AN EMPIRICAL ASSESSMENT OF THE MACROECONOMIC EFFECT OF A SHIFT FROM INDIRECT TO DIRECT TAXES IN COTE D'IVOIRE

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ABSTRACT: *This paper examines the long-run effect of a shift from indirect to direct taxes for Cote d'Ivoire using data for the period 1960 to 2006. The residual-based test of Gregory and Hansen (1996) is employed to test for cointegration and the Engle-Yoo (1987) three-step procedure is used to estimate the long-run effect of tax variables on real output. The results reveal that tax burden and tax mix are negatively associated with output, with tax burden having a much greater adverse effect on GDP than tax mix. The effect of the tax mix on GDP is contingent on the level of the tax burden and diminishes as tax burden increases. Our estimates also suggest that up to a threshold level of tax burden of 17.57%, increased direct taxation relative to indirect taxation is associated with decreased output. But beyond this threshold a move from indirect to direct taxes is likely to lead to higher levels of output*.

KEYWORDS: Macroeconomic, Indirect, Direct, Taxes, Cote d'Ivoire

INTRODUCTION

The last two decades have witnessed an upsurge of empirical research aimed at unravelling the relationship between various measures of fiscal variables and economic growth. In this endeavour, cross-section, panel, and time series data have been used. Attempts to underpin the growth relationship are undermined by conceptual, statistical and estimation concerns. Not surprisingly, empirical findings have been diverse. For example, studies such as Barro (1990), Engen and Skinner (1992), Kormendi and Meguire (1995), Cashin (1995), Engen and Skinner (1996), Kneller *et al.* (1999), Fölster and Henrekson (2001), Bleaney *et al*. (2001), Blanchard and Perotti (2002) and Karras and Furceri (2009) present evidence favoring the view that that taxation is negatively associated with economic growth. While others such as Katz *et al.* (1983), Koester and Kormendi (1989), Slemrod (1995) and Mendoza *et al.* (1997) do not detect any significant relationship. A strand of the literature argues that what matters for growth is not only the level of taxes but also the way in which different tax instruments are designed and combined to generate revenues. Some taxes are negatively associated with economic growth than others. For example, consumption taxes have been found to be less distortionary than taxes on capital and income (see Skinner, 1987; Wang and Yip, 1992; Kim, 1998; Widmalm, 2001; Arnold, 2008). Higher direct taxes reduce personal disposal income, discourage private investment and consumption, thereby impeding economic growth. Furthermore, higher income taxes create incentives for agents to seek out activities that minimize their tax burden, substituting away from high-taxed activities to those taxed at lower rates. By inducing this substitution, agents will engage in less productive activity, leading to lower rates of economic growth (Myles, 2000). Thus, holding constant the overall

tax burden, it is possible to obtain higher levels of output by shifting away the tax structure from income taxes towards non-distorting consumption taxes.
The tax policy in Cote d'Ivoire has been scrutinized and revised continuously since 1960. The country has undertaken a series of reforms in its tax system aiming at increasing tax revenues and promoting economic growth. Some of the reforms intended to extend the tax base, reduce exemptions and improve the collecting system by decentralizing the fiscal administration and eliminating fraud[1]. Despite these reforms, the overall tax rate shows a downward trend, declining from 21.6% in 1965 to 17.5% in 1990 and 15% in 2006. The tax performance does not meet the requirement of convergence criteria that target a level of tax revenues exceeding 17% of GDP[2]. Over the same time period, the share of indirect taxes in total tax revenues has fallen, declining from 85.5% in 1965 to 72.4% in 1990 and 70% in 2006. In the eyes of some observers, a reduction in direct taxes can help reducing tax avoidance[3] and improving tax burden as well as economic growth. With respect to economic performance, the country enjoyed two decades (1960-1980) of good economic performance and entered in a long period of economic crisis. Domestic adjustment strategies pursued during the 1980s failed to boost economic activity and to close all deficits. As a necessary response to the failure of macroeconomic policies, the country experienced the devaluation of its currency on January 11, 1994. The devaluation accompanied by structural reforms led to an encouraging recovery: economic performance has strengthened significantly from 1994 and budget deficits fell. But all will change on December 24, 1999, when rebels overthrew the government in the country's first military coup. Since that time, the political agenda of Côte d'Ivoire has been dominated by political and social tensions. Today, government is looking for revenues to rebuild the country's infrastructures and boost the economic activity.

Our objective in this paper is to contribute to the tax-growth literature by examining the case of Côte d'Ivoire looking in particular if there is any evidence that taxation plays a role in explaining the process of economic growth. More precisely, the study addresses the following questions. How are taxes related to economic output? Which component of the tax structure – the tax burden or the tax mix –has the more potent influence on the real GDP? What is the effect on real GDP of moving from indirect to direct taxes? To provide answers to these questions, we use annual time-series data for the period 1960 to 2006.

The remainder of the paper is organized as follows. Section 2 outlines the methodological framework. We first use Data Envelopment Analysis to isolate the influence of non-tax variables on economic growth. Next, we use econometric methods to estimate the separate growth effects of the tax burden and the tax mix. We undertake a thorough investigation of the unit root properties of the data. To this end, apart from using standard unit root tests, we also employ the Zivot and Andrews (1992) test. The appealing aspect of this test is that it allows one to establish the unit root properties of the data in the presence of one structural break. An important contribution here is that we calculate critical values specific to our

[1] An overview of a chronology of fiscal reforms implemented in Côte d'Ivoire from 1960 to 2006 can be found in "*Code Général des Impôts, Livre de procedures fiscales, Autres textes fiscaux, 2007*", Direction Générale des Impôts, Côte d'Ivoire.
[2] Côte d'Ivoire is member of the West African Economic and Monetary Union. This union has adopted in 1994 convergence criteria aiming at explicit targets for inflation, public debt and deficits to monitor the fiscal situation of the member countries. To meet the convergence criteria, the member countries should, among others, increase tax revenues over 17% of GDP and keep public deficit at a minimum of zero percent of GDP.
[3] Fiscal fraud amounts to 500 trillions of FCFA, some 31% of total tax revenue.

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sample size. We next apply the residual-based cointegration test developed by Gregory and Hansen (996) to examine the existence of a long-run relationship between the variables under the hypothesis of structural break. Section 3 presents and discusses the empirical results of the study. Finally, Section 4 summarizes the main findings.

2. Methodology

2.1 Model specification

Traditional growth accounting approach based on aggregate function originally developed by Solow (1956) is still the most widely used method for establishing factors influencing growth of countries. The approach has been extensively revised to incorporate human capital (Lucas, 1988) and public spending (Barro, 1990). In its simplest form, an aggregate production function is described as follows:

$$\log(Y_t) = \gamma + \alpha \log(K_t) + \beta \log(L_t) + \mu_t \quad (1)$$

where Y=aggregate output, K is the economy's capital stock; L is the labor force, and α and β are the elasticities of real GDP with respect to K and L. As Engen and Skinner (1996) explain, tax policy directly and indirectly affects the aggregate output through all variables on the right side of Eq. (1). Income, business and consumption taxes can alter the incentives to invest in physical and human capital, and therefore altering the growth rates of human and physical capital, as well as technical progress. Furthermore, tax policy can also influence the relative cost of physical and human capital and research and development expenditures, and thereby influencing input elasticities for human and physical capital and productivity growth.

We consider two separate measures of tax policy. One is the tax burden defined as the ratio of tax revenue to GDP (*b=(I+D)/Y*) and the other is the tax mix defined as the ratio of indirect tax to direct tax revenue (*m=I/D*). Since these tax variables influence all the variables on the right hand side of Eq. (1), we replace the production function Eq. (1) with:

$$Y_t = f(b_t, m_t, Z_t) + \mu_t \quad (2)$$

where Z_t is a vector of other non-tax economic growth determinants and μ_t represents an error term.

The empirical problem with Eq.(2) is that many of the non-tax variables in Z cannot be observed. If the variables in Z were uncorrelated with the tax variables, we could estimate Eq. (2) without concern for bias, even if we ignore Z and treat it as a vector of omitted variables (see Frisch-Waugh theorem). However, following the arguments of Engen and Skinner (1996), the assumption of no correlation between the non-tax variables in Z and the tax variables (tax burden and tax mix) is not reasonable. One approach for dealing with this problem is to use instrumental variables techniques. However such an option is made difficult by the selection of appropriate instruments. Also, parameter estimates are likely to be sensitive to the selection of instruments. To overcome these difficulties an alternative approach based on Data Envelopment Analysis (DEA) is adopted in this study. The objective of this linear programming model is to isolate the influence of the unobserved factors in Z on GDP prior to the estimation of Eq. (2).

Data Envelopment Analysis is a linear programming-based methodology to measure the efficiency of multiple decision-making units (DMUs) when the production process presents a structure of multiple inputs and outputs (see Charnes *et al.*, 1994; Ramanathan, 2003; Ray, 2004). It aims to measure how efficiently a unit uses the resources available to generate a set of output. Efficiency is defined as a weighted sum of outputs to a weighted sum of inputs. As its name suggests, DEA envelops the data so that observations on the "edge of the envelope" represent economic frontiers. The "edge of the envelope" is used to determine how far the remaining observations are from the frontier using a simple scaling factor. Assuming that there are *n* DMUs, each with *m* inputs and *s* outputs, the relative efficiency score of a test DMU *p* is obtained by solving the following model:

$$
\begin{aligned}
& Max_{u,v} \frac{\sum_{k=1}^{s} v_k y_{kp}}{\sum_{j=1}^{m} u_j x_{jp}} \\
& s.t. \quad \frac{\sum_{k=1}^{s} v_k y_{ki}}{\sum_{j=1}^{m} u_j x_{ji}} \leq 1, \forall i \\
& \quad v_k, u_j \geq 0 \quad \forall k, j
\end{aligned}
\tag{3}
$$

Where y_{ki} is the amount of output *k* produced by DMU *i*, x_{ji} is the amount of input *j* utilized by DMU *i*, v_k is the weight given to output *k*, and u_j the weight given to input *j*. One problem with this fractional program is that it has an infinite number of solutions. To avoid this one can impose the constraint $\sum_{j=1}^{m} u_j x_{jp} = 1$ and the programme becomes:

$$
\begin{aligned}
& Max_{u,v} \sum_{k=1}^{s} v_k y_{kp} \\
& s.t. \quad \sum_{j=1}^{m} u_j x_{jp} = 1 \\
& \sum_{k=1}^{s} v_k y_{ki} - \sum_{j=1}^{m} u_j x_{ji} \leq 0 \quad \forall i \\
& v_k, u_j \geq 0 \quad \forall k, j
\end{aligned}
\tag{4}
$$

The above problem is run *n* times in identifying the relative efficiency scores of all the DMUs. Each DMU selects input and output weights that maximize its efficiency score. In general, a DMU is considered to be efficient if it obtains a score of 1 and a score of less than 1 implies that it is inefficient. The dual problem associated with (4) is:

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$$
\begin{aligned}
& Min_{\theta,\lambda}\theta \\
& s.t. \quad \sum_{i=1}^{n}\lambda_i x_{ji} - \theta x_{jp} \leq 0, \forall j \\
& \qquad \sum_{i=1}^{n}\lambda_i y_{ki} - y_{kp} \geq 0 \quad \forall k \\
& \qquad \lambda_i \geq 0 \quad \forall i \\
& \qquad \sum_{i=1}^{n}\lambda_i = 1
\end{aligned} \tag{5}
$$

where θ_t is the efficiency score, and λ_is are dual variables.

In this study, two input variables (*m=2*) and one output variable (*s=1*) are considered for efficiency measurement. Input variables include the ratio of GDP to direct taxes (Y/D) and the ratio of GDP to indirect taxes (Y/I). The output variable is real output in index numbers, i.e. $GDP_t / \max(GDP_t)$. Thus, the linear program seeks the smallest reciprocal tax burden or, equivalently, the heaviest tax burden, which is consistent with the observed real GDP, given a history of observed tax burdens and real GDP in the economy over the time period. The linear programme to be solved is:

$$
\begin{aligned}
& Min_{\theta,\lambda}\theta \\
& s.t. \quad \sum_{t=1}^{T}\lambda_t\left(\frac{Y_t}{D_t}\right) - \theta\left(\frac{Y_p}{D_p}\right) \leq 0 \\
& \qquad \sum_{t=1}^{T}\lambda_t\left(\frac{Y_t}{I_t}\right) - \theta\left(\frac{Y_p}{I_p}\right) \leq 0 \\
& \qquad \sum_{t=1}^{T}\lambda_t y_t - y_p \geq 0 \\
& \qquad \lambda_t \geq 0 \quad \forall t \\
& \qquad \sum_{t=1}^{T}\lambda_t = 1
\end{aligned} \tag{6}
$$

As explained earlier, the programme given by (6) is solve T=47 times, once for each year in the data. A value of θ is then obtained for each year. For year *p* the program tries to find the largest increase in the indirect and direct tax burden consistent with the constraints. The first two constraints require that the increase in the direct (indirect) tax burden, as measured by the reciprocal direct (indirect) tax share of income, cannot exceed a linear combination of all other years' tax burdens. The third constraint requires that a linear combination of all other years' growth rates cannot be exceeded by the growth rate in year *p*. The final T+1 constraints force the linear combinations to be convex with non-negative weights.

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The solution θ_t is interpreted as a normalised proxy for the unobserved non-tax influences Z_t on economic output and used to filter GDP fluctuations out of the system. Thus, the estimable equation becomes:

$$\log(y_t) - \log(\theta_t) = \log(\tilde{y}_t) = \phi_0 + \phi_b b_t + \phi_m m_t + \phi_{bm} b_t \times m_t + \mu_t \quad (7)$$

We are interested in how taxation affects GDP. Taking the differential of Eq. (7) we obtain the time-varying long-run effect of the tax mix:

$$\frac{\partial \log(\tilde{y}_t)}{\partial m_t} = \phi_m + \phi_{bm} b_t \quad (8)$$

Thus, Eq.(7) allows the effect of tax mix to depend on the tax burden. If ϕ_m is positive and ϕ_{bm} is negative, the effect of tax mix on GDP falls as tax burden increases. The appropriate threshold value of tax burden would be $b^* = -\frac{\phi_m}{\phi_{bm}}$. When $b_t \geq b^*$, a tax mix that places excessive emphasis on indirect taxes relative to direct taxes has a negative effect on GDP. Of course, if ϕ_m and ϕ_{bm} are both positive (negative), then tax mix would have an unambiguously positive (negative) real effect on GDP, that increases (decreases) in magnitude with the level of the tax burden. A zero or insignificant interaction would indicate that the effects of the tax burden and the tax mix on GDP are independent. This would mean that the effect of changes in tax mix on GDP is not influenced by the level of tax burden.

Testing for Unit Root

Eq. (7) represents the long-run equilibrium relationship and may form a cointegration set provided all the variables are integrated of order one. Before carrying out the estimation of this equation, we have to determine the order of integration of each variable. For purposes of unit root testing, most existing empirical studies apply the well known Augmented Dickey and Fuller (ADF, 1979) and Phillip and Perron (PP, 1988) unit root tests. However, it is well documented that these standard tests are biased towards the nonrejection of the null hypothesis of a unit root in the presence of structural changes, thus using these tests could produce spurious results (Perron, 1989; Zivot and Andrews, 1992). For this reason, we apply the Zivot and Andrews (ZA, 1992) one-break test, which identifies possible periods of change in the time series based upon a series of dummy variables. Like other unit root tests (see Banerjee et al., 1992; Perron and Vogelsang, 1992; Lumsdaine and Papell, 1997; Perron, 1997; Saikkonen and Lutkepohl, 2002), this test has the advantage of not requiring the *a priori* specification of the possible timing of structural breaks. It allows the break date to be endogenously determined within the model. We use two versions of the ZA test. Model A allows for a change in intercept, while model C allows for a change in both the intercept and slope. Model A involves running the following regression:

$$\Delta H_t = \mu + \beta t + \theta DU_t + \alpha H_{t-1} + \sum_{t=1}^{k} c_j \Delta H_{t-j} + e_t \quad (9a)$$

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Model C takes the following form:

$$\Delta H_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \alpha H_{t-1} + \sum_{t=1}^{k} c_j \Delta H_{t-j} + e_t \qquad (9b)$$

where Δ is the first difference operator, DU_t and DT_t are dummy variables for a mean shift and a trend shift defined as: $DU_t = 1$ if $t > T_b$ and 0 otherwise; $DT_t = t - T_b$ if $t > T_b$ and 0 otherwise. The *k* extra regressors are included to address the problem of autocorrelation in the error term e_t. A test of the unit root hypothesis has the null $\alpha = 0$. The alternative hypothesis is that the series H_t is trend stationary with a structural break in the trend function. The searching for breakpoint (T_b) is performed by running a set of regressions and by choosing the breakpoint for which the *t*-statistic t_α for α is minimized. The lag length *k* is selected using the general-to-specific approach proposed by Perron (1989), i.e. we use a critical value of 1.60 to determine the significance of the *t*-statistic on the last lag. Given that our sample size is relatively small (47 observations) we set k_{max}=4. Whilst asymptotic critical values are available for this test, Zivot and Andrews (1992) warn that with small sample sizes the distribution of the test statistic can deviate substantially from this asymptotic distribution. To circumvent this problem, they suggest a Monte Carlo method to calculate exact critical values (see Zivot and Andrews, 1992, p. 262). Following this methodology, we estimate an *ARMA(p,q)* model for each ΔH_t, with *p* and *q* selected according to the Akaike Information Criterion (AIC). The implied ARMA process is then used as the data generating process for generation of 5000 sample specific series under the null hypothesis of a unit root with no structural break.

Testing for Cointegration

There exist several methods for testing for cointegration between two or more variables. The residual-based test of Engle and Granger (1987) and the system-based approach pioneered by Johansen (1988) are two widely used econometric tools for cointegration analysis. Although these approaches are well documented in the empirical literature, they are not immune to criticism. In particular, the power of these standard tests of cointegration may be substantially reduced when applied to series which experience structural change in their long-run cointegrating relationship. We therefore apply the Gregory and Hansen (1996,a,b) test for cointegration that allows for structural breaks in the cointegrating vector. Applying the similar approach by Zivot and Andrews (1992), the Authors revise the Engle and Granger model to consider the regime shift via residual-based cointegration technique. The technique is to test the null hypothesis of no cointegration against the alternative of cointegration with regime shifts in the trend as well as the slope coefficients. Gregory and Hansen (1996a, b) presented four models for testing cointegration:

Model C: Level shift: $z_t = \mu_1 + \mu_2 \varphi_t + \alpha x_t + e_t$ (10a)

Model C/T: Level shift with trend: $z_t = \mu_1 + \mu_2 \varphi_t + \beta t + \alpha x_t + e_t$ (10b)

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Model C/S: Regime shift: $z_t = \mu_1 + \mu_2\varphi_t + \alpha_1 x_t + \alpha_2\varphi_t x_t + e_t$ (10c)

Model C/S/T: Regime and Trend Shift: $z_t = \mu_1 + \mu_2\varphi_t + \beta_1 t + \beta_2 t\varphi_t + \alpha_1 x_t + \alpha_2\varphi_t x_t + e_t$ (10d)

where z_t is the dependent variable and $x_t = (b_t, m_t, b_t \times m_t)$ the vector of regressors, φ_t is the dummy variable which introduces the structural change, and defined as $\varphi_t = 0$ if $t \le \tau$, and $\varphi_t = 1$ otherwise; τ denotes the timing of the change point. In the general model C/S/T, μ_1, β_1 and α_1 represent the cointegrating coefficients before the regime shift, and μ_2, β_2 and α_2 denote the changes in the coefficients at the time of the shift. As in Zivot and Andrews tests, τ is determined using a grid search procedure, with all values in the central 70% of the sample being considered. For each value of τ, the above models are estimated with the resulting residuals $\hat{e}_t(\tau)$ saved and employed to compute the $ADF(\tau)$ statistic. The break point is determined by finding the minimum values for the $ADF(\tau)$ statistics across all possible breaks. Asymptotic critical values are provided in Gregory and Hansen (1996a, b).

DATA AND EMPIRICAL RESULTS

The data used for the empirical analysis are annual and cover the period from 1960 to 2006. Data on total tax revenues and its breakdown in direct and indirect taxes are obtained from the National Institute of Statistic and those on GDP are from 2010 world development indicators of the World Bank (WDI, 2010). Real GDP has been taken in index numbers by dividing each value by the maximum.

As explained earlier, before we proceed to cointegration tests, we have to test for the order of integration of the variables under consideration. This is to ensure that the variables used in the regressions are not subject to spurious correlation. Furthermore the Gregory and Hansen (1996) test is applicable for I(1) processes. To ascertain the order of integration of the series, we begin through applying the unit root tests of Dickey and Fuller (1979), Phillips and Perron (1988) and Elliott *et al.* (1996), denoted as ADF, PP and DF-GLS respectively. The DF-GLS test is a simple modification of the conventional augmented Dickey-Fuller (ADF) *t*-test as it applies generalized least squares (GLS) detrending prior to running the ADF test regression. Compared with the ADF tests, the test has the best overall performance in terms of sample size and power (Elliott *et al.*, 1996). The tests are applied to the level variables as well as to their first differences under the models with constant and with constant and trend. The results reported in Table 1 are mixed. ADF and PP tests suggest that m_t is stationary in level whereas the DF-GLS test supports that it is not. Under the model with constant and trend, all tests indicate that $\tilde{y}_t$ is stationary. With respect to b_t, the three tests unanimously indicate non-stationarity regardless of the model used for the test. However, when applied to the first differences of the series, all tests suggest stationarity.

Table 1: Standard tests for unit root

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Series	Level			First-difference		
	ADF	PP	DF-GLS	ADF	PP	DF-GLS
Model I: drift and no trend						
$\tilde{y}_t$	-2.358	-2.187	-2.271*	-7.290*	-22.265*	-7.587*
b_t	-1.892	-1.567	-1.938	-7.269*	-11.548*	-6.958*
m_t	-4.046*	-5.195*	-0.587	-6.504*	-6.514*	-5.408*
Model II: drift and trend						
$\tilde{y}_t$	-4.286*	-4.277*	-4.105*	-7.213*	-21.693*	-6.865*
b_t	-3.306	-3.155	-2.897	-7.359*	-21.298*	-7.033*
m_t	-3.977*	-4.672*	-1.990	-6.763*	-6.935*	-6.734*
Critical values 5%						
Model I	-2.926	-2.926	-1.948	-2.929	-2.928	-1.948
Model II	-3.510	-3.510	-3.190	-3.515	-3.513	-3.190

Notes: ** (*) denotes rejection of the null hypothesis at the 10% (5%) level.

The power of the standard unit root tests may be substantially reduced when applied to series which experience structural breaks. As before mentioned, Cote d'Ivoire experienced a series of tax reforms since 1960 in response to persistent budget deficits and low economic performance. These reforms have been intensified with the adoption of adjustment programmes in 1981, but also since 1994 with the adoption of convergence criteria[4]. Hence, there might be some possibility of structural breaks in the series originating from these events that can render unit root and cointegration tests biased. Therefore, the robustness of the results obtained from standard tests is checked by employing the ZA sequential one-break unit root test. The results are reported in Table 2. As can be seen, results indicate that all series exhibit behaviour consistent with unit root non-stationarity. Thus, the series $\tilde{y}_t$ and m_t become non-stationary when a break is allowed. The fact that results for $\tilde{y}_t$ and m_t are not consistent with those obtained from the ADF, PP and DF-GLS tests suggests that regime shifts in the series are highly significant. The significance of the estimated coefficients on the post-break constant dummy (θ) or the post-break slope dummy (γ) provides evidence that at least one structural break in the intercept and slope has occurred during the sample period.

[4] See footnote 2.

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Table 2: Zivot and Andrews test for unit root

Coefficients	GDP ($\tilde{y}_t$)		Tax burden (b_t)		Tax mix (m_t)	
	Model A	Model C	Model A	Model C	Model A	Model C
Tb	1988	1988	1987	1987	1990	1990
α	-0.904	-0.967	-0.548	-0.906	-0.378	-0.373
	(**-6.370**)	(**-6.352**)	(**-4.820**)	(**-6.152**)	(**-4.524**)	(**-4.128**)
β	0.001	0.0006	0.0001	0.0006	-0.037	-0.036
	(0.968)	(0.419)	(0.337)	(1.980)	(-2.609)	(-2.181)
θ	0.187	0.179	-0.028	-0.037	0.570	0.585
	(4.012)	(3.798)	(-3.341)	(-4.248)	(1.983)	(1.920)
γ	-	0.004	-	-0.002	-	0.027
		(1.121)		(-2.630)		(-0.163)
Lag length (k)	0	0	0	0	0	0
Exact critical values						
1%	-7.599	-7.756	-6.994	-7.498	-6.174	-6.484
5%	-6.736	-6.995	-6.353	-6.727	-5.468	-5.836
10%	-6.340	-6.660	-6.047	-6.358	-5.167	-5.483

Note: The null hypothesis is that the series is an integrated process without a break, the alternative hypothesis is that the series is stationary about a broken trend. The lag length *k* is selected using the general-to-specific approach proposed by Perron (1989) with a maximum lag length set equal to k_{max}=4. Critical values are calculated from simulation with 5000 replications following the procedure described in Zivot and Andrews (1992).

Given that all the variables are I(1), we can now proceed to testing for the presence of a long-run relationship between them. Before we test for cointegration with structural break, we test for parameter stability of the long-run relationship using the three statistics provided by Hansen (1992) — SupF, MeanF and L_c—which have the null hypothesis that the parameters are stable. When calculated probability values are greater than 0.05 then the null hypothesis is accepted. Results reported in Table 3 reject the null hypothesis at the 5% level, implying instability of the long-run parameters in our model.

Table 3: Hansen (1992) tests for parameter stability

Test statistic	Model	
	Without trend	With trend
SupF	3332.339*	3917.115*
	(0.010)	(0.010)
MeanF	587.499*	769.910*
	(0.010)	(0.010)
LC	20.794*	24.422*
	(0.010)	(0.010)

Note: Figures in parentheses are *p*-values. * denotes that the null hypothesis of parameter stability is rejected at the 5% level.

We apply the Gregory and Hansen test to accommodate a structural break in the long-run relationship. Results of Table 4 indicate that the no cointegration null hypothesis can be rejected under any of the four models. The break dates of 1989 and 1991 correspond with the

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dates of implementation of adjustment programmes in Cote d'Ivoire (see Kanbur (1990) and Demery (1994) for more detail on these programmes).

Table 4: Gregory and Hansen Cointegration Test Results

Model	Test statistic ADF*	Break point Tb	Critical values	
			5%	10%
Level shift (C)	-5.654*	1989	- 4.92	-4.69
Level shift with trend (C/T)	-5.527*	1989	- 5.29	-5.03
Regime shift (C/S)	-7.861*	1991	- 5.50	-5.23
Regime and Trend Shift C/S/T	-8.046*	1991	-5.96	-5.72

Note: Critical values are obtained from Gregory and Hansen (1996a: 109) and Gregory and Hansen (1996b: 559).* denotes that the null hypothesis of no-cointegration is rejected at the 5% level.

Formal hypothesis testing in regard to the value of the cointegrating parameters cannot be directly carried out because the estimated standard errors are not consistent. For that reason, we use the three-step estimation procedure suggested by Engle and Yoo (1987) to correct the estimates and standard errors of the long-run relationship. Results are reported in Table 5. The long-run coefficient on the interaction term is negative and significant. This implies that while the effect of tax mix on GDP at low levels of tax burden is positive and significant, as tax burden increases, the effect of tax mix on GDP diminishes[5]. Also, the relation between tax burden and GDP is contingent on tax structure. The effect of tax burden on GDP decreases with the tax mix. Overall these findings provide evidence that what matters for growth is not only the level of taxes but also the way in which tax components are designed.

[5] To help with the interpretation of these results, recall that we are interested in $\frac{\partial \log(\tilde{y}_t)}{\partial m_t} = \hat{\phi}_m + \hat{\phi}_{bm} b_t$. The estimated long-run coefficient on tax mix (m) equal to 0.031+0.426=0.457 captures the effect of tax mix on GDP when tax burden approaches zero, while the coefficient on the interaction term (-0.293-2.304=-2.597) is capturing the degree that the effect of tax mix (held constant) changes as tax burden (not constant) moves away from zero.

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Table 5: OLS parameter estimates and summary statistics

Variable	Coefficient	T-ratio
Constant	9.470*	62.385
b_t	-3.570*	-6.242
m_t	0.031	1.140
$b_t \times m_t$	-0.293*	-2.039
Dum_t	-1.394*	-5.029
$b_t \times Dum_t$	7.740*	4.417
$m_t \times Dum_t$	0.426*	5.991
$b_t \times m_t \times Dum_t$	-2.304*	-5.167
Log Likelihood	114.672	
R^2 Adjusted	0.971	
AIC	-4.539	
DW statistic	2.287	

Note: * denotes statistical significance at 5%.

Table 6 presents the range of computed growth elasticities with respect to the tax burden and the tax mix. The tax burden elasticity is negative in every year and has a mean value of –4.547 implying that each 1 percentage point increase in the tax burden leads to a 4.547% decrease in real GDP. The elasticity of economic output with respect to the tax mix has a mean value of -0.003, implying that a 5 percentage point increase in the tax mix will result in a 0.015% decrease in real GDP. Thus, over the period, increases in the tax burden have been much more damaging than have increases in the tax mix.

Table 6: Tax burden and tax mix elasticities

Statistic	Tax burden	Tax mix (I/D)
Mean	-4.547	-0.003
Median	-4.568	-0.025
Minimum	-7.227	-0.052
Maximum	-1.974	0.088

To get an estimate of how important the tax structure has been in enhancing the growth effects of tax burden, one can ask the hypothetical question of how much a one standard deviation decrease in the tax mix would enhance real GDP, holding constant the mean level of tax burden in the sample[6]. It turns out that adopting tax mixes that place more emphasis on direct taxes relative to indirect taxes would allow the economy to experience a 4% increase in real output, where the net effect being measured is $\phi_m \times \sigma_m + \phi_{bm} \times mean(b_t) \times \sigma_m$. Considering the mean level of tax burden in the sample (18.68% of GDP), our estimates suggest that a decrease from the average to the minimum level of the tax mix in the sample

[6] The mean value for tax burden is 18.68% in the sample. The standard deviation of tax mix is equal to 1.44.

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corresponds to a 4.4% increase in GDP[7]. We can derive the threshold value of tax burden that allows the effect of tax mix to shift. Taking differentiation of the estimated equation with respect to the tax mix, m_t, and solving for b_t yields $b^* = 17.542\%$. For $b_t \leq 17.542\%$ the growth effect of tax mix is positive with an average elasticity of 0.051, while for $b_t > 17.542\%$ the tax mix elasticity is -0.028. This suggests that up to a tax burden of 17.572%, increased indirect tax'share in total tax revenue has a growth-enhancing effect, but beyond this level increased indirect tax relative to direct tax has a negative externality on the economic activity. It is worth noting that the threshold value found here is very close to the level of 17% recommended by the West African Economic and Monetary Union. The actual tax burden is 15.05%, nearly 20% beneath the period mean observed value. The actual tax mix is 2.32, some 38% beneath the period mean value. Holding constant the actual level of tax burden, a shift from direct to indirect taxes will lead to higher levels of output. A one-off shift from the actual tax structure to a tax burden exceeding 17.54% would reduce the level of GDP unless the tax mix increases sufficiently to offset the adverse effect.

CONCLUSION

This study examines the long-run effect of the structure of taxes on aggregate economic activity in Cote d'Ivoire over the period 1960 to 2006. Prior to estimation, the analytical framework captures the influence of non-tax variables on output through Data Envelopment Analysis. The existence of a cointegration relationship between real GDP, tax burden and tax mix has been detected following the residual-based test of Gregory and Hansen (1996). The results reveal that tax policy has not been pro-growth over the sample period. Tax burden and tax mix are negatively associated with output, with tax burden having a much greater adverse effect on GDP than has a tax mix that places an excessive emphasis on indirect taxes. This suggests that any negative effect from an increase in the share of indirect taxes should be more than offset by the positive effect associated with a reduction in the overall tax burden. The effect of the tax mix on GDP is contingent on the level of the tax burden and diminishes as tax burden increases. The estimates also suggest that up to a threshold level of tax burden of 17.57%, increased direct taxation relative to indirect taxation is associated with decreased output. But beyond this threshold a move from indirect to direct taxes is likely to lead to higher levels of output.

These findings provide evidence that tax structure matters for long-run level of real GDP. They show that reducing the overall tax burden is a more potent way of enhancing economic output than is fine-tuning the share of indirect taxes in total tax revenue. More output can be produced without increasing the tax burden. Thus, holding constant the actual tax burden at its level of 15%, a shift from direct to indirect taxes will increase the level of aggregate output. As tax policy has not been growth-promoting, government should try to return taxes back to the economy in an efficient manner so that they contribute to enhance economic growth and expand the tax base.

[7] The mean observed tax mix for the period is 3.815 (a tax mix comprised of 21% direct tax and 79% indirect tax) and the minimum value is 2.22 (a tax mix comprised of 31% direct tax and 69% indirect tax).

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