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Ricardian Equivalence, Liquidity Constraints, and the Yaari-Blanchard Effect: Tests For Developing Countries

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Abstract

Adjustment programs in developing countries have emphasized the importance of reducing fiscal deficits in order to improve private sector saving and investment performance. Recent theoretical analyses associated with the Ricardian equivalence proposition, however, suggest that, in the limit, changes in the level of public sector savings may be completely offset by a change in private savings. This offset would occur because changes in the level of government savings imply changes in the level of future taxation, which in turn affects current private sector saving. Empirical tests of the model for a sample of developing economies do not support the equivalence proposition owing to the prevalence of liquidity constraints.

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Summary

Fiscal policy plays a central role in programs for short-run stabilization and medium-term adjustment in developing countries. Large and increasing public sector deficits generally lie behind excessive expansion of aggregate demand and hence widening current account deficits and rising inflation. Moreover, fiscal deficits are commonly perceived to absorb domestic saving and displace private investment, thereby inhibiting medium-term growth and adjustment. Stabilization programs, therefore, typically envisage a reduction in fiscal deficits, coupled with a change in the composition of public sector spending from consumption to investment.

Recent theoretical developments in macroeconomics, however, suggest that the above analysis of the effects of fiscal policy on savings and investment may not hold under some circumstances. For example, the "Ricardian Equivalence" proposition suggests that, among other things, public sector saving decisions may influence the saving behavior of the private sector, since changes in public sector savings represent signals for tax policy in the future. Individuals, in order to smooth their consumption path over their lifetimes, would change their saving behavior to take into account any future taxation that is suggested by the current changes in public sector debt or deficits. If the equivalence proposition holds, total domestic saving may remain unresponsive to changes in public sector saving.

In view of the important implications of the Ricardian Equivalence proposition for the role of fiscal policy in the stabilization process, this paper tests its empirical relevance in developing economies. The approach used tests whether the conditions required for the proposition hold empirically. Specifically, equivalence would not be obtained if the so-called Yaari-Blanchard effect prevailed--i.e., the discount rates of the private and the public sector were significantly different--or if a significant proportion of the population were liquidity constrained, or if the tax system was distortionary. The empirical estimates derived in the paper suggest that full Ricardian Equivalence can be rejected in 15 of the 16 countries in the sample. The model that was developed distinguished between the sources of the deviation from equivalence. In keeping with earlier studies, no evidence was found in support of the Yaari-Blanchard effect. Instead the data supported the conclusion that prevalence of liquidity constraints in a large number of countries in the sample, is the principal reason for the rejection of Ricardian Equivalence.

I. Introduction

Fiscal policy adjustments have come to occupy center stage in programs for short-run stabilization and medium-term adjustment in developing countries. Increases in public sector deficits have been cited as the cause of excessive expansion of aggregate demand, leading to current account deficits and inflation. At the same time, such deficits have been perceived as absorbing domestic saving and displacing private investment, thereby inhibiting medium-term growth and adjustment. Programs for stabilization and adjustment thus typically envisage a reduction in fiscal deficits, coupled with a change in the composition of exhaustive public-sector spending from consumption to investment.

The increased emphasis on the central role of fiscal adjustment in developing countries has, however, coincided with theoretical developments in macroeconomics which render the analysis of the effects of fiscal policy on the economy somewhat problematic. Specifically, the "Ricardian equivalence" proposition, recently resurrected by Barro (1974) can be shown to have the following unorthodox macroeconomic implications:

a. A reduction in the fiscal deficit brought about by an increase in taxes on the private sector will be exactly offset by a reduction in private saving, so total domestic saving cannot be increased when fiscal deficits are reduced by these means.

b. If the propensity to consume out of permanent income is unity, an increase in public sector consumption expenditures cannot be responsible either for excessive expansion of aggregate demand or for crowding out of private investment.

c. A change in the mix of exhaustive government spending from consumption to investment will reduce private saving and thus diminish the pool of domestic saving available to finance private investment.

d. The effect on aggregate demand of an increase in government spending will depend on whether the increased spending is devoted to consumption or investment, but not on whether it is accompanied by a larger fiscal deficit.

e. A temporary increase in public consumption will have a larger impact on aggregate demand than a permanent increase of the same magnitude.

Although the list is not exhaustive, it clearly demonstrates that, if empirically relevant in developing countries, Ricardian equivalence would have far-reaching and profound implications for the analysis of the effects of fiscal adjustment in such countries.

The conditions required for Ricardian equivalence to hold, however, are quite stringent. They include:

a. Private consumption must conform to the permanent income hypothesis.

b. Households' planning horizons must effectively be infinite.

c. The rate of discount applied to future income by consumers must be equal to the rate at which the public sector can borrow.

d. In forming expectations of future tax liabilities, consumers must act rationally -- specifically, the implications of the government budget identity must be incorporated into their expectations of future taxes.

Empirical testing of the Ricardian equivalence proposition has taken the form of either testing its implications -- e.g. for private consumption behavior -- or of testing the empirical validity of some subset of the conditions required for the proposition to hold. Most of the existing work takes the former approach, and results have not on the whole been favorable to the Ricardian proposition in the case of the United States (see the survey by Bernheim (1987)). Although this approach to empirical verification may be intrinsically flawed for Lucas-critique reasons (see Blejer and Leiderman (1987)), the proposition has not fared much better in direct tests of conditions (a)-(d), at least for the United States. The importance of liquidity constraints, undermining (a), has been documented, for example, by Flavin (1985) and by Hubbard and Judd (1986), while Hayashi (1982) has provided evidence that discount rates for anticipated labor income may exceed public borrowing costs, in contradiction to (c).

A theoretical rationale for the latter result has been provided by Yaari (1965) and Blanchard (1985). Even if households' planning horizons are effectively infinite (e.g., through a system of operative intergenerational transfers), the effective discount rate applied to future labor income will exceed the riskless rate of return if each dynasty's probability of surviving to the next period is not unity. This Yaari-Blanchard effect would cause Ricardian equivalence to fail, since a deferral of taxes to the future would increase the present value of the household's resources over its planning horizon.

Little empirical work on Ricardian equivalence exists for developing cuntries. Haque (1986) tested the Yaari-Blanchard effect for a sample consisting of annual data for 26 developing countries. Leiderman and Razin (1987) tested both liquidity constraints and the Yaari-Blanchard effect using monthly data for Israel. Surprisingly, in view of the results for the United States cited above and of the presumed severity of capital market imperfections in developing countries, these studies derived results consistent with Ricardian equivalence.

Unfortunately, both studies are subject to limitations which leave the empirical relevance of Ricardian equivalence for developing countries in doubt. Although Haque used a fairly large and diverse sample of developing countries, he tested for the presence of a Yaari-Blanchard effect under the maintained hypothesis that liquidity constraints are absent. Since one would expect ex ante that such constraints would be the most important reason for Ricardian equivalence to fail in developing countries, the robustness of his results to the abandonment of this maintained hypothesis must be examined. In the case of the Leiderman-Razin study, the critical issue is the extent to which the results for Israel can be generalized to other developing countries. We extend Haque's model to take account of the possible existence of liquidity constraints and apply this more general model to a sample of 16 developing countries. Our expanded model is able to separately test for the presence of liquidity constraints and of a Yaari-Blanchard effect, as in Leiderman and Razin. The two tests must be conducted simultaneously, since our model demonstrates that ignoring liquidity constraints may tend to bias the results using Haque's original model toward accepting the null hypothesis that the Yaari-Blanchard effect is absent. On the other hand, our model and empirical methodology will differ from that of Leiderman and Razin, and our sample is, of course, much broader.

The remainder of the paper is organized in three sections. The model is presented in the next section, followed by the empirical results in Section III. A summary of our conclusions is presented in the final section.

II. The Model

We begin by deriving an expression for aggregate per capita consumption for households which do not face liquidity constraints, but which may be subject to a Yaari-Blanchard effect. 1/

Household wealth is defined as the sum of human wealth (H_t^u) and non-human wealth. 2/ The latter consists of one-period bonds purchased last period (B_{t-1}^u), which are assumed to pay a fixed

1/ The final expression is given by equation (9) below. For the sake of brevity, in this section we present only the aggregate per capita version of the consumption model we have adopted for unconstrained households. For a detailed derivation of these relationships from the underlying equations for representative members of individual age cohorts, following closely along the lines of Frenkel and Razin (1987), see Haque (1986).

2/ The superscript u will denote variables pertaining to households that are not liquidity-constrained. Those corresponding to constrained households are given the superscript c.

interest rate r. 1/ Letting $R = 1+r$, interest plus amortization on one-period bonds held over from last period is RB^u_{t-1}. Thus aggregate per capita household wealth for unconstrained households is given by:

$$W^u_t = H^u_t + RB^u_{t-1} \; . \tag{1}$$

Human wealth is the present value of expected future labor income net of taxes. Following Blanchard (1985), the household is assumed to face a fixed probability γ of surviving to the next period, regardless of its age. Let Y^u_{t+j} denote labor income net of taxes in period t+j. Human wealth is therefore given by:

$$H^u_t = \sum_{j=0}^{\infty} (\gamma/R)^j \; E_t Y^u_{t+j} \; , \tag{2}$$

where E_t is an expectations operator conditional on information available at time t. 2/

The evolution of nonhuman wealth over time will depend on the path followed by consumption, denoted C^u_t. This dependence is governed by the budget constraint:

$$B^u_t = RB^u_{t-1} + Y^u_t - C^u_t \; . \tag{3}$$

This equation states that households can alter their claims to future nonlabor income by choosing to consume more or less than current disposable income. Turning to human wealth, since claims to this type of wealth cannot be traded, households cannot alter their claims to future labor income by saving or dissaving today. The dynamics of human wealth arise instead from changing expectations of future disposable labor income. As defined in (2), human wealth depends on today's expectation of future events, so the arrival of new information will cause households to revise their calculations of human wealth. Let $\tilde{H}^u_t$ signify the change in the expected value of time-t human wealth

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1/ Notice that our model does not allow for other durable assets. Most importantly, there are no consumer durables.

2/ Equation (2) and the tests based on it apply whenever there is a wedge between the discount rate for human capital and that which would apply to a riskless asset. The probability of dynastic extinction is only one justification for such a wedge.

from time t-1 to time t, conditional on the household's survival to time t. Then $\tilde{H}^u_t$ is defined as:

$$\tilde{H}^u_t = \sum_{j=0}^{\infty} (\gamma/R)^j \,[E_t(Y^u_{t+j}) - E_{t-1}(Y^u_{t+j})] \tag{4}$$

$$= \sum_{j=0}^{\infty} (\gamma/R)^j \,\tilde{Y}^u_{t+j} \;,$$

where $\tilde{Y}^u_{t+j} = E_t(Y^u_{t+j}) - E_{t-1}(Y^u_{t+j})$. Using (4), the evolution of human wealth over time can be expressed as:

$$H^u_t = E_{t-1}(H^u_t) + \tilde{H}^u_t \tag{5}$$

$$= (R/\gamma)\,(H^u_{t-1} - Y^u_{t-1}) + \tilde{H}^u_t$$

Equation (5) decomposes H^u_t into a portion which was anticipated at time t-1 and the revision of expectations between time t-1 and time t. The expectation in the first term on the right-hand side of the first equality in (5) is conditional on household survival from t-1 to t. To calculate $E_{t-1}H^u_t$ using H^u_{t-1}, it is necessary to exclude period t-1 disposable labor income (which will of course not be available at time t), multiply by $1/\gamma$ to convert the unconditional expectation $(H^u_{t-1} - Y^u_{t-1})$ into a conditional expectation, and convert period (t-1) values into period t values by applying the factor R.

Finally, household consumption is assumed, in conventional permanent-income fashion (see Flavin (1985)), to consist of a systematic portion which is proportional to household wealth and a white noise term U_t:

$$C^u_t = (1-s)\,W^u_t + U_t, \tag{6}$$

where the factor of proportionality is expressed for convenience as (1-s) and $0 < s < 1$.

In order to permit the parameters of this model to be estimated, we need to express household consumption in terms of observable variables. To do so, notice that using (1), the budget constraint (3) can be written as:

$$B_t^u = W_t^u - C_t^u - (H_t^u - Y_t^u) .$$

Moving (5) forward one period and substituting into the above expression yields:

$$B_t^u = W_t^u - C_t^u - (\gamma/R)\, H_{t+1}^u + (\gamma/R)\, \tilde{H}_{t+1}^u . \tag{7}$$

Substituting (1) and (7) into the consumption function (6) and simplifying, we have:

$$C_t^u = sRC_{t-1}^u + (1-s)\,(1-\gamma)\, H_t^u + \gamma(1-s)\, \tilde{H}_t^u + U_t - RU_{t-1} \tag{8}$$

Equation (8) is a generalization of Hall's (1978) Euler-equation approach to testing the permanent-income hypothesis. According to Hall, no information available at time t-1 should help predict time t consumption once lagged consumption is included in the regression, since all such information would already have been captured in the previous period's consumption decision. If $\gamma = 1$, the variable H_t^u drops out of (8), which reduces to Hall's formulation except for the moving average error term. As also pointed out by Flavin (1981) and Hayashi (1982), this term arises from allowing for transitory consumption in the consumption function (6).

That the existence of a Yaari-Blanchard effect should cause H_t^u to appear in equation (8) can readily be understood intuitively. Notice first that H_t^u will appear in (8) only if the anticipated component of human wealth affects current consumption even after last period's consumption is taken into account, since the unanticipated component can be subsumed into the third term on the right-hand side of (8), using the first equality in (5). Doing this produces:

$$C_t^u = sRC_{t-1}^u + (1-s)\,(1-\gamma)\, E_{t-1}(H_t^u) + (1-s)\, \tilde{H}_t^u + U_t - RU_{t-1} . \tag{8'}$$

This equation can also be written:

$$C_t^u = sRC_{t-1}^u + (1-s)\, [(1-\gamma)E_{t-1}(H_t^u) + \tilde{H}_t^u] + U_t - RU_{t-1} . \tag{8''}$$

The term in square brackets is the change in the expected value of time t human wealth from time t-1 to time t. It has two components, of which $\tilde{H}_t^u$ captures revisions in expectations of future income flows conditional on survival to time t. The first term, however, is the relevant one for our purposes. Recall that $E_{t-1}(H_t^u)$ is the expectation of H_t^u conditioned on information available at time t-1 and on survival to time t. The expectation of H_t^u not conditioned on survival is $\gamma E_{t-1}(H_t^u)$, since this takes into account the uncertainty of surviving to the next period. The first term, therefore, is the difference between the conditional and unconditional expectations -- i.e., it is the gain in the expected value of time t human wealth due to the certainty of survival at time t. Since this gain can only exist if survival was uncertain at time t-1, the term drops out and the Hall formulation is recovered if $\gamma=1$.

Equation (8) cannot be estimated directly, because human wealth is not an observable variable. To eliminate the unobservable H_t^u from equation (8), multiply the lagged value of (8) by R/γ, subtract the result from (8), and use (5):

$$C_t^u = R(s+\frac{1}{\gamma})\,C_{t-1}^u - \frac{sR^2}{\gamma}\,C_{t-2}^u - \frac{R(1-s)\,(1-\gamma)}{\gamma}\,Y_{t-1}^u + (1-s)\,\tilde{H}_t^u \tag{9}$$

$$- R(1-s)\,\tilde{H}_{t-1}^u + U_t - R(1-\frac{1}{\gamma})\,U_{t-1} - \frac{R^2}{\gamma}\,U_{t-2}\,.$$

This is the equation estimated by Haque (1986), treating $\tilde{H}_t^u$ and $\tilde{H}_{t-1}^u$ as unobservable random shocks.

Turning to liquidity-contrained households, these are assumed to consume their disposable labor income each period, so that the aggregate per capita consumption function for such households takes the form:

$$C_t^c = Y_t^c\,, \tag{10}$$

with superscript c denoting constrained households. As is conventional in the literature (see Hayashi (1982), Leiderman and Razin (1987)) we assume that total aggregate per capita consumption is a weighted average of consumption by liquidity-constrained and unconstrained households, with the weight of the latter denoted θ. We assume also that $Y_t^u = Y_t^c = Y_t$. Letting C_t denote total per capita consumption in period t, we have:

$$C_t = \theta C_t^u + (1-\theta)\,C_t^c. \tag{11}$$

Using (9) and (10), this becomes:

$$C_t = \frac{\theta R}{\gamma}(1 + s\gamma)C^u_{t-1} - \theta\frac{SR^2}{\gamma}C^u_{t-2} + (1-\theta)\ Y_t - \theta\ \frac{R(1-s)(1-\gamma)}{\gamma}\ Y_{t-1} + \varepsilon_t \tag{12}$$

where:

$$\varepsilon_t = \theta\ \{(1-s)\ \tilde{H}^u_t + U_t - R(1-s)\tilde{H}^u_{t-1} - R(1-\frac{1}{\gamma})U_{t-1} - \frac{R^2}{\gamma}U_{t-2}\}$$

Equation (12) cannot be estimated directly, since the per capita consumption of unconstrained households, C^u_t, is unobservable when $\theta \neq 1$. To write this equation in terms of observable magnitudes, use (10) and (11) to eliminate C^u_t from (12). The result is:

$$C_t = \alpha_1 C_{t-1} + \alpha_2 C_{t-2} + \alpha_3 Y_t + \alpha_4 Y_{t-1} + \alpha_5 Y_{t-2} + \varepsilon_t, \tag{13}$$

where

$$\alpha_1 = \frac{R}{\gamma}(1 + s\gamma) > 1$$

$$\alpha_2 = -\frac{sR^2}{\gamma} < 0$$

$$\alpha_3 = 1-\theta,\ 0 \leq \alpha_3 \leq 1$$

$$\alpha_4 = \frac{-R}{\gamma}(1 + \gamma s - \theta\ [\gamma + s]) < 0$$

$$\alpha_5 = (1 - \theta)\ \frac{SR^2}{\gamma};\ 0 \leq \alpha_5 \leq -\alpha_2$$

Equation (13) can be estimated consistently using instrumental variables. If we follow Hall (1978), Flavin (1981) and Hayashi (1982) in relying on the law of large numbers to render individual variations in transitory consumption negligible in the aggregate, then the terms in U_{t-j} drop out of the disturbance ε_t. In this case, any variables dated t-2 and before which help predict consumption and disposable labor income over periods t-1 and t would serve as valid instruments.

Equation (13) forms the basis for our tests. Full Ricardian equivalence requires both the absence of liquidity constraints (θ=1) and of a Yaari-Blanchard effect (γ=1). It is easily verified that if $\theta = \gamma = 1$, then $\alpha_3 = \alpha_4 = \alpha_5 = 0$. Thus, a test for Ricardian equivalence can be conducted by examining the validity of the exclusion restrictions $\alpha_3 = \alpha_4 = \alpha_5 = 0$. If these restrictions are rejected, then it remains to establish whether rejection is caused by the presence of liquidity constraints and/or of the Yaari-Blanchard effect. In the absence of liquidity constraints, $\theta = 1$. It can readily be seen that in this case $\alpha_3 = \alpha_5 = 0$. Thus, the presence of liquidity constraints could be tested by examining whether the exclusion restrictions $\alpha_3 = \alpha_5 = 0$ are rejected by the data. If they are, then full Ricardian equivalence fails at least partly because of the importance of liquidity constraints. To establish whether it <u>also</u> fails due to Yaari-Blanchard effects requires obtaining point estimates of γ and testing $\gamma = 1$. This can be done through nonlinear instrumental variable estimation of (13). This estimation would also produce point estimates of $1 - \theta$, permitting an assessment of the <u>economic</u> significance of liquidity constraints. If the data do <u>not</u> reject $\alpha_3 = \alpha_5 = 0$, then Ricardian equivalence would be rejected due to the Yaari-Blanchard effect, the economic significance of which can then be ascertained by obtaining point estimates of γ as described above.

Before proceeding to estimation, notice the relationship of (13) to the tests for the Yaari-Blanchard effect conducted by Haque (1986) under the maintained hypothesis of no liquidity constraints. Haque estimated (13) with $\alpha_3 = \alpha_5 = 0$. Since real income is a strongly serially correlated variable, and since α_3 and α_5 are both positive, the omission of Y_t and Y_{t-2} will impart a positive bias to the coefficient of Y_{t-1} if liquidity constraints are operative. Since α_4 is negative, the coefficient of Y_{t-1} will be biased toward zero--i.e., toward a finding of no Yaari-Blanchard effect though such an effect is in fact present. Thus Haque's (1986) inability to document the evidence of a Yaari-Blanchard effect in his sample of developing countries must be reevaluated in light of the possible importance of binding liquidity constraints.

III. <u>Empirical Results</u>

This section reports estimates of equation (13) in order to conduct the empirical tests described above for a sample of developing countries. The data were drawn from the World Bank's economic and social database and cover the period 1960 to 1985. The sample consisted of sixteen countries: Algeria, Egypt, India, Indonesia, Ivory Coast, Jamaica, Kenya, Korea, Malaysia, Morocco, Nigeria, Peru, Philippines, Portugal, Thailand and Turkey. Apart from data considerations such as the availability of a reasonable length of series, the choice of the sample was determined by a desire to maintain a geographical balance and to obtain a sample that was representative of the various categories of developing countries. The geographical distribution that was obtained in the sample chosen is as follows: six African countries, six Asian countries, two Western Hemisphere countries, and two European countries. According to the classification used in the World Economic Outlook (1987) of the IMF, the sample includes two low income countries (i.e. per capita income according to World Bank estimates less than $410); five of the fifteen most heavily indebted countries; nine primary producers (i.e. countries whose exports of agricultural and mineral primary products other than fuel accounted for over fifty percent of their exports in 1980); three fuel exporters; three mineral exporters; and an exporter of manufactures.

According to the model, the theoretically-appropriate dependent variable is real private per capita consumption, excluding purchases of durables, but including the imputed services of the stock of consumer durables. However, data on consumer expenditures on durables were not available for the countries in our sample. Therefore, the dependent variable in each case is total real private consumption expenditures. 1/ For the independent variables, a measure of real per-capita labor income net of taxes is required. However, in most cases reliable estimates of labor income and tax revenues for the sample period were not available. We follow other researchers (see, for example, Flavin (1985)) in using total income as a proxy for labor income. Where revenues were available, it was not possible to fully separate them from non-tax revenues, which in some cases were items such as oil revenues. Consequently, the course adopted was to proxy disposable income by per-capita nominal GNP divided by the consumer price index. The proxy is expected to be reasonable since most of the countries in the sample have tax bases that are largely unresponsive to changes in incomes, and labor incomes that are highly correlated with GNP.

As indicated in Section II, the appropriate estimation procedure for equation (13) to take account of the correlation between the right-hand side variables and the error term as well as of the moving-average structure of the latter is generalized instrumental variables. Our set of instruments consisted of the second lags of both consumption and income, as well as of real government consumption, real domestic investment, real output in industrial countries, the world real interest rate, and each country's terms of trade index.

The results of generalized instrumental variable estimation of equation (13) are reported in Table 1. The signs and magnitudes of the coefficients conform quite closely to the expectations of the theory. The point estimate of α_1 exceeds unity in 14 of 16 cases, while that of α_2 is negative as expected in all but one case. Similarly, the estimated value of α_3 is positive and less than one for every country. The coefficient of lagged income, α_4, is negative as expected in every case. The sign of the estimated coefficient was at variance with the theory most frequently in the case of α_5. Even in this

1/ If consumer durables tend to be luxury goods, this may be a much better approximation for our sample of developing countries than it would normally be in industrial-country applications, since the share of durables in consumer expenditures may be quite small. We have no independent confirmation of this, however.

Table 1. Estimates of Equation (13) for Eighteen Developing Countries 1/

Country	α_1	α_2	α_3	α_4	α_5	L_1	L_2	R^2
ALGERIA	1.535 (3.608)	-0.553 (-1.385)	0.328 (1.992)	-0.512 (-1.841)	0.195 (1.266)	13.951	15.569	0.980
EGYPT	0.513 (0.947)	0.057 (0.146)	0.564 (1.812)	-0.104 (-0.138)	-0.193 (-0.382)	10.670	10.207	0.945
INDIA	1.217 (3.472)	-0.252 (-0.726)	0.411 (1.918)	-0.546 (-1.002)	0.162 (0.419)	24.835	30.014	0.944
INDONESIA	1.934 (2.422)	-0.975 (-1.360)	0.647 (2.699)	-1.116 (-1.654)	0.492 (1.213)	15.624	20.133	0.990
IVORY COAST	0.818 (2.414)	-0.045 (-0.175)	0.371 (2.489)	-0.275 (-1.496)	0.051 (0.429)	43.617	44.193	0.959
JAMAICA	1.632 (3.353)	-0.574 (-1.313)	0.306 (1.325)	-0.409 (-0.906)	0.064 (0.229)	16.280	15.786	0.813
KENYA	1.172 (3.384)	-0.433 (-1.299)	0.526 (1.907)	-0.448 (-0.847)	0.091 (0.296)	31.873	31.043	0.635
KOREA	1.628 (3.374)	-0.591 (-1.096)	0.186 (1.223)	-0.175 (-0.542)	-0.036 (-0.171)	27.976	29.422	0.998
MALAYSIA	1.593 (4.758)	-0.560 (-2.045)	0.211 (2.269)	-0.068 (-0.377)	-0.168 (-1.381)	33.508	32.994	0.996
MOROCCO	1.684 (2.671)	-0.728 (-1.240)	0.579 (1.936)	-1.183 (-1.630)	0.641 (1.386)	4.226	2.124	0.965
NIGERIA	1.900 (1.760)	-0.918 (-0.922)	0.435 (2.355)	-0.665 (-1.101)	0.240 (0.573	18.982	19.701	0.897
PERU	1.355 (2.873)	-0.485 (-1.530)	0.286 (2.675)	-0.232 (-1.124)	0.042 (0.394)	26.870	22.882	0.961
PHILIPPINES	1.575 (2.954)	-0.592 (-1.153)	0.104 (0.877)	-0.036 (-0.153)	-0.055 (-0.415)	16.768	16.047	0.995
PORTUGAL	1.035 (2.055)	-0.454 (-1.662)	0.444 (4.198)	-0.450 (-1.820)	0.309 (2.143)	23.188	34.650	0.993
THAILAND	2.217 (2.025)	-1.423 (-1.228)	0.788 (2.333)	-1.685 (-1.577)	1.033 (1.255)	40.304	39.741	0.984
TURKEY	1.403 (3.197)	-0.599 (-1.254)	0.639 (2.102)	-0.982 (-1.390)	0.485 (1.261)	32.674	29.003	0.986

1/ Numbers in parentheses are t-statistics.

instance, however, α_5 exhibits the correct positive sign in 12 of 13 countries and none of the incorrectly-signed estimates differs significantly from zero. Furthermore, as anticipated, α_5 is smaller than α_2 in absolute value in 14 of 16 cases. The results appear poorest for Egypt. In this country three of the five estimated coefficients are outside the predicted range. In no other country does more than one of the estimates fall outside of the predicted range.

Although the magnitudes of the coefficients are therefore in close agreement with expectations, the estimation of individual coefficients is not generally very precise. Except for the coefficients of the first lag of consumption and of concemporaneous income, estimates are not generally different from zero at conventional levels of significance. Since the time series involved are highly serially correlated, one suspects that standard errors may be increased by multicollinearity. Since our tests invclve exclusion restrictions on sets of independent variables, correlation among the elements of such sets will not diminish their effectiveness. We are interested not in the <u>individual</u> contributions of the excluded variables, but rather in their <u>joint</u> contribution. To the extent that multicollinearity involves linear combinations of excluded and included variables, on the other hand, our tests will be less likely to reject the null hypothesis.

Column 6 of Table 1 reports the likelihood ratio statistics for the Ricardian equivalence restrictions $\theta = \Upsilon = 1$. This statistic is distributed $\chi^2(3)$, with critical values 7.81 (5 percent probability of Type I error) and 11.34 (1 percent). The restrictions are rejected at the 1 percent confidence level in 14 of 16 cases, and at the 5 percent level in one additional case. We thus fail to reject full Ricardian equivalence in only one case. As suggested in the last section, such rejections could be brought about due to the presence of liquidity constraints, of Yaari-Blanchard effects, or both. To test for the presence of liquidity constraints, we tested the exclusion restrictions $\alpha_3 = \alpha_5 = 0$ which are implied by $\theta = 1$. The resulting likelihood ratio statistics are reported in column 7 of Table 1. These statistics are distributed χ^2 (2), with 5 percent critical value of 5.99 and 1 percent critical value of 9.21. For each of the 14 rejections of Ricardian equivalence, we are also able to reject the absence of liquidity constraints. Thus, full Ricardian equivalence fails in the vast majority of the developing countries in our sample, and in each case it does so at least in part because some fraction of aggregate consumption is attributable to liquidity-constrained households whose consumption responds more strongly to current income than would be implied by the permanent income hypothesis.

To assess the magnitude of the fraction of households subject to such constraints and to determine whether the Yaari-Blanchard effect characterizes the behavior of unconstrained households, we re-estimated equation (13) using nonlinear instrumental variables to extract estimates of the underlying parameters R, s, γ, and θ. Preliminary estimates yielded values of R less than unity in several cases, implying negative real interest rates. To restrict R to values greater than unity, our procedure involved a search over values of R ranging from 1.01 to 1.10, with increments of 0.01. We chose the value that minimized the weighted sum of squared residuals, and the resulting parameter estimates are reported in Table 2. 1/

Turning first to the prevalence of liquidity constraints, estimates of $(1-\theta)$ are reported in column 5 of Table 2. These range from 0.182 in Korea to 0.713 in Thailand consistent with liquidity-constrained fractions of households amounting to about one fifth in the former and more than two thirds in the latter. Of the 16 countries in the sample our estimates indicate that the fraction of liquidity constrained households in the total exceed 30 percent in 10 cases. By contrast, estimates of the fraction of liquidity-constrained households in the United States cluster around 20 percent (Hayashi (1982), Flavin (1981), and Hubbard and Judd (1986)).

Estimates of γ appear in column 4 of Table 2. Before discussing these, notice that the average length of a household's planning horizon in formulating its consumption decisions is given by $1/(1-\gamma)$. Thus with annual data, the difference between a 20-year horizon and an infinite horizon is the difference between $\gamma = 0.95$ and $\gamma = 1$. Very sharp estimates of γ are therefore needed to discriminate between these rather different cases. Our estimates are not sufficiently precise for any of the countries examined to permit us to discriminate between these alternatives. We therefore cannot reject the infinite-horizon case (i.e., the absence of Yaari-Blanchard effects) for any country. However, we can go much further than this. In all but two cases, the point estimates of γ are at least unity. Furthermore, γ is always statistically different from zero. Thus there is strong evidence that unconstrained households operate with multi-year planning horizons in our sample of developing countries, and our point estimates are not suggestive of finite horizons for such households. In short, our results do not support the presence of a Yaari-Blanchard effect in developing countries. This result is consistent with both Haque (1986) and Leiderman and Razin (1987).

IV. Summary and Conclusions

In view of the emphasis currently placed on fiscal adjustment in developing countries and of the profound implications of Ricardian equivalence for the effects of fiscal policy actions, it is important to assess the empirical relevance of this phenomenon in developing

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1/ The point estimates of γ and Θ did not prove very sensitive to the restrictions on R.

Table 2. Estimates of Underlying Parameters 1/

Country	R	s	γ	1-θ	F (2,21)	R^2
ALGERIA	1.01	0.444 (1.614)	1.018 (21.648)	0.204 (1.063)	328.421	0.969
EGYPT	1.01	0.139 (0.458)	1.014 (27.943)	0.446 (1.493)	131.604	0.926
INDIA	1.01	0.457 (1.097)	1.041 (5.466)	0.344 (1.108)	50.678	0.828
INDONESIA	1.01	0.589 (1.426)	1.099 (10.066)	0.584 (2.903)	820.738	0.987
IVORY COAST	1.01	0.176 (0.517)	1.006 (39.598)	0.368 (2.559)	75.747	0.878
JAMAICA	1.01	0.830 (1.452)	0.997 (30.142)	0.184 (0.536)	9.230	0.468
KENYA	1.01	0.374 (0.685)	1.000 (18.762)	0.546 (0.879)	18.786	0.641
KOREA	1.01	0.596 (2.147)	0.967 (15.551)	0.182 (1.416)	2796.880	0.996
MALAYSIA	1.02	0.550 (4.695)	1.051 (14.573)	0.234 (2.301)	1387.940	0.992
MOROCCO	1.08	0.763 (2.941)	1.084 (9.930)	0.609 (3.067)	200.991	0.950
NIGERIA	1.09	0.638 (2.763)	1.083 (6.705)	0.359 (2.634)	99.669	0.905
PERU	1.07	0.441 (8.438)	1.320 (5.576)	0.246 (2.106)	223.523	0.955
PHILIPPINES	1.02	0.426 (5.279)	1.034 (23.686)	0.209 (3.361)	1786.060	0.994
PORTUGAL	1.03	0.785 (2.033)	1.021 (21.000)	0.423 (2.784)	435.231	0.976
THAILAND	1.01	0.814 (1.870)	1.004 (22.521)	0.713 (10.045)	772.956	0.987
TURKEY	1.01	0.095 (0.249)	1.016 (52.695)	0.468 (3.058)	558.001	0.982

1/ Numbers in parentheses are t-statistics.

countries. Many of the pitfalls of empirical testing for Ricardian equivalence can be avoided by testing the empirical validity of the conditions required for the proposition to hold. Since these conditions are rather stringent, and since there are reasons to expect that they would particularly fail to hold in developing countries, the warranted <u>a priori</u> judgment would seem to be that the applicability of Ricardian equivalence would be limited in this setting, whatever its empirical merits in industrial countries.

It is surprising, therefore, that initial work by Haque (1986) and Leiderman and Razin (1987) conclude precisely the opposite. In this paper we generalize the work of both sets of authors by testing for liquidity constraints and for Yaari-Blanchard (finite-horizon) effects in a relatively large and diverse sample of developing countries. Although the nature of our data limits the precision of our estimates, our general model fits the data rather well, producing feasible coefficient estimates with signs in accordance with theoretical expectations in almost all cases. Full Ricardian equivalence can be rejected for 15 of 16 countries at conventional levels of statistical significance, and point estimates indicate that a much larger proportion of the population behaves as if it were liquidity-constrained in these countries than is typically found in studies for the United States. However, in keeping with the results of previous studies, we have uncovered no evidence that unconstrained households exhibit short time horizons.

Our results therefore question the applicability of full Ricardian equivalence to developing countries. This conclusion, however, is due to the importance of liquidity constraints, not of finite horizons. Although we do not suggest that effects arising from the discounting of future tax obligations can safely be ignored, the evidence does not compel us to reassess the weight currently attached to such considerations for policy purposes in developing countries.

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