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Government Revenue-Expenditure Nexus: Evidence from Latin America

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We utilize the Engle-Granger bivariate cointegration approach to test several hypotheses concerning the temporal relationship between revenues and expenditures relative to real GDP. In the case of Chile and Paraguay we find evidence of bi-directional causality between revenues and expenditures supporting the fiscal synchronization hypothesis. Under this scenario the fiscal authorities of Chile and Paraguay should try to raise revenues and cut spending simultaneously in order to control their respective budget deficits. For Colombia, Ecuador, and Guatemala we find evidence of casuality from revenues to expenditures thus supporting the tax-spend hypothesis. Under this scenario the fiscal authorities of Colombia, Ecuador, and Guatemala should focus attention on adjusting revenues in order to control spending and the size of budget deficits.

I. Introduction

The macroeconomic ramifications of budget deficits have been widely discussed in the literature. Within the context of economic growth there seems to be a consensus that persistent budget deficits have adverse effects on a country's national savings and investment which could slow a country's growth potential. Understanding the relationship between government spending and taxation is important in evaluating the government's role in the distribution of resources. The focus of this paper is to examine the intertemporal relationship between government revenues and government expenditures for a sample of five Latin American countries.[1] For many less developed countries deficit financing via borrowing or money creation have greatly affected the saving and growth prospects of these countries. Thus, many developing countries must undergo fiscal restraint to control the growth of budget deficits (Edwards (1995)). The bulk of the research in this area has dealt with industrialized countries within the G7 and OECD with the exception of studies by Ram (1988b) and Baffes and Shah (1990, 1994).

The discussion of the causal link between revenues and expenditures has resulted in several hypotheses. The tax-spend hypothesis suggests that changes in revenues induce changes in expenditures. The spend-tax hypothesis suggests the opposite in that changes in expenditures induce changes in revenues. The fiscal synchronization hypothesis argues that revenue and expenditure decisions are made jointly. Another view relates to the institutional separation

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1. Fishlow (1990) provides an excellent overview of the debate on the role of the state in Latin American economic development.

of the expenditure and taxation decisions of government. This perspective suggests that revenues and expenditures are independent of one another.

We wish to test the validity of these hypotheses in the case of five Latin American countries. Utilizing cointegration analysis and error correction models, inferences can be made concerning the respective hypotheses set forth. Section II will provide a brief overview of the hypotheses along with a review of the empirical literature on the tax-spend debate. Section III discusses the methodology and data used in the analysis. Section IV provides the empirical results while section V makes concluding remarks.

II. Literature on the Tax-Spend Debate

Several hypotheses have been set forth to describe the temporal relationship between revenues and expenditures.[2] First, the tax-spend hypothesis advanced by Friedman (1978) argues that changes in government revenues lead to changes in government expenditures. Friedman suggests that tax increases will only lead to expenditure increases resulting in the inability to reduce budget deficits. Buchanan and Wagner (1978) agree that taxes affect government expenditures but in a slightly different way. Within the Buchanan-Wagner framework, increases in government spending are due to indirect taxation. When spending is financed by other means than direct taxation the public perceives the price of government spending to be less with indirect taxation than what it would be under direct taxation. This form of indirect taxation originates through higher interest rates as a result of higher government spending (crowding out) and inflation. Buchanan and Wagner would argue that fiscal illusion results in that higher taxes lead to a decrease in government spending, opposite the result set forth by Friedman.

Second, the spend-tax hypothesis suggests that changes in government expenditures lead to changes in government revenues. Peacock and Wiseman (1979) argue that temporary increases in government expenditures due to "crises" can lead to permanent increases in government revenues often called the "displacement effect". Utilizing the Ricardian equivalence proposition Barro (1974) argues that government borrowing today results in an increased future tax liability which is fully capitalized by the public. Thus, under Barro's analysis fiscal illusion is absent in that increases in government spending lead to increases in taxes.

Third, Musgrave (1966) as well as Meltzer and Richard (1981) suggest that voters compare the marginal benefits and marginal costs of government services when formulating a decision in terms of the appropriate levels of government revenues and expenditures. Thus, revenue and expenditure decisions are jointly determined under this fiscal synchronization hypothesis. A fourth hypothesis mentioned by Baghestani and McNown (1994) relates to the institutional separation of the expenditure and taxation decisions of government. This perspective suggests that revenues and expenditures are independent of one another.

The empirical literature on the tax-spend debate has yielded mixed results due in part

2. In addition to the literature pertaining to the federal level of government there have been numerous studies at the state and local levels of government. For a more detailed discussion of this topic see the studies by Manage and Marlow (1987, 1988), Chowdhury (1988), Ram (1988a), Miller and Russek (1989), Holtz-Eakin, Newey, and Rosen (1989), Joulfaian and Mookerjee (1990a, 1990b), and Payne (1997).

to the various time periods analyzed, lag length specifications used, and methodology. Generally, the methodology used in these studies has been to test for Granger causality within a vector autoregressive model; however, some of the studies test for Granger causality within an error-correction framework.

In the case of the United States, Blackley (1986), Ram (1988a), Bohn (1991), and Hoover and Sheffrin (1992) provide evidence to support the tax-spend hypothesis while Anderson *et al*. (1986), Von Furstenberg *et al*. (1986), Jones and Joulfaian (1991) and Ross and Payne (1998) find support for the spend-tax hypothesis. Manage and Marlow (1986), Miller and Russek (1989), and Owoye (1995) suggest the fiscal synchronization hypothesis is valid for the United States while Baghestani and McNown (1994) support the institutional separation hypothesis.

In the case of Canada, the studies by Ahiakpor and Amirkhalkhali (1989) and Payne (1997) support the tax-spend hypothesis while the evidence of Owoye (1995) supports the fiscal synchronization hypothesis. Regarding the remaining G7 countries Owoye (1995) finds the tax-spend hypothesis is valid for Italy and Japan while the fiscal synchronization hypothesis is supported in France, Germany, and the United Kingdom. In the case of Greece, Provopoulos and Zambaras (1991) as well as Hondroyiannis and Papapetrou (1996) provide evidence of the spend-tax hypothesis while Katrakilidis (1997) finds evidence in favor of fiscal synchronization. Ram (1988b) examines twenty-two countries comprising both developed and less developed economies. Using constant price measures of revenues and expenditures, Ram finds support for the tax-spend hypothesis in El Salvador, Philippines, Thailand, and the United Kingdom; support for the spend-tax hypothesis in Honduras and New Zealand; and support for the fiscal synchronization hypothesis in Nicaragua. The remaining eighteen countries display an absence of causality in either direction thus lending support for the institutional separation hypothesis.[3] In a study of OECD countries, Joulfaian and Mookerjee (1991) find support for the tax-spend hypothesis in Italy and Canada; support for the spend-tax hypothesis in the United States, Japan, Germany, France, United Kingdom, Austria, Finland, and Greece; and support for the fiscal synchronization hypothesis in Ireland. Baffes and Shah (1990, 1994) have extended this analysis for Argentina, Brazil, Chile, Mexico, and Pakistan. Baffes and Shah find that for Brazil, Mexico, and Pakistan strong bi-directional causality exists between revenues and expenditures, while for Argentina and Chile expenditures appear to cause revenues.

As one can see there appears to be some disparity in the results of the studies reported. The task of this paper is to extend this line of literature to a sample of five Latin American countries which have not been examined in the literature with the exception of Chile. The following section will elaborate on the methodology to be used in this study along with a description of the data.

III. Methodology and Data

The following countries will be used in this study: Chile 1954-1993, Colombia 1950-1993, Ecuador 1951-1994, Guatemala 1958-1994, and Paraguay 1958-1993. The annual data

3. Ram's study also reports causality tests using current price data as well as various lag lengths.

are obtained from the *International Financial Statistics* CD-ROM database. For each country all variables are in real terms and converted into natural logarithms defined as follows.[4] Allowance is made for the impact of movements in real GDP by scaling the revenue and expenditure variables by real GDP.

RY	Real Government Revenues/Real GDP
EY	Real Government Expenditures/Real GDP

Given our discussion in the previous section let us briefly outline the approach taken to determine the presence of cointegration and the resulting error correction terms to be used in formulating the error correction models.

Granger (1986), Engle and Granger (1987), Engle and Yoo (1987), Johansen (1988), Stock and Watson (1987), as well as Johansen and Juselius (1990) have examined the causal relationship between two variables when a common trend exists between them. If two time series are respectively nonstationary, but some linear combination of them is a stationary process then the two time series are said to be cointegrated. A time series is said to be covariance stationary if its mean, variance, and covariances are all invariant with respect to time, in which case it is integrated of order zero, $I(0)$. If the time series requires first-order differencing to achieve stationarity, it is integrated of order one, $I(1)$. If there exists some linear combination of the two series which is $I(0)$, then cointegration is present.

In order to examine the stationarity of the respective time series in this study the following Augmented Dickey-Fuller (ADF) test was performed on each series:[5]

$$\Delta X_t = \alpha + \beta t + (\rho - 1)X_{t-1} + \sum_{j=1}^{N} \rho_j \Delta X_{t-j} + \varepsilon_t, \tag{1}$$

where Δ is the first difference operator; t is a linear time trend; ε_t is a covariance stationary random error and N was determined by the Schwarz criterion to ensure serially uncorrelated residuals. The null hypothesis is that X_t is a nonstationary series and is rejected if $(\rho - 1) < 0$ and statistically significant. If the respective time series are difference stationary, $I(1)$, then cointegrating regressions can be undertaken to determine whether or not linear combinations of the series are stationary.

Given the bivariate nature of our study, the Engle-Granger cointegration procedure is used next to test for the presence of cointegration between the two time series. If both time series are integrated of the same order then one can proceed with the estimation of the following cointegrating regressions. Let RY denote the respective revenue measure and EY denote the respective expenditure measure.

4. Given the lack of consistent and sufficient length of the time series data the inclusion of additional Latin American countries was prohibited.
5. Phillips and Perron (1988) use a nonparametric adjustment to the Dickey-Fuller test statistics which allows for weak dependence and heterogeneity in the error term. However, Kim and Schmidt (1990) indicate that the Phillips-Perron tests do not perform well in finite samples.

$$RY_t = \alpha_1 + \beta_1 EY_t + \mu_t \tag{2a}$$

and

$$EY_t = \alpha_2 + \beta_2 RY_t + \eta_t. \tag{2b}$$

The residuals from the above cointegrating regressions are then tested for stationarity to determine whether or not the two time series are cointegrated by using the following ADF unit root tests on the respective residuals.

$$\Delta\mu_t = \alpha_0 + \delta_1 \mu_{t-1} + \sum_{j=1}^{q} \alpha_j \Delta\mu_{t-j} + \varepsilon_{1t} \tag{3a}$$

and

$$\Delta\eta_t = \alpha_0 + \delta_2 \eta_{t-1} + \sum_{j=1}^{q} \beta_j \Delta\eta_{t-j} + \varepsilon_{2t}, \tag{3b}$$

where μ_t and η_t are the residuals from Equations (2a) and (2b); ε_{1t} and ε_{2t} represent the respective stationary random errors. The null hypothesis of nonstationarity (not cointegrated) is rejected when δ_1 and δ_2 are significantly negative. If cointegration is present the following error correction models can be used to test for Granger causality.

$$\Delta RY_t = \Omega_{10} + \sum_{j=1}^{s} \Omega_{11j} \Delta RY_{t-j} + \sum_{j=1}^{q} \Omega_{12j} \Delta EY_{t-j} + \delta\mu_{t-1} + \varepsilon_{1t} \tag{4a}$$

and

$$\Delta EY_t = \Omega_{20} + \sum_{j=1}^{s} \Omega_{21j} \Delta EY_{t-j} + \sum_{j=1}^{q} \Omega_{22j} \Delta RY_{t-j} + \lambda\eta_{t-1} + \varepsilon_{2t}, \tag{4b}$$

where ΔRY_t and ΔEY_t are first-difference stationary and cointegrated with μ_{t-1} and η_{t-1} representing the lagged values of the error terms from the cointegrating regressions given by Equations (2a) and (2b). From Equation (4a) the null hypothesis that ΔEY_t does not Granger cause ΔRY_t is rejected either if the coefficients Ω_{12j}'s are jointly significant, or if the coefficient on the error correction term is significant. If the coefficient δ is significant, then the null hypothesis of no long-run equilibrium relationship can be rejected. Likewise, from Equation (4b) the null hypothesis that ΔRY_t does not Granger cause ΔEY_t is rejected either if the coefficients Ω_{22j}'s are jointly significant, or if the coefficient on the error correction term is significant. If the coefficient λ is significant, then the null hypothesis of no long-run equilibrium relationship can be rejected. One can interpret the lagged changes in the independent variables in Equations (4a) and (4b) as representing the short-run causal impact while the error correction terms provide the adjustments of ΔRY_t and ΔEY_t towards their respective long-run equilibrium.

IV. Empirical Results

Table 1 presents the ADF unit root test statistics for the variables in both levels and first-differences. Based on the ADF test statistics, all variables are integrated of order one which means the respective time series are stationary in first-differences.[6] Given the respective revenue and expenditure measures are integrated of the same order we proceed to test for cointegration using the Engle-Granger bivariate methodology. Equations (2a) and 2(b) were estimated by ordinary least squares and the respective residuals were tested for stationarity via ADF unit root tests given by Equations (3a) and (3b). Table 2 displays the cointegrating regressions and ADF unit root tests.[7] The respective revenue and expenditure measures are cointegrated for all five countries.

Table 1 ADF Unit Root Tests

Country	Variables	Levels	First-differences
Chile	RY	-1.6370	-3.0127[b]
	EY	-1.7602	-3.0285[b]
Colombia	RY	-0.5903	-4.1057[a]
	EY	-0.9497	-4.1809[a]
Ecuador	RY	-0.8849	-4.2897[a]
	EY	-0.7454	-3.1626[b]
Guatemala	RY	-0.4359	-3.8149[a]
	EY	-0.1050	-3.1032[b]
Paraguay	RY	-1.2376	-3.6020[b]
	EY	-1.1759	-3.3121[b]

Finite sample critical values for the ADF unit root tests are drawn from MacKinnon (1991). a denotes significance at 1% level and b denotes significance at 5% level.

Table 2 Engle-Granger Bivariate Cointegration Tests ADF Statistics

Country	Cointegration Regression	R^2	ADF
Chile	$RY = 0.015 + 0.992\ EY$	.9996	-3.4642[b]
	$EY = 0.017 + 1.008\ RY$	.9996	-3.4589[b]
Colombia	$RY = -0.073 + 1.000\ EY$	.9935	-4.0047[a]
	$EY = 0.079 + 0.993\ RY$	.9935	-3.9628[a]
Ecuador	$RY = -0.050 + 0.950\ EY$	.9943	-4.2367[a]
	$EY = 0.060 + 1.047\ RY$	.9943	-4.2560[a]

6. One lag on the augmenting term was used. In all cases the residuals were white noise.
7. For the stationarity tests of the residuals from the cointegrating regressions the number of lags on the augmenting term was determined by choosing the lag structure (either one or two lags) that minimized Akaike's information criterion. This method of choosing the variable lag length from the data has been shown by Hall (1994) to considerably improve the ADF unit root test.

Table 2 (Continued)

Country	Cointegration Regression	R^2	ADF
Guatemala	$RY = -0.147 + 1.001\ EY$	.9780	-2.8408[c]
	$EY = 0.127 + 0.977\ RY$	.9780	-2.8477[c]
Paraguay	$RY = 0.020 + 0.956\ EY$	.9966	-5.0448[a]
	$EY = -0.021 + 1.043\ RY$	.9966	-5.0601[a]

Finite sample critical values for the ADF unit root tests are drawn from Mackinnon (1991). a denotes significance at 1% level and b denotes significance at 5% level; and c denotes significance at the 10% level.

Next, we estimate the error correction models given in Equations (4a) and (4b) to test for Granger causality. As we mentioned earlier in addition to the standard F-statistics of the coefficients on the group of lagged variables usually reported in Granger causality tests examination of the t-statistic on the respective error correction terms can also lead one to infer casuality.[8] For instance, from Equation (4a) a statistically significant coefficient on μ_{t-1} suggests that expenditures cause revenues thereby supporting the spend-tax hypothesis. Likewise from Equation (4b) a statistically significant coefficient on η_{t-1} suggests that revenues cause expenditures thereby supporting the tax-spend hypothesis. If both error correction terms from Equations (4a) and (4b) are statistically significant then the fiscal synchronization hypothesis is supported.

Tables 3A-3E present the results of the error correction models. The lag specification was determined by the Schwarz criterion. In no case were more than one lag used. The Lagrange multiplier chi-squared test for serial correlation of up to two lags was used. The error correction models were free of serial correlation via the Lagrange multiplier test, LM (2). Tests for first-order autoregressive conditional heteroscedasticity (ARCH) in the residuals were undertaken using the Lagrange multiplier chi-squared test. In no case were ARCH effects present. In the case of Chile, Table 3A, there appears to be bi-directional causality between revenues and expenditures as evident by the statistically significant lagged changes in revenue and expenditure variables along with the error correction terms. Thus, the fiscal synchronization hypothesis appears to be supported. This finding is contrary to the spend-tax results of Baffes and Shah (1990) due in part to their use of a different time period as well as the absence of cointegration between revenues and expenditures. Moreover, Baffes and Shah (1990) did not take into account movements in real GDP. In Tables 3B and 3C, Colombia and Ecuador, the error correction terms in the expenditure equations are significant lending support for the tax-spend hypothesis. Table 3D reports the results for Guatemala which support the tax-spend hypothesis given the significant coefficients on the lagged change in revenues and the error correction term in the expenditure equation. Table 3E displays the results for Paraguay. The lagged change in expenditures is significant in both the revenue and expenditure equations. Moreover, the error correction term is significant in the expenditure equation providing support for the fiscal synchronization hypothesis.[9]

8. Note that in the case where there is only one lagged value, a standard t-test replaces the joint significance test.
9. The error correction models estimated were free of structural change based on the Chow test.

Table 3A Error Correction Models Chile (t-statistics in parentheses)

Independent Variables	Dependent Variables	
	ΔRY	ΔEY
Constant	-0.187 (-2.00)	-0.192 (-2.06)
ΔRY_{t-1}	-2.518 (-2.15)[b]	-2.993 (-2.56)[a]
ΔEY_{t-1}	2.782 (2.47)[a]	3.236 (2.88)[a]
μ_{t-1}	2.372 (2.64)[a]	
η_{t-1}		-2.856 (-3.22)[a]
Adj. R^2	.2828	.3370
LM (2) [p-value]	.0281 [.9860]	.1592 [.9235]
ARCH (1) [p-value]	.0001 [.9908]	.0006 [.9805]
F-statistic	5.862[a]	7.268[a]

a denotes significance at 1% level; b denotes significance at 5% level; and c denotes significance at 10% level.

Table 3B Error Correction Models Colombia (t-statistics in parentheses)

Independent Variables	Dependent Variables	
	ΔRY	ΔEY
Constant	-0.113 (-3.84)	-0.100 (-2.88)
ΔRY_{t-1}	-0.023 (-0.10)	-0.246 (-0.93)
ΔEY_{t-1}	0.185 (0.99)	0.460 (2.09)[b]
μ_{t-1}	-0.014 (0.08)	
η_{t-1}		-0.486 (-2.51)[a]
Adj. R^2	-.0246	.1141
LM (2) [p-value]	4.557 [.1024]	.0453 [.9776]
ARCH (1) [p-value]	.1040 [.7470]	.0246 [.8755]
F-statistic	.6716	2.760[c]

a denotes significance at 1% level; b denotes significance at 5% level; and c denotes significance at 10% level.

Table 3C Error Correction Models Ecuador (t-statistics in parentheses)

Independent Variables	Dependent Variables	
	ΔRY	ΔEY
Constant	-0.053 (-1.85)	-0.033 (-1.26)
ΔRY_{t-1}	0.278 (1.10)	0.072 (0.31)
ΔEY_{t-1}	0.313 (1.42)	0.665 (3.27)[a]
μ_{t-1}	-0.067 (0.31)	
η_{t-1}		-0.488 (-2.60)[a]
Adj. R^2	.3050	.5285
LM (2) [p-value]	4.355 [.1134]	.1234 [.9402]
ARCH (1) [p-value]	1.588 [.2076]	.6295 [.4275]
F-statistic	6.997[a]	16.317[a]

a denotes significance at 1% level; b denotes significance at 5% level; and c denotes significance at 10% level.

Table 3D Error Correction Models Guatemala (t-statistics in parentheses)

Independent Variables	Dependent Variables	
	ΔRY	ΔEY
Constant	-0.068 (-2.54)	-0.043 (-1.48)
ΔRY_{t-1}	0.419 (1.61)[c]	0.509 (1.83)[b]
ΔEY_{t-1}	-0.175 (0.88)	-0.011 (0.05)
μ_{t-1}	0.004 (0.02)	
η_{t-1}		-0.331 (-1.75)[b]
Adj. R^2	.0143	.2336
LM (2) [p-value]	.9807 [.6124]	.0581 [.9713]
ARCH (1) [p-value]	.1397 [.7086]	.1150 [.7346]
F-statistic	1.164	4.454[b]

a denotes significance at 1% level; b denotes significance at 5% level; and c denotes significance at 10% level.

Table 3E Error Correction Models Paraguay (t-statistics in parentheses)

Independent Variables	Dependent Variables	
	ΔRY	ΔEY
Constant	-0.074 (-2.76)	-0.080 (-2.51)
ΔRY_{t-1}	-0.189 (-0.63)	-0.577 (-1.69)[c]
ΔEY_{t-1}	0.480 (1.70)[b]	0.858 (2.65)[a]
μ_{t-1}	-0.046 (0.13)	
η_{t-1}		-0.946 (-2.39)[b]
Adj. R^2	.1292	.1474
LM (2) [p-value]	1.596 [.4502]	1.407 [.4950]
ARCH (1) [p-value]	.9714 [.3382]	.1815 [.6701]
F-statistic	2.632[c]	2.902[c]

a denotes significance at 1% level; b denotes significance at 5% level; and c denotes significance at 10% level.

V. Concluding Remarks

This paper has attempted to extend the literature on the tax-spend debate to a sample of five Latin American countries. We utilize the Engle-Granger bivariate cointegration approach to test several hypotheses concerning the temporal relationship between revenues and expenditures relative to real GDP. In the case of Chile and Paraguay we find evidence of bi-directional causality between revenues and expenditures which lends support for the fiscal synchronization hypothesis. Under this scenario the fiscal authorities of Chile and Paraguay should try to raise revenues and cut spending simultaneously in order to control their respective budget deficits. For Colombia, Ecuador, and Guatemala we find evidence of the casuality from revenues to expenditures thus supporting the tax-spend hypothesis. Under this scenario the fiscal authorities of Colombia, Ecuador, and Guatemala should focus attention on adjusting revenues in order to control spending and the size of budget deficits.

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