violence resulted in a 23.4 % [$e^{-2.266891}$= 0.766 ⇒ {(0.7657-1)*100%} = –23.43%] decrease in total real tax revenues compared with the period of absence of political violence. Besides this, it should be noted that although the long-run effect of inflation (INF) is negative as expected, it is not statistically significant. The money supply (M3) and the real interest rate (RIR) are also not statistically significant.

The dynamics of short-run causality in the Olivera-Tanzi Effect model

Following the results of the Bounds test, we estimated the short-run dynamic relationship of the selected ARDL (4, 2, 0, 0, 0, 1) model of equation (7). The estimated ECM results of are given in Table 9 below.

Table 9. The error correction model (ECM) of the Olivera-Tanzi effect

Variable	Coefficient	Std. Error	t-Statistic	Prob.
ΔLnRFR(-1)	-0.412164***	0.076614	-5.379724	0.0000
ΔLnRFR(-2)	-0.353987***	0.073473	-4.817895	0.0000
ΔLnRFR(-3)	-0.224434***	0.072146	-3.110812	0.0028
ΔINF	-0.006843***	0.002241	-3.053829	0.0033
ΔINF(-1)	-0.005395**	0.002242	-2.406092	0.0192
ΔRIR	-0.005464	0.003690	-1.481037	0.1437
CointEq(-1)	-0.363681***	0.037737	-9.637322	0.0000

Note (**): Significant at the 5%; (***): Significant at the 1%.
Note: Adjusted R-squared= 65%. DW=1.91. *p-value incompatible with t-Bounds distribution.
Source: Authors' computation using Eviews 12.

The estimates obtained in Table 9 above show that the adjustment coefficient towards the long-run equilibrium CointEq(-1) is negative and statistically significant at the 1% level. Hence, there is a causal relationship among variables (Narayan, 2005). This reinforces the results found with the Bounds test, i.e., there is a error correction mechanism and, therefore, a long–run relationship among the variables. Thus, this negative value of the error correction term ECT (-0.36) for the real total tax revenue equation indicates that 36 % of the variations between the short-run and long-run equilibrium values of real total tax revenue in Burundi are corrected each quarter. Following Ayoub (2020), we calculated the speed at which real total tax revenues return to equilibrium after any eventual exogenous shock using the following formula:

$$(1-\alpha) = (1-|\widehat{\omega}|)^t \quad (8)$$

Thus, by applying the logarithm, we got:

$\Leftrightarrow Log(1-\alpha) = t * \log(1-|\widehat{\omega}|)$

$$\Leftrightarrow t = \frac{Log(1-\alpha)}{\log(1-|\widehat{\omega}|)} \quad (9)$$

Where,

t = number of quarters;

$\widehat{\omega}$ = the estimated value of the error correction coefficient; and

α = the percentage of shock to be removed (i.e., 95%).

By substituting the values in formula (9), we found that eliminating 95% of a potential exogenous shock in real total tax revenues requires 7 quarters before their return to equilibrium.

$$t = \frac{Log(1-0.95)}{\log(1-|-0.36|)} = \frac{Log(0.05)}{log\ (0.64)} = \frac{-1.301}{-0.194} = 6.7$$

Additionally, the results in Table 9 above also show that past values of the dependent variable have a reversionary trend effect for the next period. As the coefficients of the actual and past values of inflation are negative and statistically significant, inflation has a negative effect on real total tax revenues and does not appear to change its effect over time. This confirms our first research question and hypothesis. In the short-run, a unit increase in the actual inflation rate leads to a 0.68% decrease in real total tax revenue, and a unit increase in the past inflation rate leads to a 0.54% decrease in real tax revenues. Thus, in Burundi, in the short term, the Olivera-Tanzi effect is valid.

4.3 Testing the Validity of the Patinkin Effect

Based on the Patinkin effect theory, this paper also studies the equational relationship expressed as follows :

$$\mathrm{RPE} = \mathrm{f(INF, RGDP, M3,\ RIR)} \quad (10)$$

Where,

RPE= real public expenditures as the dependent variable of the model.

INF= inflation rate;

RGDP= real Gross Domestic Product;

M3= money supply; and

RIR= the real interest rate are the independent variables.

From relation (10), to test the cointegration among the variables in the Patinkin effect model and, following Pesaran, Shin, and Smith (2001), the form of the ARDL model expressed as an unrestricted error–correction model (ECM) is as follows:

$$\Delta \text{Ln RPE}_t = \alpha_0 + \sum_{n-1}^{p} \emptyset_1 \Delta \text{LnRPE}_{t-i} + \sum_{n-0}^{q} \emptyset_2 \Delta \text{INF}_{t-i} + \sum_{n-0}^{q} \emptyset_3 \Delta \text{LnRGDP}_{t-i} + \sum_{n-0}^{q} \emptyset_4 \Delta \text{LnM3}_{t-i} + \sum_{n-1}^{p} \emptyset_5 \Delta \text{RIR}_{t-i} + \beta_1 \text{LnRFR}_{t-1} + \beta_2 \text{INF}_{t-1} + \beta_3 \text{LnRGDP}_{t-1} + \beta_4 \text{LnM3}_{t-1} + \beta_5 \text{RIR}_{t-1} + \varepsilon_t \quad (11)$$

Cointegration is tested using to the following hypotheses:

The null hypothesis is: $H_0: \emptyset_1 = \emptyset_2 = \emptyset_3 = \emptyset_4 = \emptyset_5 = 0$: There is no cointegration (there is no long-run relationship).
The alternative hypothesis is: $H_a: \emptyset_1 \neq \emptyset_2 \neq \emptyset_3 \neq \emptyset_4 \neq \emptyset_5 \neq 0$: There is cointegration (there is a long-run relationship).
The decision criterion is based on the comparison between the empirical F–statistic and the lower–bound I (0) and upper - bound I (1) at the 10%, 5%, 2.5%, and 1 % significance levels. If the empirical F–statistic is lower than the lower-bound values, we fail to reject the null hypothesis; therefore, there is no cointegration (the long-run relationship) between the variables. If the empirical F–statistic lies within the range of the lower–and upper–bound values, the test is inconclusive. If the empirical F–statistic is greater than the upper–bound values, the null hypothesis of no cointegration is rejected (Narayan, 2005); in this instance, there is cointegration (the long-run relationship) between the variables (Pesaran et al., 2001; Kripfganz and Schneider, 2023). If cointegration is established, the long-run relationship is estimated based on the ARDL model expressed as follows:

$$\text{Ln RPE}_t = \alpha_0 + \beta_1 \text{LnRPE}_{t-1} + \beta_2 \text{INF}_{t-1} + \beta_3 \text{LnRGDP}_{t-1} + \beta_4 \text{LnM3}_{t-1} + \beta_5 \text{RIR}_{t-1} + \varepsilon_t \quad (12)$$

The short-run dynamic relationship is estimated based on the ARDL model expressed as follows:

$$\Delta \text{Ln RPE}_t = \alpha_0 + \sum_{n-1}^{p} \emptyset_1 \Delta \text{LnRPE}_{t-i} + \sum_{n-0}^{q} \emptyset_2 \Delta \text{INF}_{t-i} + \sum_{n-0}^{q} \emptyset_3 \Delta \text{LnRGDP}_{t-i} + \sum_{n-0}^{q} \emptyset_4 \Delta \text{LnM3}_{t-i} + \sum_{n-1}^{p} \emptyset_5 \Delta \text{RIR}_{t-i} + \omega ECM_{t-1} + \varepsilon_t \quad (13)$$

α_0: is the constant term; $\emptyset_i$: are the short-run elasticities and coefficients; β_i: are the long-run elasticities and coefficients; ECM_{t-1}: is the error correction term lagged for one period; ω: is the speed of adjustment; Δ: is the first difference operator; Ln: is the natural logarithm; p: is the number of lags for the endogenous variable; q: is the number of lags for the exogenous variables, and ε_t: is the white noise.

Presentation of Results of the Patinkin Effect Testing

The exogenous variables and their data remained the same as in the Olivera-Tanzi effect model. The results of the ADF and PP tests on these variables remain the same as those shown in Tables 1 and 2. Recall that these results showed that the exogenous variables were integrated of order I (0) and I (1). Therefore, here, in Table 10, we present simultaneously and exclusively the results of the ADF and PP unit root tests on the endogenous variable: real public expenditures (RPE).

Table 10. Results of the ADF Unit Root Test

Model		**ADF Statistics**	Phillips-Perron
		At level	
Variable		LnRPE	
With Constant	t-Statistic	0.031891	-1.324173
	Prob.	0.9580	0.6146
		NO	NO
With Constant & Trend	t-Statistic	-6.197599***	-6.476574***
	Prob.	0.0000	0.0000
Without Constant &Trend	t-Statistic	3.764320	4.396619
	Prob.	0.9999	1.0000
		NO	NO
		At first difference	
With Constant	t-Statistic	-9.481667***	-40.11843***
	Prob.	0.0000	0.0001
With Constant & Trend	t-Statistic	-9.435329***	-39.72926***
	Prob.	0.0000	0.0001
Without Constant &Trend	t-Statistic	-9.875030***	-14.89771***
	Prob.	0.0000	0.0000

Note (***): Significant at the 1%; NO: Not significant.
*Mackinon (1996) one sided P-Value.
Source: Authors' computation using Eviews 12.

The results of the two tests, presented in Table 10 above, show that the endogenous variable is integrated of order I(1). Overall, the results obtained reveal the possibility of a cointegrating relationship between the different variables of the Patinkin effect model. From the equation (11), we used Akaike information criterion (AIC) to select the optimal ARDL model–the one that yielded statistically significant results with the fewest parameters. The estimation results for the optimal ARDL model gave us an ARDL (1, 1, 3, 2, 0) model, as shown in Table 11 below:

Table 11. Estimates of the optimal ARDL (1, 1, 3, 2, 0) model Selected

Variable	Coefficient	Std. Error	t-Statistic	Prob.*
LnRPE(-1)	-0.207975	0.111712	-1.861716	0.0673
INF	0.002061	0.004454	0.462580	0.6453
INF(-1)	0.010026	0.004390	2.283615	0.0258
LnRGDP	1.331204	0.354951	3.750385	0.0004
LnRGDP(-1)	0.345378	0.356848	0.967859	0.3368
LnRGDP(-2)	0.165029	0.322274	0.512077	0.6104
LnRGDP(-3)	1.048007	0.320606	3.268826	0.0018
LnM3	-0.704424	0.558081	-1.262226	0.2115
LnM3(-1)	-1.005300	0.940664	-1.068713	0.2893
LnM3(-2)	-1.9141105	0.597482	3.203622	0.0021
RIR	-0.004054	0.003715	-1.091271	0.2793
c	-51.36443	12.44487	-4.127357	0.0001
R-squared	0.962237	Mean dependent var.		26.30798
Adjusted R-squared	0.955644	S.D. dependent var.		0.652980
S.E. of regression	0.137524	Akaike info criterion		-0.984393
Sum squared resid	1.191506	Schwarz criterion		-0.613595
Log likehood	48.91475	Hannan-Quinn criterion		-0.836338
F-statistic	145.9369	Durbin-Watson stat.		1.983193
Prob(F-statistic)	0.000000			

Note (*): p-values and any subsequent tests do not account for the model.
Source: Authors' computation using Eviews 12.

The specification obtained in the ARDL (1, 1, 3, 2, 0) model is generally satisfactory. The model explains almost 96.2% of the observed variability in real tax revenues. Regarding robustness tests, the null hypothesis is accepted by all tests; therefore, the residuals meet the model's validity conditions, namely the absence of autocorrelation and heteroscedasticity and the presence of normality (Table 12).

Table 12. Diagnostic tests on the selected ARDL (1, 1, 3, 2, 0) model.

Test hypothesis	Test	F-statistic	Prob.
Autocorrelation	Breusch-Godfrey Serial Correlation LM test	0.516818	0.6723
Heteroscedasticity	ARCH test	0.733188	0.5358
Normality	Jarque-Berra test	4.691932	0.095755

Source: Authors' computation using Eviews 12.

Furthermore, according to the CUSUM and CUSUMSQ graphs (Figure 8) tests, the bleu plots do not travel beyond the bounds of the 95 % significance level, showing that model is stable.

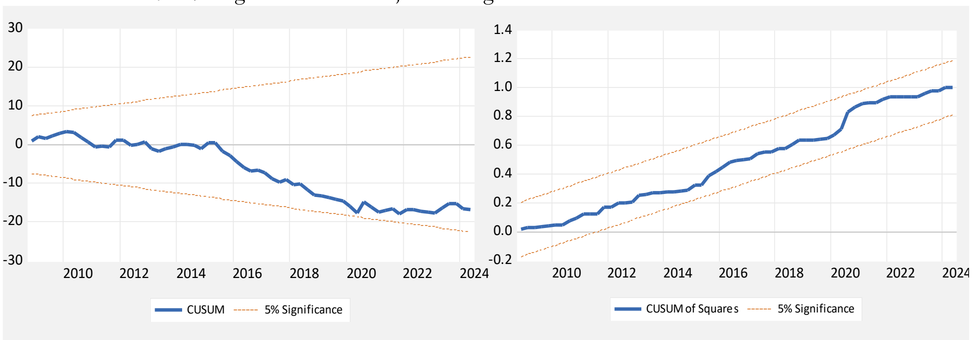

Figure 8. CUSUM stability test on the selected ARDL (1, 1, 3, 2, 0) model.
Source: Authors' computation using Eviews 12.

Hereinafter, Table 13 provides values for the Bounds test, which uses Fisher's exact test to verify the cointegration hypotheses. Thus, we test the null hypothesis of the absence of cointegration against the alternative hypothesis of the existence of a cointegrating relationship using the traditional approach of Pesaran et al.(2001) and the approach of Narayan (2005) for small samples. The bounds test table shows that the empirical Fisher statistic, which has the value 20.77875, is greater than all the lower and upper bounds values. This leads us to reject the null hypothesis of no cointegrating relationship and, consequently, to accept the alternative hypothesis of the existence of a cointegrating relationship between the variables. This evidence of cointegrating relationships, therefore, allowed us to proceed with the estimation of the long- and short-run relationships of our ARDL (1, 1, 3, 2, 0) cointegrating model.

Table 13. ARDL (1, 1, 3, 2, 0) Bounds Test for Cointegration

Test Statistic	Value	Signif.	Lower Bounds: I(0)	Upper Bounds: I(1)
F-Statistic	20.77875***	10%	2.2	3.09
k = 4		5%	2.56	3.49
		2.5%	2.88	3.87
		1%	3.29	4.37

Note (***): Significant at the 1%; k: stands for the number of regressors.
Source: Authors' computation using Eviews 12.

The bounds test Table 13 above shows that the empirical Fisher statistic, which has the value 20.77875, is greater than all the lower and upper bounds values. This leads us to reject the null hypothesis of no cointegrating relationship and, consequently, to accept the alternative hypothesis of the existence of a cointegrating relationship among the variables. This evidence of cointegrating relationships, therefore, allowed us to proceed with the estimation of the long- and short-run relationships of our ARDL (1, 1, 3, 2, 0) cointegrating model.

The long-run equilibrium relationship in the Patinkin Effect model

The results of the long-run estimate coefficients of the selected ARDL (1, 1, 3, 2, 0) model are presented in Table 14 below.

Table 14. Econometric results for the long-run ARDL (1, 1, 3, 2, 0) selected model (real public expenditures is endogenous variable), 2005: Q1-2024:Q2.

Variable	Coefficient	Std. Error	t-Statistic	Prob.
INF	0.010006***	0.002169	4.612252	0.0000
LnRGDP	2.392117***	0.452352	5.288179	0.0000
LnM3	0.169194	0.108805	1.555014	0.1250
RIR	-0.003356	0.003061	-1.096348	0.2771
c	-42.52110***	9.147622	-4.648322	0.0000
EC = LnRPE - (0.00100.INF + 2.3921.LnPIB + 0.1692.LnM3 - 0.0034.RIR-42.5211)				

Note (***): Significant at the 1%.
Source: Authors' computation using Eviews 12.

Table 14 provides the long-run coefficients. The results show that the long-run inflation rate is positive and significant. Therefore, a 1 % point increase in inflation, all other things being equal, leads to a 1 % increase in real public expenditures. This result contradicts theoretical predictions of the Patinkin effect. However, it aligns with Nzirorera (2016)'s findings, which indicate that public expenditures grows rapidly during periods of inflation. The aforementioned result, therefore, reinforces the Olivera-Tanzi effect. The other variable that positively and significantly impacts real public expenditures is real GDP. In the long run, a 1% increase in real GDP is associated with a 2.39% increase in real public expenditures.

The dynamics of short-run causality in the Patinkin Effect model

Following the results of the Bounds test, we estimated the short-run dynamic relationship of the selected ARDL (1, 1, 3, 2, 0) model in equation (13). The estimated ECM results are given in Table 15 below.

Table 15. The error correction model (ECM) for the Patinkin effect

Variable	Coefficient	Std. Error	t-Statistic	Prob.
ΔINF	0.002061	0.003630	0.567687	0.5723
ΔLnRGDP	1.331204***	0.293493	4.535726	0.0000
ΔLnRGDP(-1)	-1.213036***	0.293107	-4.138539	0.0001
ΔLnRGDP(-2)	-1.048007***	0.271271	-3.863317	0.0003
ΔLnM3	-0.704424	0.478142	-1.473251	0.1457
ΔLnM3(-1)	-1.914105***	0.517872	-3.696095	0.0005
CointEq(-1)	-1.207975***	0.104133	-11.60031	0.0000

Note (**): Significant at the 5%; (***): Significant at the 1%.

Note: Adjusted R-squared= 68%. DW=1.98. *p-value incompatible with t-Bounds distribution.
Source: Authors' computation using Eviews 12.

The estimates obtained in Table 15 above show that the adjustment coefficient CointEq(-1) is negative and statistically significant, indicating the existence of an error–correction mechanism and, consequently, a long–run relationship among the variables. However, surprisingly, the value of this coefficient is plausible. A value of -1.21 indicates a highly plausible speed of adjustment in real public expenditures. The model data are quarterly, and the model has undergone various diagnostic tests. It has been found to have no instability issues. Thus, for a well-specified and stable model, once the error–correction coefficient is negative and statistically significant, the range of coefficient values should be $-2 < \omega < 0$. Within this interval, if the estimated value is $-1 < \omega < 0$, then the model converges monotonically toward its equilibrium. If $-2 < \omega < -1$, then the model converges in an oscillatory (damped) manner to its equilibrium. In other words, in the latter case, the dependent variable will oscillate (increase, then decrease) around its equilibrium value before stabilizing. The other extreme cases concern a poorly specified or unstable (in this case, the coefficient would have a value of $\omega < -2$) and the case where there is no error–correction mechanism (in this case, the coefficient would have a value of $0 > \omega$) (Taha, 2025). In our case, the value of -1.21 shows that the real public expenditures model, although converging, is damped in its return to equilibrium. By substituting into formula (9), we found that eliminating 95% of a potential exogenous shock in real public expenditures requires two quarters before their return to equilibrium.

$$t = \frac{Log(0.05)}{\log(0.21)} = 1.92$$

Additionally, the results show that, in the short–run, the semi-elasticity of the inflation rate is positive. It is not statistically significant; thus, contemporaneous inflation does not have a significant impact on real public expenditures. However, real GDP has a positive and significant effect on real public expenditures. In the immediate term, a 1 % increase in real GDP leads to a 1.33 % increase in real public expenditures. However, this effect appears to change over subsequent quarters and become negative. The other significant variable is the money supply, whose immediate effect remains negative but is not statistically significant. However, in subsequent quarters, it negatively and significantly affects real public expenditures. Thus, a 1% increase in the money supply leads to a 1.91% decrease in real public expenditures. The results in Tables 14 and 15 refute the second hypothesis and answer the second research question of this study. In the long run, inflation, rather than decreasing real public expenditures, increases them. In the short run, its positive effect on real public expenditures is not significant. Thus, in Burundi, the Patinkin effect does not apply.

5 Conclusions and Recommendations

For nearly two decades, the Burundian economy has faced public finance management problems. On the one hand, public spending has increased dramatically, and revenues are no longer sufficient to cover it. As a result, the economy finds itself in a situation of recurring budget deficits. On the other hand, inflation continues its upward trend, becoming increasingly chronic and pervasive, steadily eroding the purchasing power of the average citizen (Kimolo et al., 2024; UNICEF, 2023). Although this situation is underappreciated, the coexistence of this twin problem—chronic deficits and chronic inflation (BAD, 2024)—raises fundamental questions about the impact of inflation on public finances.

Hitherto, studies show that inflation reduces the real value of fiscal revenues, a phenomenon known as the "Olivera–Tanzi effect". It occurs because there is a time lag between recognition and determination on one side and tax collection, which is realized later with interest accruing due to administrative delays in assessment, evaluation, and eventual collection of taxes (Tanzi et al., 1987; Anušić and Švaljek, 1996). This is an especially important dynamic in economies where tax systems are affected by bureaucratic inefficiencies and low institutional capacity. Thus, when inflation increases, the real value of tax revenues drops and, therefore, budget deficits worsen. Yet other studies postulate that when inflation increases, the real value of public expenditures drops, which therefore helps shrink budget deficits; a phenomenon known as the "Patinkin effect." As inflation increases, the real value of the money that the government spends does not increase, so eventually deficits widen. The Patinkin effect is seen as the flip side of the Olivera-Tanzi effect. This really matters in developing countries, where budgets are pretty rigid, and governments cannot adapt easily.

Therefore, this study aims to empirically analyze the Olivera-Tanzi and Patinkin effects in the Burundian economy. To achieve this objective, the study uses total real tax revenue and total real government expenditures as dependent variables. Inflation is the main independent variable, whereas real GDP, the money supply, and the real interest rate are control variables that may influence the relationship between the main independent variable and the dependent variables. The data are quarterly, ranging from 2004:Q1 to 2024:Q2.

The ADF and PP unit root tests showed that the models' series were integrated of order I(0) and I(1), respectively. This revealed that the ARDL approach and the bounds test for cointegration technique should be used. Since the diagnostic tests did not detect any anomalies in the real public expenditures (Patinkin effect) model, the CUSUM and CUSUMSQ stability tests showed that the real tax revenues (Olivera-Tanzi) model experienced potential breaks, particularly from the second quarter of 2015 onward. This break is particularly significant in Burundi, as it corresponds to the political violence that plagued the country from the beginning of that quarter. Accordingly, this paper introduced a dummy variable (PV = Political Violence) into the ARDL model of real tax revenues to capture these structural breaks. For the Olivera-Tanzi effect model, an optimal ARDL (4, 2, 0, 0, 0, 1) was chosen, while for the Patinkin effect model, an optimal ARDL (1, 1, 2, 3, 0) was chosen using the Akaike Information Criterion (AIC). Following the bounds test approach of Pesaran et al. (2001), the long-term cointegration relationship was tested for both ARDL models, and the results of this test revealed that there was a long-term relationship in these.

Concerning the results of the Olivera-Tanzi effect model, in the short run, this study shows that 36 % of the variation in real tax revenues is corrected each quarter. The model revealed that 95% of a potential exogenous shock to real tax revenues requires seven quarters before returning to equilibrium. Additionally, the results show that past values of real tax revenues have a reversionary trend effect for subsequent periods. In the short run, a one–unit increase in the contemporaneous inflation rate leads to a 0.68% decrease in real tax revenues, and a one–unit increase in the past inflation rate leads to a 0.54% decrease in real tax revenues. Thus, in Burundi, in the short run, the Olivera–Tanzi effect is valid. In the long run, a 1 % change in real GDP leads to a 2.80 % change in real tax revenues. Furthermore, in the long-run, the presence of political violence resulted in a 23.4 % decrease in real tax revenues compared with the period of absence of political violence.

As far as the results of the Patinkin effect model are concerned, in the short run, the model, although converging, is damped in its return to equilibrium. It revealed that 95% of a potential exogenous shock in real public expenditures requires two quarters before returning to equilibrium. Additionally, the contemporaneous inflation does not have a significant impact on real public expenditures. A 1% increase in the money supply leads to a 1.91% decrease in real public expenditures, but in the second quarter. Conversely, real GDP has a positive and significant effect on real public expenditures. In the immediate term, a 1 % increase in real GDP leads to a 1.33 % increase in real public expenditures. However, this effect appears to change over subsequent quarters and becomes negative.

In the long run, a 1% increase in real GDP is associated with a 2.39% increase in real public expenditures. Conversely, a 1%–point increase in inflation, leads to a 1 % increase in real public expenditures. This result contradicts theoretical predictions of the Patinkin effect. However, it aligns with Nzirorera (2016)'s findings, which indicate that public expenditures grows rapidly during periods of inflation. This result, therefore, reinforces the Olivera–Tanzi effect rather than the Patinkin effect. Thus, in Burundi, the Patinkin effect postulate does not hold.

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The article followed all ethical standards appropriate for this kind of research.

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