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University of Maryland

March 1996

INTERNATIONAL MONETARY FUND

STAFF PAPERS

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Vol. 43 No.1 March 1996

IMF *Staff Papers*
Vol. 43, No. 1 (March 1996)

Saving Behavior in Low- and Middle-Income Developing Countries

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The relationship between real interest rates, saving, and growth is a central issue in development economics. Using macroeconomic data for a cross-section of countries, we estimate a model in which the intertemporal elasticity of substitution varies with the level of wealth. The estimated parameters are used to calculate, in the context of a simple endogenous growth model, the responsiveness of saving to real interest rate changes for countries at differing stages of development. The hypothesis that the saving rate, and its sensitivity to the interest rate, are a rising function of income finds strong empirical support. [JEL E21, F41, F43, O11, O16, O57]

THE IMPACT OF CHANGES in real interest rates on saving, investment, and economic growth is a central issue in development economics. According to one familiar view (see, for example, McKinnon (1973) and Shaw (1973)), an increase in real interest rates in developing countries should encourage saving and expand the supply of credit available to domestic investors, thereby enabling the economy to grow more quickly. Indeed, a number of liberalization programs supported by the international financial institutions over the years have had as their explicit objective to increase interest rates from levels that in many cases were substantially negative in real

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terms. While increases in real interest rates have often been the outcome of such liberalization episodes (see, for example, Galbis (1993)), their impact on domestic saving and investment has been unclear.[1]

There is little consensus in the empirical literature on the interaction between saving and the real rate of interest (see, for instance, Savastano (1995) and Schmidt-Hebbel and others (1992) for a review of this literature). Some researchers have been unable to detect much of an effect of changes in real interest rates on domestic saving in developing countries. For example, Giovannini (1985) finds that in only 5 of the 18 developing countries in his sample are consumption and saving sensitive to changes in the real interest rate. He concludes that, for the majority of cases, "the response of consumption growth to the real rate of interest is insignificantly different from zero" (p. 215) and that one should therefore expect "negligible responses of aggregate saving to the real rate of interest [in developing countries]" (p. 197). Rossi (1988) also finds that "increases in the real rate of return are not likely to elicit substantial increases in savings, especially in low-income developing countries" (p. 104). In a model with a single consumption good, Ostry and Reinhart (1992) confirm these findings but when a disaggregated commodity structure that allows for traded and nontraded goods is assumed, these authors find higher and statistically significant estimates of the intertemporal elasticity of substitution. However, regional differences emerge, with Asian countries showing a greater responsiveness to real interest rate changes.[2]

The finding of a zero or near-zero interest rate sensitivity of saving in a number of developing countries has led researchers to consider a number of alternative hypotheses that could help to explain this result. One such hypothesis is that consumption in developing countries may be more related to subsistence considerations—particularly in the case of low-income countries—than to intertemporal consumption smoothing.[3] If households must first achieve a subsistence consumption level, letting intertemporal considerations guide their decisions only for that portion of their budget left after subsistence has been satisfied, then the intertemporal elasticity of substitution and the interest-rate sensitivity of private saving will be close to zero for countries at or near subsistence consumption levels, and will rise thereafter.

[1] For a view running contrary to the McKinnon-Shaw hypothesis, see van Wijnbergen (1983). For an analysis of the Uruguayan experience with financial liberalization, see de Melo and Tybout (1986).

[2] Ostry and Reinhart (1992) attributed these differences to the presence of more binding liquidity constraints in Africa and Latin America than in Asia. Using a reduced form approach, similar regional differences in the interest-rate sensitivity of saving were found by Gupta (1987).

[3] For models that stress the role played by subsistence considerations in consumption/saving decisions, see Rebelo (1992) and Easterly (1994).

A second reason why the intertemporal elasticity of substitution may be lower for low-income countries concerns the relative share of *necessities* in the budgets of relatively poor households. If necessities (for example, food) are less substitutable through time than other goods, then the intertemporal elasticity of substitution will be lower for households with a larger proportion of necessities in their budgets than for households in which such goods are less important. The implication is that for relatively poor countries, where budget shares of food are relatively high, the interest-rate elasticity of saving would be relatively low.[4]

It is indeed a well-established empirical regularity that food consumption accounts for a markedly lower share of total expenditure in high-income than in low-income countries (see Mitchell and Ingco (1993) and Putnan and Allshouse (1993)). As shown in Table 1, food consumption accounts for less than 20 percent of total expenditure in most industrial countries and for only 8 percent of total consumption in the United States. For middle-income countries such as Mexico and Thailand, the share is often 30–40 percent, while for the poorer countries the share of food approaches 60–70 percent.

Empirical support for these hypotheses is found in Atkeson and Ogaki (1993). Specifically, using a panel of data on Indian households, they find that the intertemporal elasticity of substitution of the richest six households in their data set is approximately 1.6 times that of the poorest six households. Thus, from their results, the effects of the level of income on the intertemporal elasticity of substitution would indeed appear to be economically significant.

There are, however, additional reasons why saving may be less responsive to changes in real interest rates in low-income than in middle-income countries. Rossi (1988), for example, argues that low-income developing countries are characterized by pervasive liquidity constraints that imply that consumption growth in such countries is more likely to follow income growth than changes in expected rates of return.[5] The empirical evidence appears to point to the presence of liquidity constraints in many developing countries; however, Haque and Montiel (1989) highlight that the severity of these constraints varies considerably across countries. More recently,

[4] On the basis of a similar argument, Rebelo (1992) argues that financial liberalization in low-income developing countries is unlikely to produce large effects on saving and economic growth. For a discussion of the effects of financial market deregulation in developing countries, see Galbis (1993). For an analysis that highlights stylized facts concerning the differences in saving behavior between low- and middle-income developing countries, see two recent World Bank volumes (World Bank (1993 and 1994a)), dealing, respectively, with the performance of the economies of East Asia and Africa.

[5] Deaton (1989) has also emphasized the importance of liquidity constraints in explaining consumption/saving behavior in developing countries.

Table 1. *Food as a Percent of Total Personal Expenditure for Selected Countries (1990)*

Country	Percent
Low-income countries	
Honduras	44.5
India	50.6
Sierra Leone	68.0
Sri Lanka	51.3
Sudan	63.5
Average for group	55.6
Lower middle-income countries	
Colombia	29.5
Ecuador	32.4
Fiji	25.8
Jamaica	15.0
Jordan	39.8
Philippines	55.2
Thailand	27.3
Average for group	32.1
Upper middle-income countries	
Greece	32.4
Malaysia	25.8
Mexico	37.0
South Africa	28.6
Venezuela	28.6
Average for group	30.5
High-income countries	
Australia	14.8
Belgium	16.1
Canada	11.8
Sweden	15.4
United Kingdom	11.8
United States	8.0
Average for group	13.0

Sources: Putnam and Allshouse (1993); and Mitchell and Ingco (1993).
Note: For classification of economies by income level see World Bank (1994b).

Vaidyanathan (1993) shows that the incidence of liquidity constraints among households is inversely related to the degree of economic development, which would imply—following Rossi [(19]88)—that saving in poorer countries should be less responsive to intere[st ra]te changes.[6]

[6] Vaidyanathan (1993) also finds that financial li[beral]ization in developing countries reduced the severity of borrowing constraints. Although no direct tests were undertaken, the implication would be that financial liberalization, by reducing the fraction of households for which liquidity constraints are binding, should increase the interest-rate sensitivity of private saving. For a direct test of this hypothesis, see Bayoumi (1993) for the case of the United Kingdom, and Ostry and Levy (1995) for the case of France.

Lastly, it could be argued that failure to detect a systematic relationship between saving and real interest rates across countries and across time may also be due to considerable variation in the economic significance and information content of the real rates of return themselves. Lack of sophistication and depth in domestic financial markets or direct regulation may result in nominal interest rates that do not adequately reflect expectations about the underlying economic fundamentals. This problem may be particularly severe for some of the poorer countries. In addition, lack of information on inflation expectations is a problem that may, as the literature on the peso problem has shown, be especially relevant for high-inflation countries and/or episodes.

While a number of hypotheses have been put forward to *suggest* that saving behavior in developing countries is affected by the level of development, there has been little systematic empirical investigation of this issue at the macroeconomic level in the literature. Yet from a policy perspective, the issue is of central importance since, for example, the effects on investment and growth of a number of policy actions frequently undertaken by developing countries—including financial liberalization, the reduction of capital controls, and the transmission of fiscal and commercial policy changes to the current account—will all be governed to some degree by the responsiveness of saving to interest rate changes.[7]

With this in mind, the purpose of this paper is to quantify empirically the response of consumption/saving to changes in the real rate of interest. We proceed in two steps. First, we use macroeconomic data for a sample of countries with diverse income levels to estimate a model that allows the intertemporal elasticity of substitution to vary with the level of wealth. We then use the estimated parameters to calculate, in the context of a simple endogenous growth model, the elasticity of saving with respect to changes in the real rate of interest.

[7] In addition to policy-induced shocks, the transmission of a temporary terms of trade disturbance, through its effect on the *consumption rate of interest*, depends crucially on the sensitivity of saving to intertemporal relative prices (see, for example, Svensson and Razin (1983) and Ostry (1988)). In the area of commercial policies, noncredible liberalizations will also generate changes in consumption rates of interest as households may view the reduction in import prices (associated with tariff reductions) as a temporary phenomenon (see, for example, Razin and Svensson (1983), Calvo (1987, 1988, and 1989), Edwards and Ostry (1990), and Ostry (1991a and 1991b)). The extent to which such noncredible liberalizations will induce a consumption boom (and a current account deterioration) therefore depends on the intertemporal elasticity of substitution in consumption. The finding that this parameter is low in a number of countries may help to rationalize the empirical findings of Ostry and Rose (1992) that tariffs have little systematic effect on saving and current account behavior. Finally, the transmission of fiscal policy changes (which engender movements in domestic interest rates) to the current account will be governed in part by the responsiveness of private saving to real interest rates (see, for example, Frenkel and Razin (1992) and Ostry (1994)).

The remainder of the paper is organized as follows. Section I presents some stylized facts on saving behavior and income levels and focuses on the differences between low- and middle-income developing countries. Section II describes the analytical framework, while Section III discusses the empirical methodology and summarizes the estimation results. The response of saving to changes in real interest rates is examined in the context of a simple endogenous growth model in Section IV. The implications for policy and future research are taken up in the final section.

I. Stylized Facts

A model, such as the one developed in this paper, that emphasizes the role of subsistence consumption has two main predictions about saving behavior. First, saving rates should increase with the level of wealth at the initial stages of development, with the largest increases in the saving rate occurring as a country moves from low- to middle-income levels.[8] Second, saving should become more responsive to changes in real interest rates as countries become richer. The first prediction follows directly from the role of subsistence consumption in the low-income developing countries, and the fact that the share of subsistence in total consumption declines with income.[9] The second prediction may also be related to subsistence considerations, since intertemporal incentives should only affect that portion of the budget left over after subsistence has been achieved, that is, discretionary income.

With regard to saving rates, a model that stresses the role of liquidity constraints offers a different prediction; if poor consumers cannot borrow but face an uncertain income stream, the demand for precautionary saving rises (see, for instance, Deaton (1989)).[10] Hence, it may be the poorest liquidity-constrained consumers that have relatively high saving rates. However, as with the subsistence model, the responsiveness of saving to changes in interest rates rises with the level of wealth as liquidity constraints become less binding.

Among countries in various income groups, the patterns of saving rates that emerge are broadly consistent with the predictions of the subsistence model. As Table 2 highlights for selected countries, private saving is, on average, considerably lower for the poorest developing countries, where the

[8] In fact, in such models (see Rebelo (1992) and Sarel (1996)), the saving rate need not increase in the transition from middle- to high-income levels.

[9] See, for example, Rebelo (1992).

[10] This precautionary motive, without explicitly modeling liquidity constraints, is empirically investigated in Ghosh and Ostry (1994).

Table 2. *Personal Saving Rates for Selected Countries*[a]
(1985–1993 averages, unless otherwise noted)

Country	GNP per equivalent adult in 1985 $ 1980–87 average	Personal savings as a percent of GDP
Low-income countries		
Tanzania[b]	639.5	−1.0
Burkina Faso[b]	644.6	1.0
Bangladesh	889.2	13.5
Madagascar	916.8	4.3
Togo	937.9	14.0
Somalia	1,146.4	6.2
Ghana	1,164.1	6.1
Haiti	1,210.4	4.5
Kenya	1,197.9	18.2
Sierra Leone	1,341.0	8.1
Nigeria	1,603.5	9.5
Pakistan	1,672.0	23.0
Honduras[c]	1,679.9	7.5
Guyana[c]	1,833.0	14.3
Sri Lanka	2,156.1	19.8
Egypt	2,158.3	29.8
Average for group	1,324.4	11.2
Lower middle-income countries		
Bolivia	2,047.9	12.2
Côte d'Ivoire[c]	2,057.6	12.7
Cameroon	2,170.4	11.0
El Salvador	2,203.0	15.0
Philippines[c]	2,432.0	16.2
Morocco[c]	2,472.4	20.9
Dominican Republic	2,811.4	14.8
Thailand	2,901.4	22.8
Paraguay	3,082.4	14.6
Tunisia[c]	3,773.4	14.5
Peru	3,786.5	24.4
Turkey	3,931.6	21.4
Iran	3,962.5	20.0
Colombia	4,164.0	12.7
Poland[c]	4,360.6	26.8
Chile	4,587.8	12.8
Average for group	2,805.8	17.1
Upper middle-income countries		
Mauritius	4,406.6	24.2
Korea	4,409.5	25.4
Argentina	4,994.5	18.5
Brazil[c]	5,099.8	17.4
Portugal	5,280.9	21.3
South Africa	5,770.9	22.8

Table 2. *(concluded)*

(1985–1993 averages, unless otherwise noted)

Country	GNP per equivalent adult in 1985 $ 1980–87 average	Personal savings as a percent of GDP
Upper middle-income countries *(continued)*		
Malaysia	5,824.4	18.6
Greece	6,232.5	25.8
Mexico	6,968.8	13.8
Venezuela[c]	7,672.1	11.4
Trinidad and Tobago[c]	11,161.0	15.1
Average for group	6,165.5	19.5
High-income countries		
Ireland	7,170.9	22.0
Spain	7,477.8	20.7
Israel	10,572.9	16.9
Austria	11,147.3	23.3
United Kingdom	11,462.6	15.2
Italy	11,613.1	25.7
Belgium	11,675.1	23.2
Japan	11,819.9	25.5
Netherlands	12,013.8	24.9
Finland	12,019.5	19.4
France	12,775.6	19.1
Australia	13,841.5	18.8
Switzerland	16,079.1	23.5
Canada	16,529.3	21.5
United States	18,194.5	16.4
Average for group	12,292.9	21.1

Sources: IMF, *World Economic Outlook*; Savastano (1994); World Bank (1994a).
[a] For classification of economies by income level, see World Bank (1994b).
[b] Average for 1987–91 from World Bank (1994a).
[c] Average for 1985–92 from Savastano (1995).

saving rate is about one half that of the high-income group.[11] In fact, such differences also appear within regions. According to the World Bank (1994a), median gross domestic saving as a percentage of GDP (1987–91 average) was 5.6 percent for the low-income African countries and 19.0 percent for the middle-income African countries (the average was 7.7 percent).

As predicted, the relationship between the level of income and the saving rate appears to be nonlinear for the countries in our sample (see Rebelo

[11] We adopt the World Bank (1994b) classification of economies. See also Aghevli and others (1990).

(1992)); the largest increases in the saving rates occur in the transition from low- to lower middle-income, where the average personal saving rate rises by 5.9 percentage points (Table 2). The average for the upper middle-income countries is still 2.4 percentage points above that of the lower middle-income group, but there appears to be less of a difference (1.6 percentage points) between the average saving rates in the high-income and upper middle-income countries in our sample.

Still, there appear to be sharp differences in saving rates that cannot be accounted for by the subsistence model. For example, saving rates in Latin America are well below those observed in many Asian countries, despite similar income levels.[12] As pointed out in World Bank (1993), gross domestic saving as a percentage of GDP was nearly 40 percent for the high-performing (middle-income) Asian economies in 1990; it is argued that these relatively high saving rates were an engine of growth in many of these countries, since they financed higher rates of investment as well.

The role of real interest rates in saving behavior is more difficult to gauge. One problem—which is particularly important in Africa—is that financial markets remain thin and governments set interest rates, frequently at nonmarket levels. As pointed out in a recent World Bank study on Africa (World Bank (1994a)), "most countries [in Africa] have few banks, and . . . there is little scope for 'true' market-determination of interest rates" (p. 114). This feature of credit markets in low-income countries may itself make saving less responsive to interest rates.

Nevertheless, there is some evidence that financial savings increased as a result of the increase in real interest rates associated with liberalization of financial markets, both in Africa and elsewhere among developing countries. For example, among the Asian countries, the increase in real interest rates in Taiwan Province of China in 1949 contributed to a sharp increase in savings. Similar results were achieved by Indonesia and Korea and, more recently, by Argentina, Chile, Mexico, and Pakistan (see World Bank (1993) and the references therein). There is also some evidence that reform programs in Africa have succeeded in raising domestic savings. For example, as documented in World Bank (1994a), median gross domestic saving rates climbed 3.3 percentage points for the six African countries with a large improvement in macroeconomic policies (which consisted of five main indicators, one of which was interest rate policy), compared with a *decline* of

[12] Income distribution within a country is often argued to exert an independent influence on saving behavior, but we are unaware of any systematic empirical investigation of this hypothesis. Kaminsky and Pereira (forthcoming) do find evidence that social indicators, which reflect the skewness of income distribution, help explain the downward rigidity of consumption in Latin American countries relative to the Asian countries in their sample.

3.3 percentage points for countries with policy deterioration. In general, moving real interest rates from sharply negative to mildly positive levels seems to be a positive factor in mobilizing domestic savings, although there is no particular evidence that the effectiveness of such policies rises with the level of development, as the subsistence model would suggest. An investigation of this issue is undertaken in the following three sections.

II. Analytical Framework

We begin by describing the maximization problem faced by a representative household in a given country. Since our ultimate purpose is to examine *cross-country* differences in saving behavior, we do not assume that representative households in different countries have identical preferences. As in Ostry and Reinhart (1992), we adopt a two-good framework that distinguishes between traded and nontraded goods. As argued in that paper, it is important to estimate the intertemporal elasticity of substitution allowing for two goods in order to avoid biases arising from changes in the relative price of traded and nontraded goods (the real exchange rate). We allow for cross-country variation in both the intratemporal elasticity of substitution between traded and nontraded goods and the intertemporal elasticity of substitution. The assumption is that the latter varies systematically with the level of real income. We begin by describing, in equations (1)–(7) below, the optimization problem at the *national* level.[13]

Consider then an economy with an infinite-lived representative household whose objective is to choose a consumption stream that maximizes

$$\frac{\sigma}{(\sigma-1)} E_0 \sum_{t=0}^{\infty} \beta^t \left(am_t^{1-1/\varepsilon} + n_t^{1-1/\varepsilon}\right)^{\frac{1-1/\sigma}{1-1/\varepsilon}}, \qquad a, \beta, \varepsilon, \sigma > 0, \ \beta < 1, \tag{1}$$

subject to the series of budget constraints

$$p_t m_t + q_t n_t = p_t \overline{m}_t + q_t \overline{n}_t + \overline{x}_t + B_t - \left(1/R^*_{t-1}\right)B_{t-1}, \qquad \forall t \geq 0, \tag{2}$$

and the transversality condition

$$\lim_{t\to\infty} \prod\nolimits_{i=0}^{t} \left(1/R^*_i\right)B_t = 0, \tag{3}$$

where E_0 is the expectations operator conditional on information available at time 0; B_t denotes the real level of debt carried from period t to period

[13] For a fuller discussion of the underlying model at the national level, see Ostry and Reinhart (1992).

$t+1$ with B_{-1} given; $(1/R_t^*) - 1 = r_t^*$ is the real interest rate (in terms of the numeraire) on the debt so R_t^* is the associated world real discount factor; m (n) denotes consumption of importables (nontradables) and p (q) denotes the relative price of m (n) in terms of the numeraire; β is the subjective discount factor; and ε (σ) denotes the intratemporal (intertemporal) elasticity of substitution.[14] An intratemporal elasticity of substitution greater (less) than one implies gross substitutability (complementarity) between traded and nontraded goods; a value of unity corresponds to the logarithmic utility case. The intertemporal elasticity of substitution reflects the sensitivity of consumption (and therefore saving) to changes in intertemporal prices (i.e., the consumption rates of interest), with higher values indicating greater sensitivity.

The problem of the representative consumer in a given country is to choose an optimal sequence $\{m_t, n_t, B_t\}$ that maximizes equation (1) subject to equations (2) and (3). The first-order necessary conditions for an optimum are

$$E_t\left\{\frac{p_t}{R_t^* p_{t+1}}\left[\frac{am_{t+1}^{1-1/\varepsilon}+n_{t+1}^{1-1/\varepsilon}}{am_t^{1-1/\varepsilon}+n_t^{1-1/\varepsilon}}\right]^{\frac{\sigma-\varepsilon}{\sigma(\varepsilon-1)}}\left[\frac{m_{t+1}}{m_t}\right]^{-\frac{1}{\varepsilon}}\right\}=\frac{1}{\beta}, \tag{4}$$

$$E_t\left\{\frac{q_t}{R_t^* q_{t+1}}\left[\frac{am_{t+1}^{1-1/\varepsilon}+n_{t+1}^{1-1/\varepsilon}}{am_t^{1-1/\varepsilon}+n_t^{1-1/\varepsilon}}\right]^{\frac{\sigma-\varepsilon}{\sigma(\varepsilon-1)}}\left[\frac{n_{t+1}}{n_t}\right]^{-\frac{1}{\varepsilon}}\right\}=\frac{1}{\beta}, \tag{5}$$

and

$$a(n_t/m_t)^{1/\varepsilon} = p_t/q_t. \tag{6}$$

Equation (4) is the intertemporal Euler equation associated with importables consumption in two consecutive periods; it states that the marginal utility cost of giving up one unit of m at time t should be equated to the expected utility gain from consuming one more unit of m at $t+1$. Equation (5) is the analogous condition relating the marginal rate of substitution between consumption of good n at t and $t+1$ to the relevant intertemporal relative price. Finally, equation (6) is the nonstochastic first-order condition equating the intratemporal marginal rate of substitution between importables and nontradables to the corresponding relative price ratio. It can be verified that equations (4)–(6) are not independent. Specifically, combining

[14] Following the discussion in Section I, we allow for cross-country differences in the elasticities of substitution in the estimation.

equation (6) with either of the two remaining equations yields the third. Therefore, given that the nonstochastic first-order condition holds, equations (4) and (5) do not provide independent restrictions on the evolution of consumption through time.

The main difference between the model described above and standard Euler-equation models that have been applied previously to developing countries (for example, Giovannini (1985) and Rossi (1988)) relates to the goods structure. In the above model, we allow for consumption of tradables and nontradables since the latter appear to account for a large fraction of the consumption basket in many developing countries. Moreover, imposing a one-good structure can seriously bias the resulting estimates of preference parameters. If the real exchange rate (the relative price of traded and nontraded goods) varies over time, as is the case in most developed and developing countries, aggregate consumption will respond to those price changes. This is because a change in the real exchange rate alters the consumption rate of interest. As such channels are ignored in a one-good structure, the empirical results from such models may be biased.[15] For example, the finding that intertemporal elasticities of substitution are insignificantly different from zero in the majority of developing countries was found by Ostry and Reinhart (1992) not to be robust to a relaxation of the one-good assumption.

Given time-series data on importables and nontradables consumption, and on interest rates, and import, export, and nontradables prices, it is possible to estimate the system consisting of equations (4)–(6) and recover the main parameters of interest. Since equation (6) must hold identically (in the absence of measurement error), and since equations (4) and (5) are not independent given that equation (6) holds, we can eliminate equation (4) or (5) from the estimation. The restrictions on the joint behavior of consumption of importables and nontradables, the terms of trade, and the relevant rate of return, implied by the maximization of the expected utility function given by equation (1) subject to the constraints given in equations (2) and (3), are summarized in equation (4). In addition, given the assumption of rational expectations, we can use equation (4) to define the disturbance

$$u_t = \left\{ \frac{p_t}{R_t^* p_{t+1}} \left[\frac{a m_{t+1}^{1-1/\varepsilon} + n_{t+1}^{1-1/\varepsilon}}{a m_t^{1-1/\varepsilon} + n_t^{1-1/\varepsilon}} \right]^{\frac{\sigma-\varepsilon}{\sigma(\varepsilon-1)}} \left[\frac{m_{t+1}}{m_t} \right]^{-\frac{1}{\varepsilon}} \right\} - \frac{1}{\beta}, \tag{7}$$

where u_t must be uncorrelated with any variable that is in the information set of agents at time t.

[15] For a discussion of the bias issue, see Ostry and Reinhart (1992) and Ogaki and Reinhart (1995).

Now that we have fully described the optimization problem of a representative household in a given country, we must specify how intertemporal parameters governing saving behavior vary systematically across countries. In what follows, we take a particularly simple approach motivated by a Stone-Geary preference specification (as described, for example, by Rebelo (1992)). We adopt a specification in which the intertemporal elasticity of substitution is an increasing function of the gap between permanent income and the subsistence level of consumption, namely,

$$\sigma_i = \sigma\left(1 - \frac{\gamma}{y_i^p}\right), \tag{8}$$

where σ_i denotes the intertemporal elasticity of substitution in country i; γ is a constant that reflects subsistence consumption, and y_i^p is a measure of permanent income in country i. Clearly, equation (8) is similar to the Stone-Geary preference specification, with permanent income replacing consumption. Conceptually, we focus on income rather than consumption in order to make transparent the connection between the intertemporal elasticity of substitution and the level of development (i.e., income per capita). Equation (8) shows that the interest-rate sensitivity of aggregate consumption (as well as the level of saving itself) will be lower as the ratio, γ/y_i^p approaches unity and wealth is only sufficient to support a subsistence level of consumption.

III. Estimation Strategy and Empirical Results

We begin by describing our data set, before moving to a description of the econometric methodology.

Data Issues

The parameters of the representative household's utility function outlined in the previous section are estimated using annual time-series data for thirteen countries. The low-income countries in the sample are Egypt, Ghana, India, Pakistan, and Sri Lanka; the low-middle-income countries consist of Colombia, Costa Rica, Côte d'Ivoire, Morocco, and the Philippines; and three upper middle-income countries—Brazil, Korea, and Mexico—are also included in the analysis. Data coverage for each country begins in 1968 and ends anywhere between 1983 and 1992.

As equations (7) and (8) highlight, estimation of the intertemporal and intratemporal elasticities of substitution requires data on household consumption of traded and nontraded goods, on the terms of trade, and on permanent income or wealth. While time series on the terms of trade are readily available, consumption data are generally not disaggregated into traded and nontraded goods, and estimates of wealth are scarce. The methodology applied to disaggregate consumption is described in detail in Ostry and Reinhart (1992). Basically, we assume that the nontraded goods sector of the economy consists of public and private services while traded goods production emanates from the agriculture, mining, and industrial sectors. Consumption of importables can then be constructed using the supply-side data and data on exports and imports of consumer goods. Consumption of nontraded goods is calculated residually as total private consumption less consumption of importables. The relevant price deflators for the consumption of traded and nontraded goods are price indices for imports and services, respectively. Deposit rates of interest were used when available; in their absence, a money market rate was employed. Finally, as a proxy for permanent income, an average of real income per equivalent adult for the period 1980–87 was used.[16]

Methodology and Results

In this subsection, we apply Cooley and Ogaki's (forthcoming) two-step procedure to obtain estimates of the intratemporal elasticity of substitution between traded and nontraded goods and of the intertemporal elasticity of substitution. An alternative procedure described in Ostry and Reinhart (1992) uses a cointegration approach to obtain first-stage estimates of the intratemporal parameter and Hansen and Singleton's (1982) Generalized Method of Moments (GMM) approach to obtain (second-step) estimates of the intertemporal parameters.

The First Step: Estimating the Intratemporal Elasticity of Substitution

Many economic time series can be modeled as being difference stationary with drift. Under such conditions, the notions of stochastic and deterministic cointegration are useful in examining the interaction between two or more variables of interest.[17] Suppose that the components of a vector series $\mathbf{X}(t)$ are I(1) processes with drift; if a linear combination of $\mathbf{X}(t)$, say $\boldsymbol{\lambda}'\mathbf{X}(t)$, is *trend stationary*, the components of $\mathbf{X}(t)$ are said to be (stochas-

[16] All series are available from the authors upon request.

[17] Stochastic cointegration and the deterministic cointegration restrictions were defined by Ogaki and Park (1989) and Campbell and Perron (1991). Efficiency gains from estimating the cointegrating vectors by imposing the deterministic cointegration restriction were discussed by West (1988) for the one-regressor case and by Hansen (1992) and Park (1992) for the multiple-regressor case.

tically) cointegrated with a cointegrating vector $\boldsymbol{\lambda}$. Consider an additional restriction that the cointegrating vector eliminates the deterministic trend as well as the stochastic trend, so that $\boldsymbol{\lambda}'\mathbf{X}(t)$ is *stationary*. This restriction is called the deterministic cointegration restriction.

Standard unit root tests suggest that it is reasonable to model the relative price of traded and nontraded goods and the consumption ratio (of nontraded to traded goods) as I(1) processes.[18] We now focus on the relationship between these two variables. The intratemporal first-order condition (equation 6) implies that the log of the relative price and the log of the consumption ratio are cointegrated with the deterministic cointegration restriction.[19] If cointegration obtains, we can recover a consistent estimate of the intratemporal elasticity of substitution, ϵ.[20] To test for cointegration, we employ Park's (1992) Canonical Cointegrating Regression (CCR) procedure.[21]

There are several advantages associated with CCR. First, the estimated parameters, a and ε in this case, are not only consistent (as is OLS), but also asymptotically efficient and median unbiased. Second, unlike OLS, the asymptotic distributions of CCR estimators are nuisance-parameter free and normal conditioned on the regressors. This feature allows for the usual interpretation of the standard errors and therefore for hypothesis testing. Third, as Monte Carlo experiments show (see Park and Ogaki (1991)), CCR estimators have better small-sample properties than Johansen's estimators. This makes CCR particularly attractive for the present application, where the data are annual and the sample size is limited.

Rearranging terms, taking logs, and introducing a disturbance term (to allow, for example, for measurement error) equation (6) becomes

$$\ln(n_t / m_t) = \mathrm{e} \ln(a) + \mathrm{e} \ln(p_t / q_t) + u_t. \tag{9}$$

Equation (9) is likely to suffer from simultaneity bias, since relative prices are determined endogenously, and the error term may be serially correlated.[22] To correct for the potential presence of such nuisance parameters, long-run covariances are estimated and used to transform the data. As

[18] The results of the Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) unit root tests for all the series of interest are not reported but are available from the authors upon request.

[19] As shown by Hall (1978), consumption is a random walk when the real interest rate is assumed to be constant. Since we allow the real interest rate to vary over time, the first difference of the log of consumption can be serially correlated.

[20] This will also yield a point estimate for the parameter a, which is related to the consumption share.

[21] See Ogaki (1993a) for a more detailed explanation of CCR-based estimation and testing.

[22] Note that the OLS estimator is consistent but not efficient because of long-run simultaneity bias.

shown in Park (1992), the transformed data will be cointegrated with the same parameter vector as the original data.

Under the null of cointegration, an important property of the CCR procedure is that linear restrictions can be tested by χ^2 tests.[23] These χ^2 tests are used in a regression with spurious deterministic trends added to test for deterministic and stochastic cointegration. For this purpose, the CCR procedure is applied to a regression of the form

$$\ln(n_t / m_t) = \varepsilon \ln a + \sum_{i=1}^{q} \eta_i t^i + \varepsilon \ln(p_t / q_t) + u_t. \tag{10}$$

We let $H(p,q)$ denote the standard Wald statistic to test the hypothesis $\eta_{p+1} = \eta_{p+2} = \ldots \eta_q = 0$, where the variance of u_t has been replaced with elements of the long-run covariance matrix (see Park (1990)). Under the null of cointegration $H(p,q)$ converges to a χ^2_{q-p}. In particular, the H(0,1) statistic tests the deterministic cointegrating restriction, which is suggested by the intratemporal first-order condition. On the other hand, the $H(1,q)$ tests for stochastic cointegration.

Table 3 reports the CCR results.[24] As shown in column (1), with the exception of Korea, the intratemporal elasticity of substitution, ϵ, is estimated to be positive, consistent with our theoretical priors. In the case of Korea, where the estimate of ϵ is significantly negative, the H(0,1) test statistic presented in column (2) also rejects the specification implied by the model at the 0.1 percent level. It is possible that the assumption of homothetic preferences implied by the CES utility function is causing problems in this case.[25] For the remaining countries, the point estimates for ϵ range from a low of 0.38 to a high of 2.16; for 8 of the 13 countries the intratemporal elasticity of substitution exceeds unity, implying gross substitutability between traded and nontraded goods. For the Philippines, the point estimate of ϵ is positive but is not significantly so, and the H(0,1) test is significant

[23] This is in contrast to other cointegration tests, which have a null of *no cointegration*.

[24] We used Ogaki's (1993b) GAUSS CCR package for the CCR estimations. The CCR procedure requires an estimate of the long-run covariance of the disturbances in the system. We used Park and Ogaki's (1991) method with Andrews and Monahan's (1992) prewhitened HAC estimator with the QS kernel. A first-order VAR was used for prewhitening. We followed Andrews and Monahan (1992) with the maximum absolute value of the elements of Δ (their notation) set to 0.99. Andrews' (1991) automatic bandwidth estimator, S_T, was constructed by fitting an AR(1) to each disturbance.

[25] See, for example, Atkeson and Ogaki (1993), for an attempt to model nonhomothetic preferences with an extended addilog utility function. These authors assume time separability of preferences over two goods, in order to use an aggregation result over households. Finally, the failure to obtain cointegration could also arise from measurement error.

Table 3. *Canonical Cointegrating Regression Results*

Country (1)	ε (2)	$\varepsilon \ln(a)$ (3)	H(0,1) (4)	H(1,2) (5)	H(1,3) (6)	H(1,4) (7)
Brazil	2.156	−0.224	3.599	3.971	4.117	4.128
	(0.148)	(0.069)	(0.058)	(0.046)	(0.128)	(0.248)
Colombia	0.678	−0.512	5.363	0.687	0.836	1.207
	(0.214)	(0.037)	(0.021)	(0.407)	(0.658)	(0.751)
Costa Rica	1.132	−0.513	0.474	1.087	1.934	3.878
	(0.179)	(0.049)	(0.491)	(0.297)	(0.371)	(0.275)
Côte d'Ivoire	1.749	−1.424	1.147	3.078	3.730	3.884
	(0.208)	(0.219)	(0.284)	(0.079)	(0.155)	(0.274)
Egypt	0.440	−0.457	0.221	3.752	3.877	4.707
	(0.240)	(0.244)	(0.638)	(0.053)	(0.144)	(0.195)
Ghana	0.634	0.867	0.115	0.710	1.280	1.599
	(0.331)	(0.295)	(0.734)	(0.400)	(0.527)	(0.660)
India	1.547	1.666	1.828	1.227	1.637	2.339
	(0.410)	(0.109)	(0.176)	(0.268)	(0.441)	(0.505)
Korea	−5.257	−1.650	15.877	3.957	5.168	14.481
	(1.022)	(0.190)	(0.000)	(0.047)	(0.075)	(0.002)
Mexico	1.707	−0.914	0.005	0.000	2.670	3.009
	(0.414)	(0.131)	(0.946)	(0.992)	(9.263)	(9.390)
Morocco	1.070	−0.735	0.263	4.051	4.054	4.119
	(0.181)	(0.092)	(0.608)	(0.044)	(0.132)	(0.249)
Pakistan	2.075	0.054	0.558	2.256	3.942	4.215
	(0.267)	(0.081)	(0.455)	(0.133)	(0.139)	(0.239)
Philippines	0.382	0.142	7.989	0.695	2.040	2.304
	(0.333)	(0.088)	(0.005)	(0.404)	(0.361)	(0.512)
Sri Lanka	1.587	−0.421	0.669	0.674	1.737	1.887
	(0.157)	(0.088)	(0.414)	(0.419)	(0.419)	(0.596)

Note: Park and Ogaki's (1991) method with Andrews's (1991) automatic band-width parameter estimator was used to estimate long-run correlation.

at the 1 percent level, implying that the null of deterministic cointegration is rejected; however, the null of stochastic cointegration cannot be rejected (columns (3) and (4)) at standard significance levels. For the other countries, none of the H(0,1) test statistics are significant at the 1 percent level and only a few of them are significant at the 5 percent level, indicating that the null hypothesis of deterministic cointegration cannot be rejected.

To summarize, we find evidence that the CES specification of preferences is not rejected by the data for all the sample countries except Korea and, possibly, the Philippines. On this basis, we proceed with this specification when estimating the intertemporal elasticity of substitution.

The Second Step: Estimating the Intertemporal Parameters

In the first step, consistent estimates of the intratemporal parameters were obtained via a cointegration regression. In the second step, we impose the country-specific estimated values of ε and a and apply GMM to the Euler equation defined by equation (7) in order to obtain estimates of the intertemporal parameters. This two-step procedure does not alter the asymptotic distribution of the GMM estimators or test statistics because our cointegrating regression estimator is super-consistent and converges at a rate faster than $T^{1/2}$, where T is the sample size.

As discussed previously, we assume that the intertemporal elasticity of substitution of country i satisfies the Stone-Geary condition given in equation (8), reproduced below

$$\sigma_i = \sigma\left(1 - \frac{\gamma}{y_i^p}\right), \tag{11}$$

where y_i^p is a measure of permanent income of country i, and γ is a constant that reflects the subsistence level. In the GMM estimation, we fix γ and estimate σ. Since it is difficult to assess the real dollar value of the subsistence basket (i.e., γ), we allow γ to vary across a fairly broad range of values. Its minimum level of 100 is slightly below the value of US\$123 (1985 prices) found for India by Atkeson and Ogaki (1993) and its upper level of 400 was determined by the income level of the poorest countries since, for equation (9) to make sense, it must be that $(1 - \gamma/y_i^p) \geq 0$. In any case, the sensitivity of the results with respect to the choice of γ is reported.

We apply GMM to the Euler equation defined by equation (7) in a panel data set of countries, imposing equation (11) as a cross-country restriction.[26] For each country, a and ϵ are set at the values obtained in the first step cointegrating regression for the country in question. We restrict the discount factor, β, to be the same across countries in order to obtain more precise estimates of σ, which is the main parameter governing the responsiveness of saving to changes in the real rate of interest. Hence, in the second step, we estimate two parameters, σ and β. Because the sample size is not the same for all countries, we use the panel data estimation method described in Ogaki (1993d). Since there are many moment restrictions from many countries, we use only constants as instruments and avoid the use of lagged variables as instruments. Because lagged instrumental variables are not used, our method is robust to the time aggregation problem.

[26] Hansen/Heaton/Ogaki's GAUSS GMM package (see Ogaki (1993c)) is used for the GMM estimations in this paper.

Table 4 presents the second-step GMM results. We report results for twelve of the developing countries in our sample (all except Korea) for different values of γ. The panel excludes Korea because the CES specification is rejected by the data in this case. Our point estimates of σ are positive and significant while the point estimates of β are greater than one. Hansen's J-test statistics do not reject the overidentifying restrictions implied by the model at conventional levels. An attractive feature of the results is that they are not very sensitive to the choice of γ for the broad range used.

We recover the country-specific value of the intertemporal elasticity of substitution, σ', by employing the Stone-Geary specification given in equation (11). Figure 1 plots, for the sample countries, the intertemporal elasticity of substitution against the ratio of the country's permanent income to U.S. income, for the case of $\gamma = 400$. The intertemporal elasticity of substitution rises markedly with the level of income when low- to middle-income countries are compared; the change from middle- to high-income levels makes less of a difference. The nonlinearities that were evident in the relationship between the saving rate and income are also present here.

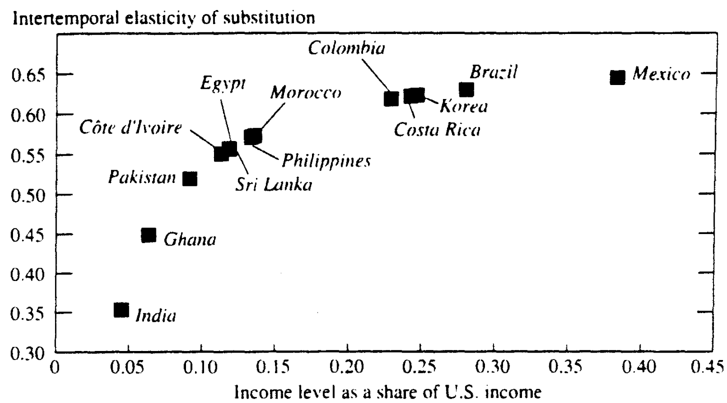
To examine the implications of our results for a broader set of countries, we use the estimated value of σ and measure permanent income in the same way as for "within-sample" countries. In column (4) of Table 5, we provide the estimate of each country's σ_i for the case where $\gamma = 400$. In addition, we report (in columns (3) and (5)) the range of values that result from adding and subtracting one standard error from the point estimate. As Table 5 makes clear, the range of variation of σ_i is wide, from a low value of about 0.05 for Uganda and Ethiopia, the poorest countries in our sample, to a high of about

Table 4. *The Intertemporal Parameters: Generalized Method of Moments Results*

Panel size	γ (1)	σ (2)	β (3)	J_T (4)	Degrees of freedom (5)
12 countries	100	0.596	1.051	14.600	10
		(0.208)	(0.015)	(0.147)	
12 countries	250	0.606	1.055	13.531	10
		(0.201)	(0.020)	(0.195)	
12 countries	300	0.615	1.056	13.312	10
		(0.201)	(0.015)	(0.207)	
12 countries	350	0.628	1.057	13.180	10
		(0.202)	(0.020)	(0.214)	
12 countries	400	0.646	1.057	13.151	10
		(0.204)	(0.020)	(0.215)	

Notes: In columns (2) and (3), the standard errors are in parentheses. Column (4) reports Hansen's J-test and the corresponding asymptotic *P*-values are given in parentheses. The last column gives the degrees of freedom for Hansen's J-statistic.

Figure 1. *The Intertemporal Elasticity of Substitution for Sample Countries*



Notes: The point estimates for the intertemporal elasticity of substitution and the income level and as a share of U.S. income are taken from columns (4) and (2), respectively, of Table 5.

0.64 for the United States and several other high-income countries. These estimates are in line with several earlier studies: Atkeson and Ogaki (1993), who examine Indian panel data; Ostry and Reinhart (1992), who focus on regional patterns; and Reinhart and Végh (1995), who apply a monetary model to several chronic-inflation countries. The implications for saving behavior of these cross-country differences are taken up in the next section.

IV. Saving and the Rate of Interest

Having estimated the key parameters that characterize household consumption and saving behavior, we now turn to the implications of our estimates for the interaction between saving and the real rate of return. As noted earlier, the attempt to encourage saving by raising real interest rates is at the heart of adjustment programs in a number of low- and middle-income developing countries. Higher saving, it is argued, can finance higher investment and lead to faster growth. In addition, the external current account effects of fiscal policy changes that alter domestic interest rates will be sensitive to the elasticity of private saving with respect to changes in the rate of return. To examine the interactions between interest rates, saving, investment, and growth, it is necessary to have some model in which these variables are determined endogenously. Rebelo's (1991) endogenous

Table 5. *The Intertemporal Elasticity of Substitution: Low- and Lower Middle-Income Countries*

	GNP per equivalent adult in 1985$ (1980–87 average)		Intertemporal elasticity of substitution		
Country	Level (1)	As a share of U.S. level (2)	Lower bound (3)	Point estimate (4)	Upper bound (5)
Low-income countries					
Uganda[a]	430.8	0.024	0.032	0.047	0.061
Ethiopia[b]	435.1	0.024	0.036	0.052	0.069
Zaïre	448.4	0.025	0.049	0.070	0.092
Chad[a]	593.0	0.033	0.146	0.212	0.277
Tanzania	639.5	0.035	0.169	0.243	0.318
Guinea[a]	640.3	0.035	0.169	0.244	0.319
Burkina Faso	644.6	0.035	0.171	0.247	0.323
Mali	644.8	0.035	0.171	0.247	0.323
Burundi	671.3	0.037	0.182	0.263	0.343
Malawi	727.9	0.040	0.203	0.293	0.383
Myanmar[a]	768.2	0.042	0.216	0.312	0.407
India*	829.1	0.046	0.233	0.336	0.440
Bangladesh[a]	889.2	0.049	0.248	0.358	0.468
Niger	897.6	0.049	0.249	0.360	0.471
Central African Republic	898.1	0.049	0.250	0.361	0.471
Afghanistan[a]	907.8	0.050	0.252	0.364	0.475
Nepal[a]	909.2	0.050	0.252	0.364	0.476
Madagascar	916.8	0.050	0.254	0.366	0.479
Togo	937.9	0.052	0.258	0.373	0.487
Gambia[a]	940.0	0.052	0.259	0.373	0.488
Rwanda	962.3	0.053	0.263	0.380	0.497
Zambia	1,054.0	0.058	0.279	0.403	0.527
Somalia	1,146.4	0.063	0.293	0.423	0.553
Ghana*	1,164.1	0.064	0.295	0.427	0.558
Kenya	1,197.9	0.066	0.300	0.433	0.566
Haiti	1,210.4	0.067	0.301	0.435	0.569
Sudan	1,300.8	0.071	0.312	0.450	0.589
Mozambique	1,342.3	0.074	0.316	0.456	0.597
Pakistan*	1,672.0	0.092	0.342	0.494	0.647
Sri Lanka*	2,156.1	0.119	0.367	0.529	0.692
Egypt*	2,158.3	0.119	0.367	0.530	0.692
Average for group	972.1	0.053	0.233	0.337	0.441
Average for the ten poorest countries	587.6	0.032	0.133	0.192	0.251
Lower middle-income countries					
Côte d'Ivoire*	2,057.6	0.113	0.363	0.524	0.685
Philippines*	2,432.0	0.134	0.376	0.543	0.710
Morocco*	2,472.4	0.136	0.377	0.545	0.712

Table 5. *(continued)*

Country	GNP per equivalent adult in 1985$ (1980–87 average): Level (1)	GNP per equivalent adult in 1985$ (1980–87 average): As a share of U.S. level (2)	Intertemporal elasticity of substitution: Lower bound (3)	Intertemporal elasticity of substitution: Point estimate (4)	Intertemporal elasticity of substitution: Upper bound (5)
Dominican Republic	2,811.4	0.155	0.386	0.558	0.729
Congo[b]	2,860.7	0.157	0.387	0.559	0.731
Thailand	2,901.4	0.159	0.388	0.560	0.733
Paraguay	3,082.4	0.169	0.392	0.566	0.740
Jordan	3,600.5	0.198	0.400	0.578	0.756
Ecuador	3,666.8	0.202	0.401	0.579	0.757
Tunisia	3,773.4	0.207	0.402	0.581	0.760
Peru	3,786.5	0.208	0.402	0.581	0.760
Turkey	3,931.6	0.216	0.404	0.584	0.764
Iran[a]	3,962.5	0.218	0.405	0.584	0.764
Algeria	3,993.5	0.219	0.405	0.585	0.765
Colombia*	4,164.0	0.229	0.407	0.588	0.768
Poland	4,360.6	0.240	0.409	0.590	0.772
Panama[b]	4,422.9	0.243	0.409	0.591	0.773
Costa Rica*	4,487.9	0.247	0.410	0.592	0.774
Chile	4,587.8	0.252	0.411	0.593	0.776
Fiji	4,605.4	0.253	0.411	0.594	0.776
Iraq[a]	5,092.5	0.280	0.415	0.599	0.783
Average for group	3,669.2	0.202	0.398	0.575	0.752
Upper middle-income countries					
Gabon[a]	4,990.5	0.274	0.414	0.598	0.782
Argentina	4,994.5	0.275	0.414	0.598	0.782
Brazil*	5,099.8	0.280	0.415	0.599	0.783
Taiwan	5,166.1	0.284	0.415	0.600	0.784
Yugoslavia	5,207.6	0.286	0.415	0.600	0.785
Portugal	5,280.9	0.290	0.416	0.601	0.786
South Africa	5,770.9	0.317	0.419	0.605	0.791
Malaysia	5,824.4	0.320	0.419	0.605	0.792
Hungary	5,883.3	0.323	0.419	0.606	0.792
Uruguay	5,926.6	0.326	0.420	0.606	0.793
Greece	6,232.5	0.343	0.421	0.608	0.795
Malta	6,282.9	0.345	0.421	0.609	0.796
Syria	6,667.8	0.366	0.423	0.611	0.799
Mexico*	6,968.8	0.383	0.424	0.613	0.801
Venezuela	7,672.1	0.422	0.427	0.616	0.806
Average for group	5,864.6	0.322	0.419	0.605	0.791
High-income countries					
Ireland	7,170.9	0.394	0.425	0.614	0.803
Israel	10,572.9	0.581	0.433	0.625	0.818
Singapore[a]	10,966.2	0.603	0.434	0.626	0.819
United Kingdom	11,462.6	0.630	0.434	0.627	0.820

Table 5. *(concluded)*

Country	GNP per equivalent adult in 1985 $ (1980–87 average)		Intertemporal elasticity of substitution		
	Level (1)	As a share of U.S. level (2)	Lower bound (3)	Point estimate (4)	Upper bound (5)
Hong Kong	11,474.9	0.631	0.434	0.627	0.820
Italy	11,613.1	0.638	0.435	0.628	0.821
Japan	11,819.9	0.650	0.435	0.628	0.821
Denmark	12,406.6	0.682	0.435	0.629	0.823
France	12,775.6	0.702	0.436	0.630	0.823
Sweden	12,940.9	0.711	0.436	0.630	0.824
Australia	13,841.5	0.761	0.437	0.631	0.825
Iceland	14,087.9	0.774	0.437	0.632	0.826
Norway	14,408.3	0.792	0.438	0.632	0.826
Switzerland	16,079.1	0.884	0.439	0.634	0.829
Canada	16,529.3	0.908	0.439	0.634	0.829
United States*	18,194.5	1.000	0.440	0.636	0.831
Kuwait[b]	20,033.0	1.101	0.441	0.637	0.833
United Arab Emirates[a]	30,904.5	1.699	0.444	0.642	0.839
Average for group	14,712.4	0.786	0.436	0.631	0.825

[a] Average for 1980–85.
[b] Average for 1980–86.
Notes: The lower (upper) bounds are constructed by subtracting (adding) one standard error to the point estimate. An asterisk denotes that the country was included in the sample.

growth model is particularly well suited to the issue at hand because it allows us to calculate the effects of different values of the intertemporal elasticity of substitution on saving rates very easily.

A Simple Model of Endogenous Growth

Since the aim is illustrative, we make a series of simplifying assumptions and focus on a linearized, continuous-time version of Rebelo's (1991) model. The reader is referred to these papers for further details.

The household problem is outlined in equations (1)–(7), except that now we focus on a continuous-time version of the model and assume that the technology is such that the traded good can be transformed into the nontraded good at a constant rate, so that the equilibrium relative price of these goods is constant.[27] Hence, total consumption is given by

[27] For estimation purposes this simplifying assumption turns out to be too restrictive and is rejected by the data (see Ostry and Reinhart (1992)). However, the aim here is to describe the interaction between saving and real rates under the simplest of settings.

$$c_t = m_t + n_t. \tag{12}$$

Production takes place under a linear technology and employs a single type of capital good that is a composite of physical and human capital

$$y_t = Az_t, \tag{13}$$

where A represents the technology level and y_t (z_t) is output (capital). The linear technology, which is a common feature of a class of endogenous growth models (see Romer (1989)), ensures that the rate of return to capital does not decline as the capital stock increases. Finally, output is either consumed or used for capital accumulation, namely,

$$y_t = c_t + \dot{z}_t. \tag{14}$$

Production efficiency equates the marginal product of capital to the rate of return

$$A = 1 + r, \tag{15}$$

where r is the real rate of interest.

The standard equations of motion for consumption and the accumulation of capital that arise from the optimization problem are given by

$$\frac{\dot{c}_t}{c_t} = \sigma_i(r - \delta) \tag{16}$$

and

$$\frac{\dot{z}_t}{z_t} = \sigma_i(r - \delta), \tag{17}$$

where δ, which is greater than zero, represents the constant rate of time preference and σ_i is, as before, the intertemporal elasticity of substitution for country i.[28] Hence, in this economy all real variables grow at the same constant rate.

Saving is defined as

$$S_t = y_t - c_t = \sigma_i(r - \delta)z_t \tag{18}$$

and the saving rate, s_t, is given by

$$s_t = \sigma_i\left[1 - \frac{(1+\delta)}{(1+r)}\right]. \tag{19}$$

[28] Note that $\beta = 1/(1 + \delta)$.

To determine the response of the saving rate to changes in the real rate of interest, we differentiate equation (19) with respect to r and obtain

$$\frac{\partial s}{\partial r} = \sigma_i \frac{1+\delta}{(1+r)^2}. \tag{20}$$

Equation (20) highlights the key role that the intertemporal elasticity of substitution plays in determining how saving rates react to changes in real interest rates.[29] Specifically, as σ_i approaches zero, saving declines, growth declines, and saving ceases to respond to real interest rates altogether. In the remainder of this section, the estimated parameters are used to calculate—under the assumption that these economies share a common technology and thus face a common interest rate—the response of saving to real interest rate changes for a broad spectrum of countries with very different income levels.

Results

Using equation (20), the point estimates of σ_i reported in Table 5, and a variety of plausible values for the real rate of interest and the subjective discount factor, we can calculate the implied response of saving to changes in the real rate of interest under various scenarios.

There are several features of the results presented in Table 6 that are worth noting. First, as implied by this simple analytical framework, the intertemporal elasticity of substitution plays a central role in determining how much (or how little) saving rates respond to changes in the real rate of interest. Indeed, the saving elasticities presented in Table 6 closely resemble the σ_i reported earlier (Table 5). Second, it follows from the previous observation that the cross-country variation is wide. For the poorest countries, a one percentage point rise in the real rate of interest should elicit a rise in the saving rate of only about one tenth of one percentage point;[30] for the wealthiest countries, the rise in the saving rate in response to a similar change in the real interest rate is about two thirds of a percentage point.[31]

[29] In our context, these reactions should be thought of as steady-state effects.

[30] These results are consistent with the result that regressions of the saving rate against the real interest rate in low-income developing countries fail to find evidence of a significant coefficient on the rate of return variable: see, for example, Savastano (1994).

[31] Although the model used here ignores a number of important determinants of household behavior, it is interesting to note that in the context of a more complicated model that included, inter alia, the effects of financial deregulation on household saving behavior (see Ostry and Levy (1995)), the magnitude of the implied elasticities for the case of France is broadly similar to what is reported here.

Table 6. *The Interest Sensitivity of Saving: Low- and Lower Middle-Income Countries*[a]

Country	r = 0.03 discount factor = 0.01 (1)	r = 0.04 discount factor = 0.01 (2)	r = 0.05 discount factor = 0.01 (3)	r = 0.03 discount factor = 0.02 (4)	r = 0.03 discount factor = 0.03 (5)
Low-income countries					
Uganda[b]	0.044	0.043	0.043	0.045	0.045
Ethiopia[c]	0.050	0.049	0.048	0.050	0.051
Zaïre	0.067	0.065	0.064	0.067	0.068
Chad[b]	0.201	0.198	0.194	0.203	0.206
Tanzania	0.232	0.227	0.223	0.234	0.237
Guinea	0.232	0.228	0.224	0.235	0.237
Burkina Faso	0.235	0.230	0.226	0.237	0.240
Mali	0.235	0.230	0.226	0.237	0.240
Burundi	0.250	0.245	0.241	0.253	0.255
Malawi	0.279	0.273	0.268	0.282	0.285
Myanmar[b]	0.297	0.291	0.285	0.300	0.303
India*	0.320	0.314	0.308	0.324	0.327
Bangladesh[b]	0.340	0.334	0.328	0.344	0.347
Niger	0.343	0.337	0.330	0.347	0.350
Central African Republic	0.343	0.337	0.330	0.347	0.350
Afghanistan[b]	0.346	0.340	0.333	0.350	0.353
Nepal[b]	0.347	0.340	0.334	0.350	0.354
Madagascar	0.349	0.342	0.336	0.352	0.356
Togo	0.355	0.348	0.342	0.359	0.362
Gambia[b]	0.356	0.349	0.342	0.359	0.363
Rwanda	0.362	0.355	0.348	0.365	0.369
Zambia	0.384	0.377	0.370	0.388	0.392
Somalia	0.403	0.395	0.388	0.407	0.411
Ghana*	0.406	0.398	0.391	0.410	0.415

Table 6. *(continued)*

Country	r = 0.03 discount factor = 0.01 (1)	r = 0.04 discount factor = 0.01 (2)	r = 0.05 discount factor = 0.01 (3)	r = 0.03 discount factor = 0.02 (4)	r = 0.03 discount factor = 0.03 (5)
Kenya	0.412	0.404	0.397	0.416	0.421
Haiti	0.414	0.406	0.399	0.419	0.423
Sudan	0.429	0.420	0.412	0.433	0.437
Mozambique	0.434	0.426	0.418	0.439	0.443
Pakistan*	0.471	0.462	0.453	0.476	0.481
Sri Lanka*	0.504	0.494	0.485	0.509	0.514
Egypt*	0.504	0.495	0.485	0.509	0.515
Average for group	0.321	0.315	0.309	0.324	0.327
Average for ten poorest countries	0.183	0.179	0.176	0.184	0.186
Lower middle-income countries					
Côte d'Ivoire*	0.499	0.489	0.480	0.504	0.509
Philippines*	0.517	0.507	0.498	0.522	0.528
Morocco*	0.519	0.509	0.499	0.524	0.529
Dominican Republic	0.531	0.521	0.511	0.536	0.542
Congo[c]	0.532	0.522	0.512	0.538	0.543
Thailand	0.534	0.523	0.513	0.539	0.545
Paraguay	0.539	0.528	0.518	0.544	0.550
Jordan	0.550	0.540	0.529	0.556	0.561
Ecuador	0.551	0.541	0.531	0.557	0.563
Tunisia	0.553	0.543	0.532	0.559	0.565
Peru	0.553	0.543	0.533	0.559	0.565
Turkey	0.556	0.545	0.535	0.562	0.567

Iran[b]	0.556	0.546	0.535	0.562	0.568
Algeria	0.557	0.546	0.536	0.563	0.568
Colombia*	0.559	0.549	0.538	0.565	0.571
Poland	0.562	0.551	0.541	0.568	0.574
Panama[c]	0.563	0.552	0.542	0.569	0.575
Costa Rica*	0.564	0.553	0.542	0.569	0.575
Chile	0.565	0.554	0.544	0.571	0.577
Fiji	0.565	0.554	0.544	0.571	0.577
Iraq*	0.570	0.559	0.549	0.576	0.582
Average for group	0.547	0.537	0.527	0.553	0.559
Upper middle-income countries					
Gabon[b]	0.569	0.558	0.548	0.575	0.581
Argentina	0.569	0.558	0.548	0.575	0.581
Brazil*	0.570	0.559	0.549	0.576	0.582
Taiwan	0.571	0.560	0.549	0.577	0.583
Yugoslavia	0.571	0.560	0.550	0.577	0.583
Portugal	0.572	0.561	0.550	0.578	0.584
South Africa	0.576	0.565	0.554	0.582	0.588
Malaysia	0.576	0.565	0.555	0.582	0.588
Hungary	0.577	0.566	0.555	0.583	0.589
Uruguay	0.577	0.566	0.555	0.583	0.589
Greece	0.579	0.568	0.557	0.585	0.591
Malta	0.579	0.568	0.558	0.585	0.591
Syria	0.582	0.571	0.560	0.588	0.594
Mexico*	0.583	0.572	0.561	0.589	0.595
Venezuela	0.587	0.575	0.564	0.593	0.599
Average for group	0.576	0.565	0.554	0.582	0.588

Table 6. *(concluded)*

Country	r = 0.03 discount factor = 0.01 (1)	r = 0.04 discount factor = 0.01 (2)	r = 0.05 discount factor = 0.01 (3)	r = 0.03 discount factor = 0.02 (4)	r = 0.03 discount factor = 0.03 (5)
High-income countries					
Ireland	0.584	0.573	0.562	0.590	0.596
Israel	0.595	0.584	0.573	0.602	0.608
Singapore[b]	0.596	0.585	0.574	0.602	0.609
United Kingdom	0.597	0.586	0.575	0.603	0.610
Hong Kong	0.597	0.586	0.575	0.603	0.610
Italy	0.598	0.586	0.575	0.604	0.610
Japan	0.598	0.586	0.575	0.604	0.610
Denmark	0.599	0.587	0.576	0.605	0.611
France	0.599	0.588	0.577	0.606	0.612
Sweden	0.600	0.588	0.577	0.606	0.612
Australia	0.601	0.589	0.578	0.607	0.613
Iceland	0.601	0.590	0.579	0.607	0.614
Norway	0.602	0.590	0.579	0.608	0.614
Switzerland	0.603	0.592	0.581	0.610	0.616
Canada	0.604	0.592	0.581	0.610	0.616
United States	0.605	0.594	0.582	0.611	0.618
Kuwait[c]	0.607	0.595	0.584	0.613	0.619
United Arab Emirates[b]	0.611	0.599	0.588	0.617	0.623
Average for group	0.601	0.589	0.578	0.607	0.613

[a] Effect (in percentage points) on the saving ratio of a 1 percentage point increase in the real interest rate. The lower (upper) bounds are constructed by subtracting (adding) one standard error to the point estimate. An asterisk denotes that the country was included in the sample.
[b] Average for 1980–85.
[c] Average for 1980–86.

Third, as columns (2) and (3) of Table 6 highlight, the saving elasticity is not very sensitive to the level of the real rate of interest assumed nor to the subjective discount factor (columns (4) and (5)).

V. Conclusions

This paper has sought to investigate the effect of the level of development on household saving behavior. The main issue with which we were concerned was the responsiveness of saving to interest rate changes, and whether there were any economically significant behavioral differences within a sample of countries at different stages of development. This issue was argued to be of relevance to policymakers because the investment and growth effects of, for example, financial liberalization, will depend on how responsive consumption/saving is to changes in real rates of return. Other policy questions—for example, the relationship between government deficits and the current account—will also depend on the responsiveness of private saving to real interest rates to the extent that changes in public (dis)saving alter domestic rates of return.

The main conclusion that emerged from our analysis was that much of the considerable cross-country variation in both the level of saving and the responsiveness of saving to the real rate of interest could be systematically explained by the country's income level. Specifically, the hypothesis that the saving rate and its sensitivity to interest rate changes were a rising function of income found strong empirical support. There is, of course, an often wide variation in saving behavior across countries with similar income levels that remains unaccounted for by the simple framework presented here. With these limitations in mind, however, our results may help to explain why the rising real interest rates that typically accompany financial liberalization have often failed to elicit an appreciable rise in private saving.[32] They may also shed some light on the wide cross-country variation in the response of the current account to fiscal policy changes that alter domestic interest rates.

The results presented here suggest that higher saving rates may not be forthcoming, even with relatively large increases in real interest rates, if the country in question is at the lower end of the income spectrum. In addition, the simple endogenous growth model presented in Section IV suggested that the growth effects of higher interest rates would also tend to be relatively small for relatively poor countries.

[32] Nonetheless, even among the low-income countries, a move from negative to positive real interest rates, even if it has little impact on saving, may still be desirable from the point of view of macroeconomic stabilization and improving the efficiency of investment.

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