XVI

Volume XVI, Number 1, May 2013

Journal of Applied Economics

Jose Angelo Divino
Luis Felipe Vital Nunes Pereira

Are increases in government spending neutral? Evidence from Latin-American Households

UCEMA

Edited by the Universidad del CEMA
Print ISSN 1514-0326
Online ISSN 1667-6726

Journal of Applied Economics. Vol XVI, No. 1 (May 2013), 101-120

ARE INCREASES IN GOVERNMENT SPENDING NEUTRAL? EVIDENCE FROM LATIN-AMERICAN HOUSEHOLDS

Jose Angelo Divino* **and Luis Felipe Vital Nunes Pereira**
Catholic University of Brasilia

Submitted December 2010; accepted May 2012

Using a dynamic optimization model, Ricardian Equivalence (RE) is empirically tested for Argentina, Brazil, Chile and Mexico. The system of equations obtained in the theoretical model is solved using Generalized Method of Moments and Full Information Maximum Likelihood. Results indicate that the null hypothesis concerning RE cannot be rejected for Argentina, Brazil, and Chile but is strongly rejected for Mexico. Therefore, when the fiscal authority seeks to stimulate economic activity by means of tax reductions and increases in government spending, the outstanding effect might be only a rise in private savings for the first three countries.

JEL classification codes: E62, H30, H60
Key words: fiscal policy, Ricardian Equivalence, public debt

I. Introduction

In an environment of recurrent economic instability, due to currency crises, changes in exchange rate regimes, confidence crises, sudden stops, and other events with important impacts on economic activity, a major concern of Latin-American policy makers is the relationship between fiscal policy and aggregate demand. Stabilization plans launched during the recent period have attributed a decisive role to fiscal policy. However, this role might not be as effective if Ricardian

---

* J. A. Divino (corresponding author): Catholic University of Brasilia, SGAN 916, modulo B, office A-116. Zip: 70.790-160. Brasilia – DF. Brazil; phone +55(61) 3448-7192; fax +55(61)3447-4797; email jangelo@pos.ucb.br. L. F. Pereira: Catholic University of Brasilia; email luis.fpereira@catolica.edu.br. The authors would like to thank participants at the XXXVII meeting of the Brazilian Economic Association, the Editor, and two anonymous referees for helpful comments and suggestions. J. A. Divino thanks CNPq for financial support. All remaining errors are the authors' sole responsibility. The views and opinions expressed in this work are those of the authors and do not necessarily reflect those of the Brazilian National Treasury.

Equivalence (RE) holds. Under RE, a temporary tax cut, for instance, would not affect personal consumption, since the increase in disposable income would be compensated by a rise in personal savings to neutralize expected increases in future taxes in order to maintain a balanced government budget.

Another implication of RE is associated to the interaction between fiscal and monetary policies. The regime of monetary policy dominance, under which fiscal policy is passive, is essentially Ricardian. In this case, the monetary authority is not forced to monetize the public debt, and is free to pursue inflation stabilization as its major policy objective. In fact, existence of Ricardian Equivalence is taken for granted in most models which seek to derivate optimal monetary policy rules.

The goal of this paper is to test if RE holds for the major Latin-American countries, namely Argentina, Brazil, Chile, and Mexico in the period from the first quarter of 1996 to the fourth quarter of 2007. These countries were chosen for their economic and political importance in the region. In addition, they have experienced distinct arrangements for their fiscal policies during the recent period and, to avoid negative effects of international crises on the domestic economies, usually follow the common practice of adopting expansionary fiscal policies. Such measures might not have the expected effects if RE holds in these countries.

The landmark on the Ricardian Equivalence literature is Barro (1974), who was the first author to model RE and state a clear hypothesis needed to confirm its validity. However, the relationship between debt issuance and taxation was first called Ricardian Equivalence by Buchanan (1976). David Ricardo believed that the government choice to issue debt or to tax is irrelevant, since debt can be viewed as a postponement of taxes. Elmendorf and Mankiw (1998) addressed the issue of public debt and its macroeconomic effects, comparing RE to the Modigliani and Miller (1958) theorem. Accordingly, corporate financing decisions are similar to government financing decisions in public sector economics. In theory, none of them matters.

Theoretically, RE requires restrictive assumptions. It requires individuals that behave as if they have infinite horizons; complete capital markets; rational and farsighted consumers; non-distortionary or lump-sum taxes; no uncertainty regarding income and future taxes; and a balanced government budget.

Some of these hypotheses have been lately relaxed, yielding restricted versions of RE. For instance, Divino and Orrillo (2008) demonstrate the validity of RE in a general equilibrium model with incomplete markets, provided that the risk-free payoff is in the asset span. Hayford (1989) shows that RE holds in the presence of liquidity constraints when default implies partial payment of debt (positive recovery value). Bassetto and Kocherlakota (2003) demonstrate that RE holds with distortionary taxes if the government is able to decide when to collect taxes.

The first empirical works on RE were based on regressions of personal consumption against fiscal variables, such as public debt and tax revenues. Rejection of RE would depend on finding statistically significant coefficients for the fiscal variables. The results, however, are contradictory, usually depending on econometric techniques, methods of collection of fiscal variables, and sample periods. Ricciuti (2003) argues that when RE is tested using life cycle models, it is usually rejected. On the other hand, dynamic optimization models tend to validate RE. Leiderman and Razin (1988) developed an intertemporal stochastic model based on Blanchard (1985) that allows to jointly test hypotheses for RE. More specifically, they test finite horizons and liquidity constraints for Israel from 1980 to 1985 with monthly data and do not reject RE.

For Latin-American countries, tests on RE are still incipient. Khalid (1996) introduced some changes in Leiderman and Razin's model and focused the analysis on 21 developing countries, including Argentina, Brazil, and Mexico. He used annual data from 1960 to 1988 and Gross National Income as a proxy for disposable income. At the 5% significance level, he did not reject RE for only 9 countries. Cuaresma and Reitschuler (2006) tested the same model for 15 OECD countries with annual data from 1960 to 2002. Their results showed deviations from RE for Finland, United Kingdom, Ireland, Luxembourg, Netherlands, Portugal and Sweden. For the other OECD countries, RE was found to hold empirically.

The main contribution of this paper is to provide empirical evidence on the existence of different consumption behaviors across the major Latin-American countries. Under a counter-cyclical fiscal policy, for instance, the common reaction of fiscal authorities across the region is to increase government expenditures as a way to stimulate economic activity in the downturn phase of the business cycle. Our results suggest that this measure might not be effective for Argentina, Brazil and Chile but might be for Mexico. Thus, there is no space for applying a single fiscal policy rule in the region.

Indeed, RE is not rejected for Argentina, Brazil, and Chile, but is strongly rejected for Mexico. Estimated parameters result in survival probabilities statistically equal to one, meaning that individuals behave as if they had infinite horizons. Tests for the presence of liquidity constraint indicate that the percentage of individuals facing liquidity restrictions is not statistically different from zero in all countries but Mexico. We find distinct rules for public and private consumption across the select Latin-America countries, meaning that it is not clear whether increasing public expenditure will crowd-out private investment. This evidence stands in sharp contrast with Khalid's (1996) results for Argentina and Brazil. The main reason is that Khalid used annual data from 1960 to 1988 without taking into account any structural break which could affect the developing countries from his sample. This misleading modeling strategy might have severely biased his results on RE.

The paper is organized as follows. In Section II, we describe the theoretical model used in the empirical evidence. The econometric procedures and results are presented and discussed in Section III. Finally, Section IV is dedicated to the concluding remarks.

II. The model

The theoretical model is drawn from Khalid (1996), who modified the framework proposed by Leiderman and Razin (1988) to accommodate testable restrictions for RE. It is an overlapping generations model with rational agents and finite horizon. There is a survival probability, γ, that does not depend on age. The probability of living for τ periods is γ^{τ}.

The consumption of an individual with no liquidity constraints, c_t^u, is given by a linear combination between public, g_t, and private, c_t, consumptions. Thus,

$$c_t^u = c_t + \sigma g_t \Rightarrow c_t = c_t^u - \sigma g_t , \tag{1}$$

where σ indicates how individuals weigh public consumption in relation to private consumption, being also understood as the degree of substitutability between public and private consumption. If σ is close to zero, then public consumption cannot substitute private consumption.

The expected utility of a consumer with no liquidity constraints is represented by:

$$E_t \sum_{\tau=0}^{\infty} (\gamma\delta)^{\tau} U(c_{t+\tau}^{u*}), \tag{2}$$

where E_t is the expectation operator conditioned on time t information set, $c_{t+\tau}^{u*}$ is the consumption of an individual with no liquidity constraints, and δ is the discount factor.

The individual maximizes (2) subject to the following budget constraint:

$$c_t^{u*} = b_t^u + y_t^u - \left(\frac{R}{\gamma}\right) b_{t-1}^u + \sigma g_t, \tag{3}$$

where b_t^u is a government bond issued to an individual with no liquidity restriction at time t, y_t^u is the disposable income and R is the risk-free interest rate, assumed to be constant.

Individuals are also subject to a non-Ponzi scheme rule:

$$E_t \lim_{t\to\infty} \left(\frac{\gamma}{R}\right)^t b_t^u = 0. \tag{4}$$

The Bellman equation can be written as:

$$V\left(y_t^u, b_{t-1}^u\right) = \underset{b_t}{Max}\, U\left[y_t^u + b_t^u - \left(\frac{R}{\gamma}\right) b_{t-1}^u + \sigma g_t\right] + \gamma \delta E_t\left[V\left(y_{t+1}^u, b_t^u\right)\right], \tag{5}$$

subject to (3).

The solution yields the following Euler equation:

$$U'\left(c_t^{u*}\right) = \delta R E_t U'\left(c_{t+1}^{u*}\right). \tag{6}$$

Following Khalid (1996), we used a quadratic utility function, implying the certain equivalence property. This assumption allows finding a linear solution for the Euler equation such as:

$$c_t^{u*} = \beta_0 + \beta_1 E_t w_t^{u*}, \tag{7}$$

where $E_t w_t^{u*}$ is the expected present value of the individual wealth, given by the sum of the present value disposable income and government consumption.[1]

[1] See Khalid (1996) for details on the derivation.

After aggregating variables and recognizing that there is a percentage θ of individuals with and (1-θ) without liquidity constraints, one finds:

$$C_t = (1-R)\beta_0 + (1-\beta_1)RC_{t-1} + (1-\gamma)(1-\theta)\beta_1 E_{t-1}H_t + (1-\gamma)\beta_1\sigma E_{t-1}S_t + \theta Y_t \sigma G_t \tag{8}$$
$$-(1-\beta_1)R\left[\theta Y_{t-1} - \sigma G_{t-1}\right]\mu_t^*,$$

where $E_{t-1}S_t$ is the expected value, based on the time t-1 information set of the discounted present value of the aggregate government consumption.

As equation (8) depends on human wealth, H_t, which is not directly observed, it is not possible to directly test its validity. In addition, one could argue that the residuals are probably correlated with Y_t. Leiderman and Razin (1988) suggested modeling Y_t and G_t by an ARIMA (1,1,0) as a way to address this issue.

Using (8) and the estimated ARIMA(1,1,0), one can find an equation which is empirically testable:

$$C_t = \lambda_0 + \lambda_1 C_{t-1} + \lambda_2 Y_{t-1} + \lambda_3 Y_{t-2} + \lambda_4 G_{t-1} + \lambda_5 G_{t-2} + v_t, \tag{9}$$

where v_t is the random error term, assumed to be homoscedastic and not autocorrelated. The compound coefficients are given by:

$$\lambda_0 = \frac{\alpha\gamma(1-R)(1-\delta R)}{\delta R(R-\gamma)}, \tag{10}$$

$$\lambda_1 = \frac{\gamma}{\delta R}, \tag{11}$$

$$\lambda_2 = \left[\theta\left(1+\rho_1 - \frac{\gamma}{\delta R}\right) + (1-\theta)(1-\gamma)\left(1-\frac{\gamma}{\delta R^2}\right)\left(\frac{R^2(1+\rho_1) - R\gamma\rho_1}{(R-\gamma)(R-\gamma\rho_1)}\right)\right], \tag{12}$$

$$\lambda_3 = \left[(1-\theta)(\gamma-1)\left(1-\frac{\gamma}{\delta R^2}\right)\left(\frac{R^2\rho_1}{(R-\gamma)(R-\gamma\rho_1)}\right) - \theta\rho_1\right], \tag{13}$$

$$\lambda_4 = \left[\left(\frac{\gamma}{\delta R} - 1 - \rho_2\right) + (1-\gamma)\left(1-\frac{\gamma}{\delta R^2}\right)\left(\frac{R^2(1+\rho_2) - R\gamma\rho_2}{(R-\gamma)(R-\gamma\rho_2)}\right)\right]\sigma, \tag{14}$$

$$\lambda_5 = \left[\rho_2 + (\gamma - 1)\left(1 - \frac{\gamma}{\delta R^2}\right)\left(\frac{R^2(1+\rho_2) - R\gamma\rho_2}{(R-\gamma)(R-\gamma\rho_2)}\right)\right]\sigma. \tag{15}$$

In case the time series are non-stationary, it is possible to rewrite (9) as an error correction model:

$$\Delta C_t = -\varphi(C_{t-1} - \theta_0 - \theta_1 Y_{t-1} - \theta_2 G_{t-1}) + \xi\Delta Y_{t-1} + \kappa\Delta G_{t-1} + v_t, \tag{16}$$

where $\phi, \theta_0, \theta_1, \theta_2, \xi$ and κ can be expressed as:

$$\theta_0 = \frac{\lambda_0}{(1-\lambda_1)}, \tag{17}$$

$$\phi = 1 - \lambda_1, \tag{18}$$

$$\theta_1 = \frac{\lambda_2 - \lambda_3}{(1-\lambda_1)}, \tag{19}$$

$$\xi = -\lambda_3, \tag{20}$$

$$\theta_2 = \frac{\lambda_4 - \lambda_5}{(1-\lambda_1)}, \tag{21}$$

$$\kappa = -\lambda_5, \tag{22}$$

which, given the system of equations (10) to (15), allows one to write $\phi, \theta_0, \theta_1, \theta_2, \xi$ as functions of the model's structural parameters.

The system of equations (10) to (15) can be solved for the structural parameters from the estimation of the reduced form equations (9) or (16). The estimation allows to directly test restrictions implied by RE. The proposition holds empirically if it is not possible to jointly reject the assumptions that the survival probability is equal to one $(\gamma = 1)$ and the percentage of individuals facing liquidity constraints is equal to zero $(\theta = 0)$. For this purpose, one can apply the Wald test.

The model's solution generates an overidentified system of equations in C_t, Y_t, and G_t. It should not be estimated by OLS as this would result in inconsistent estimators. The assumption that explanatory variables are non-stochastic is violated. Therefore, one should use alternative estimation procedures.

In the empirical estimation, we applied both Full Information Maximum Likelihood (FIML) and Generalized Method of Moments (GMM). One should be aware that the FIML estimator is based on the assumption of normally distributed residuals, which might be too restrictive in the context of Latin American countries, especially due to structural breaks in the time series. The GMM estimator does not make any assumption on the residuals distribution. However, it is also affected by the occurrence of changes in regimes. In case there are more moment conditions than parameters to be estimated, overidentifying restrictions can be tested by the Hansen (1982) statistics. In both cases, structural breaks will be appropriately modeled by including dummy variables in the estimated regressions.

III. Econometric procedure

A. Data

The data set is quarterly and ranges from the first quarter of 1996 to the fourth quarter of 2007 for Argentina, Brazil, Chile and Mexico. This is a period of relative economic stability in the region. It starts after the launch of economic stabilization plans, which resulted in the end of hyperinflation processes and the adoption of reliable fiscal and monetary policies. The sample is restricted to the fourth quarter of 2007 to avoid the influence of the international financial crisis, which started in the U.S. in early 2008. One of the major features of that crisis was the imposition of severe international liquidity constraints. As discussed in the introduction, the presence of credit restrictions could bias the results towards the rejection of RE. The time series are seasonally adjusted and expressed in local currencies deflated by each country's consumer price index (CPI). The common source, for all variables and countries, was the IMF Statistics (www.imfstatistics.org). Disposable income (Y_t) represents total individual income, excluding taxes. As in Khalid (1996), Gross National Income is used as a proxy. Private consumption (C_t) should exclude the consumption of durable goods. However, there is no such time series available. So, the series of household consumption is used as a proxy. Government expenditure (G_t), is given by government consumption expenditure. The real interest rate (R) is represented by the quarterly factor of average real interest rates. We used the Money Market interest rate and CPI to derive the real rate.

B. Unit root tests

It is well known that traditional unit root tests, primarily those based on the classic methods of Dickey and Fuller (1979, 1981) and Phillips and Perron (1988), suffer from low power and size distortions. However, these shortcomings have been overcome by modifications in the testing procedures, such as the methods proposed by Perron and Ng (1996), Elliott, Rothenberg and Stock (1996), and Ng and Perron (2001).

The modified unit root tests, labeled $MADF^{GLS}$ and MPP^{GLS}, are applied to the time series of each country. In essence, these tests use GLS de-trended data and the modified Akaike information criterion (MAIC) to select the optimal truncation lag. Asymptotic critical values for both tests are given in Ng and Perron (2001). In addition, the test by Kwiatkowski, Phillips, Schimidt and Shin (1992), labeled KPSS, which differs from the previous ones by testing the null hypothesis of stationarity instead of unit root, is performed. Critical values are provided by Kwiatkowski, Phillips, Schimidt and Shin (1992).

The results are summarized in Table 1.[2] The deterministic terms included both constant and trend or only constant whenever the trend was not statistically significant. The optimal number of lags was chosen by the MAIC, starting the search with a maximum of 10 lags. In general, the results support the conclusion that all series have a unit root, or are integrated of first order, or I(1). At least two out of the three performed tests indicate that conclusion. Still, one should be aware that the occurrence of structural breaks might affect the power of those tests leading to no rejection of the unit root null hypothesis.

Given that structural break is a common feature across Latin American time series, we also applied unit root tests which appropriately account for the occurrence of changes in regimes. The methodology suggested by Perron (1989) allows testing stationarity by modeling three different types of exogenously-selected structural breaks. Perron (1997) innovates by allowing the time of the break to be endogenously chosen in order to avoid correlation between the time series and the break period. More recently, Lee and Strazicich (2003) argued that Perron (1997) assumed no structural break under the null hypothesis and suggest a new test procedure to overcome this deficiency.

[2] Table 1 shows only the test conclusions to save space. The complete set of results is available from the authors upon request.

Table 1. Summary of the no-break unit root tests

Variable	$MADF^{GLS}$	MPP^{GLS}	KPSS
Argentina			
Personal Consumption SA	I(1)	I(1)	I(1)
Government Expenditure SA	I(1)	I(0)	I(1)
Disposable Income SA	I(1)	I(0)	I(1)
Brazil			
Personal Consumption SA	I(0)	I(1)	I(1)
Government Expenditure SA	I(1)	I(1)	I(1)
Disposable Income SA	I(1)	I(1)	I(1)
Chile			
Personal Consumption SA	I(1)	I(1)	I(1)
Government Expenditure SA	I(1)	I(1)	I(1)
Disposable Income SA	I(1)	I(1)	I(1)
Mexico			
Personal Consumption SA	I(1)	I(1)	I(1)
Government Expenditure SA	I(1)	I(1)	I(1)
Disposable Income SA	I(1)	I(1)	I(1)

Note: I(1) means that the time series is integrated of first order while I(0) indicates that it is stationary at the standard 5% significance level. SA stands for seasonally adjusted.

The results for the Perron (1989) unit root test, which considers the existence of one exogenously-chosen break, are summarized in Table 2. The time of the break was selected, for each country, according to some potentially relevant macroeconomic event. The results show that for all countries but Mexico the time series of personal consumption, government expenditure, and disposable income have a unit root.

The endogenous break selection made by the Perron (1997) test in Table 2 confirms the previous results and, in addition, suggests that personal consumption and disposable income in Mexico also have a unit root. Only government expenditure in Mexico remains as a stationary time series. The periods of the breaks, however, differ from the ones exogenously set in the Perron (1989) test.

Finally, the test by Lee and Strazicich (2003) indicated that all three time series have a unit root in each one of the four major Latin American countries. This test, which is less restrictive than Perron (1997) under the null hypothesis, confirms the conclusion by the no-break unit root tests, according to which all time series have a unit root once the structural breaks are appropriately accounted for. The structural breaks identified here will be taken to the cointegration analysis performed in the next section.

Table 2. Summary of the structural breaks unit root tests

Variable	Perron (1989)		Perron (1997)		Lee and Strazicich (2003)	
	Result	Exogenous Break	Result	Endogenous Breaks	Result	Endogenous Breaks
Argentina						
Personal Consumption SA	I(1)	2Q01	I(1)	4Q00, 2Q03 and 1Q04	I(1)	2Q03, 1Q04, 3Q99 and 4Q01
Government Expenditure SA	I(1)	2Q01	I(1)	3Q06, 3Q01 and 4Q06	I(1)	2Q99, 2Q00 and 2Q03
Disposable Income SA	I(1)	2Q01	I(1)	2Q02, 2Q01 and 2Q04	I(1)	3Q99, 3Q04, 1Q01 and 1Q05
Brazil						
Personal Consumption SA	I(1)	2Q02	I(1)	2Q06, 4Q04 and 1Q05	I(1)	4Q02, 3Q03, 1Q00 and 1Q05
Government Expenditure SA	I(1)	2Q02	I(1)	1Q06, 3Q02 and 3Q03	I(1)	3Q02, 2Q03, 4Q99 and 4Q02
Disposable Income SA	I(1)	2Q02	I(1)	1Q06, 1Q06 and 1Q98	I(1)	1Q01, 2Q06, 3Q01 and 3Q04
Chile						
Personal Consumption SA	I(1)	2Q02	I(1)	1Q98, 4Q02 and 1Q02	I(1)	3Q98, 3Q99 and 3Q02
Government Expenditure SA	I(1)	2Q02	I(1)	2Q05, 3Q02 and 1Q04	I(1)	4Q01, 3Q03, 3Q02 and 1Q06
Disposable Income SA	I(1)	2Q02	I(1)	1Q05, 4Q01 and 2Q02	I(1)	4Q05, 2Q06, 3Q01 and 3Q05
Mexico						
Personal Consumption SA	I(0)	2Q03	I(1)	1Q03, 3Q02 and 2Q02	I(1)	3Q98, 2Q03, 1Q02 and 4Q03
Government Expenditure SA	I(0)	2Q03	I(0)	4Q06, 3Q02 and 2Q03	I(1)	3Q98, 3Q01, 4Q01 and 2Q03
Disposable Income SA	I(0)	2Q03	I(1)	3Q03, 2Q03 and 3Q01	I(1)	2Q02, 1Q03, 4Q00 and 4Q03

Note: I(1) means that the time series is integrated of first order while I(0) indicates that it is stationary at the standard 5% significance level. SA stands for seasonally adjusted.

C. Cointegration analysis

Based on the results of the previous section, which indicated that the time series have a unit root, tests for cointegration were applied. The goal is to find a linear combination of the series within a model, say $a' y_t$, where a is not null, that is stationary. Traditional tests, such as Engle and Granger (1987) and Johansen (1988), do not consider the existence of structural breaks, which could critically affect the cointegration results. To overcome this flaw, we applied the procedures proposed by Saikkonen and Lutkepohl (2000) and Johansen, Mosconi and Nielsen (2000). The time series are filtered for the structural breaks in the first stage, and then cointegration is tested in the second one. The tests were applied using only a constant as deterministic term and the same structural breaks identified by the unit root tests. Table 3 summarizes the results.

Johansen's trace test indicates that there is one cointegrating vector for Argentina, Brazil, and Chile. For Mexico, however, it suggests the existence of two cointegrating vectors. The Saikkonen and Lutkepohl test shows evidence of just one vector. Thus, based on this evidence that the variables under consideration are cointegrated, we estimated and tested restrictions imposed by RE on equation (16), which is an error correction model. The next section presents the estimation and analysis.

It is useful to remember that the restrictions imposed by RE simultaneously require that the survival probability be equal to one $(\gamma = 1)$ and the percentage of individuals facing liquidity constraints be equal to zero $(\theta = 0)$. Given the systems of equations (17) to (22) and (10) to (15) for the model's parameters, the estimation of equation (16) allows to identify the structural parameters of interest and jointly test the restriction imposed by RE.

D. Model estimation

Theoretically, it is expected that the subjective discount factor (δ) has a value between 0 and 1. The survival probability (γ)should also belong to the interval [0, 1], with 1 corresponding to the case where individuals act as if they lived forever (infinite horizon). For the substitutability between private and public consumption (σ), a negative value suggests they are complementary while a positive value indicates substitutability between those consumptions. In case σ is equal to 0, one can conclude that public consumption does not affect private consumption.

Equation (16) is estimated by FIML and system GMM (S-GMM). These two methods are considered more robust because they are not subject to the endogeneity problem. The OLS estimator will lead to highly inconsistent estimates because of the stochastic nature of some explanatory variables.[3] The estimation by S-GMM is included in the analysis because it does not impose the restrictive assumption of normally distributed disturbances, as the FIML does. We avoided the problem of weak instruments and moment conditions overidentification by using all lagged variables from the system as instruments in the S-GMM estimator. In order to model structural breaks, level dummy variables were included, corresponding to relevant macroeconomic events in each country. For Chile, no such event was observed during the period and the time of the break was endogenously selected by the Perron (1997) unit root test from Table 2. We also present the joint statistic for the Jarque Bera normality test, based on the inverse square root of the residual correlation matrix.

Table 3. Cointegration tests

Country / null hypothesis	Saikkonen and Lutkepohl (2000)	Johansen (2000) Trace Test	Structural Break
Argentina			
Rank = 0	53,68*	75,43*	3Q02
Rank = 1	12,97*	19,20	3Q02
Rank = 2	1,61	5,54	3Q02
Brazil			
Rank = 0	37,71*	54,43*	2Q02
Rank = 1	6,79	16,63	2Q02
Rank = 2	0,10	5,77	2Q02
Chile			
Rank = 0	24,35*	44,68*	2Q02
Rank = 1	5,48	21,78	2Q02
Rank = 2	2,91	7,07	2Q02
Mexico			
Rank = 0	29,63*	58,48*	3Q02
Rank = 1	11,23	26,58*	3Q02
Rank = 2	1,46	6,86	3Q02

Note:* Indicates that the null hypothesis of rank = q, with q = 0, 1 or 2, is rejected at the 10% significance level.

[3] The OLS estimation was also performed and the results are available upon request.

Table 4 presents the results of the FIML estimation. Initial values of the parameters were set according to what is theoretically expected and to achieve convergence of the solution. Thus, the initial values were given by $\gamma = 0.99$, $\delta = 1$, $\sigma = 1$, $\theta = 0$, $\rho_1 = 1.1$ and $\rho_2 = 1.1$.

Table 4. FIML estimation

	Argentina	Brazil	Chile	Mexico
γ	0,980**	0,955**	0.992**	0.958**
	(0,138)	(0,035)	(0.010)	(0.056)
δ	1,122**	1,256**	1.006**	1.526**
	(0,213)	(0,209)	(0.049)	(0.327)
θ	0,999	-0,499	-0.226	-1.034
	(1,034)	(0,717)	(0.588)	(2.412)
ρ_1	1,030**	0,454**	0.104	0.402**
	(0,099)	(0,187)	(0.164)	(0.145)
ρ_2	0,572**	0,085	0.697**	-0.239
	(0,185)	(0,233)	(0.111)	(0.173)
σ	2,317	-0,938	-0.879	-0.962**
	(2,182)	(0,760)	(3.288)	(0.443)
dummy	-4,351**	-1,705	33.391	-49.768
	(1,870)	(2,207)	(50.455)	(45.489)
dummy date	2Q01	2Q02	2Q02	3Q02
Log Likelihood	-349,899	-406,914	-796.567	-769.132
Jarque Bera	47,149*	32,162*	81.531*	53.043*

Note: * and ** indicate rejection of the null that the coefficient is zero with 10 and 5% of significance, respectively. In parentheses are the standard deviations.

The FIML produced estimated coefficients for the survival probability, γ, and the subjective discount factor, δ, according to what was expected. The coefficient on the percentage of individuals with liquidity constraints was not found to be statistically significant. Regarding the substitutability between private and public consumption, only Mexico presented a significant and negative estimated coefficient (statistically equal to -1). This means that there is a perfect complementary relationship between public and private consumption in that country.

Khalid (1996) found similar values for γ and δ using annual data in the period from 1960 to 1990 for Argentina, Brazil, and Mexico. For θ , however, the two studies found considerably different estimations. That might be because Latin-American countries were under severe credit restrictions during the heterogeneous period used in Khalid's empirical evidence. In the recent period, considered in this study, a relative economic stability was achieved in the region, which might have reduced credit constraints faced by consumers within each one of those countries.

Finally, the system of equations was estimated by S-GMM. The vector of initial values for the parameters was the same one used in the FIML estimation. The instrument set included the lagged variables of the model, yielding an exactly indentified system. However, as the set of instruments is the same for each one of the three equations of the system, the estimations are still not immune to the weak instruments problem. The S-GMM estimator is still relevant given the absence of any restrictive assumptions on the distribution of residuals, as the previous FIML estimator. The results are reported in Table 5.

Table 5. S-GMM estimation

	Argentina	Brazil	Chile	Mexico
γ	1.064**	0,9417**	0.988**	1.004**
	(0.036)	(0,032)	(0.005)	(0.002)
δ	1.151**	0,941**	0.998**	2.937**
	(0.052)	(0,018)	(0.013)	(0.542)
θ	-0.475**	0,051	-0.030	0.713**
	(0.127)	(0,083)	(0.06)	(0.061)
ρ_1	0.585**	0,638**	0.209	0.424**
	(0.068)	(0,059)	(0.151)	(0.094)
ρ_2	0.744**	0,296**	0.858**	-0.295**
	(0.081)	(0,087)	(0.059)	(0.087)
	6.052**	0,800	1.062*	-1.645**
	(2.190)	(0,678)	(0.601)	(0.213)
dummy	-0.541**	0,867	50.224**	-48.597**
	(0.122)	(1,368)	(19.547)	(24.537)
dummy date	2Q01	2Q02	2Q02	3Q02
Hansen	n.a.	n.a.	n.a.	n.a.
Jarque Bera	26.535	21,661*	49.725*	44.992*

Note: * and ** indicate rejection of the null that the coefficient is zero with 10 and 5% of significance, respectively. In parentheses are the standard deviations. n.a. means that the test does not apply because the system is exactly identified.

In general, the estimated coefficients are close to the ones obtained by the FIML estimation. The only exceptions are γ statistically greater than 1 for Mexico and δ negative for Argentina. For Mexico, this finding corroborates the conclusion on the complementary relationship between public and private consumption. The theoretical restrictions imposed by RE are tested in the next section.

E. Testing RE restrictions

As discussed in Section II, the theoretical model generates testable restrictions for RE from the estimated parameters. If RE holds, then the estimated survival probability is statistically equal to 1 and the fraction of individuals facing liquidity constraints is statistically equal to 0. In terms of the parameters, this restriction implies that $\gamma = 1$ and $\theta = 0$ simultaneously. The joint hypothesis was tested by the Wald test applied to the models estimated by FIML and GMM. The results are reported in Table 6.

Table 6. Tests for RE restrictions

	GMM	FIML
	H0: $\gamma=1$ and $\theta=0$	H0: $\gamma=1$ and $\theta=0$
Argentina	4,298	1,757
Brazil	2,864	2,363
Chile	3,868	1,065
Mexico	1547,121**	7,966**

Note: ** indicates rejection of the null at the 5% significance level.

The Wald test rejects the null hypothesis of infinite horizon and no-liquidity constraint for Mexico in both estimated models. Thus, there is strong evidence that RE does not hold for Mexico in the recent period. On the other hand, for Argentina, Brazil, and Chile, the Wald test indicated the opposite conclusion. The null hypothesis that validates RE is not rejected by both the S-GMM and FIML estimations. Thus, there is strong evidence that RE holds for Argentina, Brazil, and Chile but does not hold for Mexico during the recent period.[4]

[4] The data set covered a period of relative economic stability for Brazil, Chile, and Mexico. On the other hand, Argentina fell into default and a deep economic crisis in the 2001-2002 period. However, RE failed only for Mexico. The recent international financial crisis, which started in the U.S. in early 2008, did not enter the estimation to avoid the impact of international liquidity constraints on the results.

The result for Mexico confirms the findings by Khalid (1996), who used annual data from 1960 to 1988 to test RE for 21 developing countries. For Argentina and Brazil, however, the results are in sharp contrast. Argentina was excluded from Khalid's analysis due to rejection of the overidentification restrictions in the first stage of his estimation. The RE hypothesis was rejected for Brazil at the 5% significance level. These differences might be due to the various structural breaks that characterize Khalid's period, such as hyperinflation episodes, debt crisis, first and second oil price shocks and successive launches and failures of economic stabilization plans. These breaks were not accounted for in his estimations and might have severely biased the results. On the other hand, the quarterly period from 1996 to 2007 was marked by a relative economic stability in the region, after the successful launch of economic stabilization programs. In addition, structural breaks that showed up in the time series were appropriately modeled in the estimation procedure. Thus, the differences between our results and Khalid's might be credited to the time period and the approach adopted to deal with structural breaks in the time series.

IV. Concluding remarks

This paper provided empirical evidence on the validity of Ricardian Equivalence (RE) for Argentina, Brazil, Chile, and Mexico in the recent period of relative economic stability in Latin America. These countries were chosen for their economic and political representativeness in the region. The theoretical model, proposed by Khalid (1996), provides testable restrictions implied by RE. Alternative estimation procedures, represented by FIML and S-GMM, were applied and restrictions were tested by the Wald test.

The results showed that RE cannot be rejected for Argentina, Brazil and Chile, while it is strongly rejected for Mexico. Estimated parameters indicated that the survival probability, γ, and the fraction of individuals with no liquidity constraints, θ, are statistically equal to 1 and 0, respectively, in Argentina, Brazil, and Chile. In Mexico, however, about 70% of the individuals are affected by liquidity constraints.

Note that the previous findings do not confirm Khalid's (1996) results for Argentina and Brazil. The reasons rest on specific period characteristics and the approach applied to identify and model structural breaks in the time series. Khalid (1996) analyzed the heterogeneous period from 1960 to 1988 without any attempt to model structural breaks which affected the developing countries from his sample. This omission might have biased his results. Here, on the other hand, we considered a relatively more stable period from the first quarter of 1996 to the fourth quarter of 2007 and appropriately accounted for structural breaks in the time series of Argentina, Brazil, Chile, and Mexico.

The favorable evidence of RE in the Brazilian case is in line with recent studies on the issue of fiscal versus monetary dominance. Fialho and Portugal (2005) and Gadelha and Divino (2008), for instance, conclude that the Brazilian economy was under monetary dominance in the post 1994 period. Thus, there was an active monetary policy in the country seeking price stabilization which was backed by a Ricardian fiscal policy.

In the context of designing fiscal stabilization programs, empirical evidences of RE are highly relevant for policy making. The fiscal authority is tempted to adopt expansionary policies following a Keynesian orientation. Usually, the measures involve tax reductions and increases in government expenditure. These are, for instance, the guidelines followed by countries seeking to use fiscal policy to stimulate domestic economic activity and increase the level of employment. If RE holds, however, an expansionary fiscal policy might have limited impact on the consumption path and, as a result, on the real side of the economy. According to the previous results, this is the case of Argentina, Brazil, and Chile, but not Mexico.

References

Barro, Robert. J. (1974), Are government bonds net wealth?, *Journal of Political Economy* **82**: 1095-1117.

Bassetto, Marco, and Narayana Kocherlakota (2003), On the irrelevance of government when taxes are distortionary, *Journal of Monetary Economics* **51**: 299-304.

Blanchard, Olivier (1985), Debt deficits and finite horizons, *Journal of Political Economy* **93**: 223-47.

Buchanan, James M. (1976), Barro on the Ricardian Equivalence Theorem, *Journal of Political Economy* **84**: 373-81.

Cuaresma, Jesus C., and Gerhard Reithschuller (2006), Is the Ricardian Equivalence Proposition an aerie faerie theory for Europe?, *Economica* **74**: 682-94.

Dickey, David A. and Wayne A. Fuller (1979), Distribution of the estimators for autoregressive time series with unit root, *Journal of the American Statistical Association* **74**: 427-31.

Dickey, David A., and Wayne A. Fuller (1981), Likelihood ratio statistics for autoregressive time series with a unit root, *Econometrica* **49**: 1057-72.

Divino, Jose A., and Jaime Orrillo (2008), Ricardian Equivalence Theorem in "a few lines", LACEA-LAMES 2008 Parallel Meetings, Rio de Janeiro.

Elmendorf, Douglas and Nicholas G. Mankiw (1998), Government debt, in J. B. Taylor and M. Woodford, eds., *Handbook of Macroeconomics*. Amsterdam, North Holland.

Elliott, Graham, Thomas J. Rothenberg, and James H. Stock (1996), Efficient tests for an autoregressive unit root, *Econometrica* **64**: 813-36.

Engle, Robert, and Clive Granger (1987), Co-integration and error correction: representation, estimation and testing, *Econometrica* **55**: 251-276.

Fialho, Marcelo L., and Marcelo S. Portugal (2005), Monetary and fiscal policy interactions in Brazil: an application to the fiscal theory of the price level, *Estudos Econômicos* 35: 657-85.

Gadelha, Sergio, and Jose A. Divino (2008), Dominância fiscal ou dominância monetária no Brasil? Uma análise de causalidade, *Economia Aplicada* **12**: 659-75.

Hansen, Lars P. (1982), Generalized instrumental variables estimation of nonlinear expectations models, *Econometrica* **50**: 1269-86.

Hayford, Marc (1989), Liquidity constraints and the Ricardian Equivalence Theorem, *Journal of Money, Credit and Banking* **21**: 380-387.

Johansen, Soren (1988), Statistical analysis of cointegration vectors, *Journal of Economic Dynamics and Control* **12**: 231-54.

Johansen, Soren, Rocco Mosconi, and Bent Nielsen (2000), Cointegration analysis in the presence of structural breaks in the deterministic trend, *Econometrics Journal* **3**: 216-49.

Khalid, Ahmed M. (1996), Ricardian Equivalence: empirical evidence from developing countries, *Journal of Development Economics* **51**: 413-32.

Kwiatkowski, Denis, Peter C. B. Phillips, Peter Schmidt, and Yongcheol Shin (1992), Testing the null hypothesis of stationary against the alternative of a unit root, *Journal of Econometrics* **54**: 159-178.

Lee, Junsoo, and Mark C. Strazicich (2003), Minimum Lagrange multiplier unit root test with two structural breaks, *Review of Economics and Statistics* **85**: 1082-1089.

Leiderman, Leonardo, and Assaf Razin (1988), Testing Ricardian neutrality with an intertemporal stochastic model, *Journal of Money, Credit and Banking* **20**: 1-21.

Modigliani, Franco, and Merton Miller (1958), The cost of capital, corporation finance and the theory of investment, *American Economic Review* **48**: 261-97.

Ng, Serena, and Pierre Perron (2001), Lag length selection and the construction of unit root test with good size and power, *Econometrica* **69**: 1519-54.

Perron, Pierre (1989), The great crash, the oil price shock, and the unit root hypothesis, *Econometrica* **57**: 1361-401.

Perron, Pierre (1997), Further evidence on breaking trend functions in macroeconomic variables, *Journal of Econometrics* **80**: 355-85.

Perron, Pierre, and Serena Ng (1996), Useful modifications to unit root tests with dependent errors and their local asymptotic properties, *Review of Economic Studies* **63**: 435-65.

Phillips, Peter C. B., and Pierre Perron (1988), Testing for a unit root in time series regression, *Biometrika* **75**: 335-46.

Ricciuti, Roberto (2003), Assessing Ricardian Equivalence, *Journal of Economic Surveys* **17**: 55-78.

Saikkonen, Pentti, and Helmut Lutkepohl (2000), Testing for the cointegrating rank of a VAR process with an intercept, *Econometric Theory* **16**: 451-64.