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Private Saving and Terms of Trade Shocks

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The relationship between temporary terms of trade shocks and household saving in developing countries is examined. It is first shown that, from a theoretical standpoint, this relationship is ambiguous: private saving may rise or fall in response to a transitory terms of trade shock, depending on the values of the intertemporal elasticity of substitution and the intratemporal elasticity of substitution between traded and nontraded goods. Empirical estimates of these two parameters are obtained using data from a sample of 13 developing countries, and then used to draw implications for the response of private saving to transitory terms of trade shocks. [JEL E21, F32, F41, 010, 053, 054, 055]

THE TERMS OF TRADE have historically been one of the most important exogenous determinants of the external positions of developing countries. Over the past two decades, sharp fluctuations in world market prices for primary commodities and two oil shocks, which substantially increased the price of imported energy products for non-oil developing

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countries, were associated with increased variability in the saving, investment, and current account behavior of these countries.

The theoretical literature on the relationship between the terms of trade and the current account has focused almost exclusively on how terms of trade changes affect private saving, ignoring any additional effects on investment and public saving.[1] The traditional explanation—associated with the names of Harberger (1950) and Laursen and Metzler (1950)—suggests that an improvement in the terms of trade raises a country's real income level, measured as the purchasing power of its exports in world markets, and hence, on the assumption that the marginal propensity to consume is less than unity, raises private saving. Thus, the Harberger-Laursen-Metzler (HLM) effect, as it has become known, hypothesized that improvements in a country's terms of trade would be associated with increases in private saving, and conversely, adverse terms of trade shocks would reduce saving.

This view went largely unchallenged for nearly three decades, and was generally supported by the available empirical evidence. (See, for example, Khan and Knight (1983).) In the early 1980s, however, several studies re-examined the theoretical underpinnings of the HLM effect, a crucial building block of which was the Keynesian (static) relationship between consumption (or saving) and income. These studies, including, for example, those by Sachs (1981, 1982) and Svensson and Razin (1983), argued that household saving decisions should be derived from solutions to a dynamic optimization problem of choosing consumption levels at different points in time. As far as the HLM effect was concerned, the key insight provided by these models was that the relationship between the terms of trade and saving depended crucially on the expected duration of the terms of trade shock. For example, if households expected an improvement in the terms of trade to be permanent, then they would revise upward their estimate of permanent income in proportion to the increased purchasing power of their income today. Under the hypothesis that the marginal propensity to consume (save) out of permanent income is unity (zero), a permanent change in the terms of trade would therefore have no effect on saving, contrary to the HLM view.[2] By contrast, in a situation in which the improvement in the terms of trade was expected to be only temporary, the increase in permanent income would be smaller than the increase in current income, and saving would accordingly rise.

[1] For a discussion of investment effects of terms of trade changes in a somewhat different context, see Corden (1988).

[2] The view that permanent terms of trade shocks have no effect on the current account has been disputed by Obstfeld (1982), Ostry (1988), and, more recently, by Gavin (1990).

Therefore, the HLM hypothesis was satisfied for transitory terms of trade disturbances, but apparently not for permanent ones.

At the same time, the view that transitory changes in the terms of trade have unambiguous effects on private saving is misleading for two reasons. When a country experiences a temporary adverse terms of trade shock that raises the price of current imports relative to future imports, consumers have an incentive to postpone their purchases—that is, to save more. So, while *consumption-smoothing* considerations—the basis for the HLM effect—imply that private saving should decline in response to the temporary real income decline, the so-called *consumption-tilting* motives imply that private saving should increase as agents reduce current consumption in line with the increase in its relative price.[3] On these grounds alone, therefore, what happens to saving is theoretically ambiguous and depends on the relative magnitudes of the consumption-smoothing and tilting motives. The parameter governing this latter motive is the intertemporal elasticity of substitution. Relatively large values of this parameter imply that, in response to a given (transitory) movement in the terms of trade and, hence, in the intertemporal relative price (consumption rate of interest), consumers increase their saving by a relatively large amount; it follows that the larger is this elasticity, the greater is the increase (the smaller the fall) in private saving in response to a transitory adverse shock to the terms of trade.

In addition, however, when there are nontraded goods, an adverse terms of trade shock will lead consumers to substitute away from relatively expensive imports in favor of home goods, thereby bidding up their relative price. If the terms of trade shock is temporary, the resulting temporary real appreciation will contribute to a further increase in the consumption interest rate and, hence, a further increase in saving.[4] The parameter governing the switch from imports to home goods and, hence, the magnitude of the temporary real appreciation and increase in the consumption rate of interest is the intratemporal elasticity of substitution between tradables and nontradables. A relatively large value of this parameter implies a large increase in the consumption rate of interest and a commensurately large rise in saving. It may be concluded, therefore,

[3] A transitory adverse shock to the terms of trade raises the cost of current consumption relative to future consumption (the consumption rate of interest) because it temporarily raises the relative price of imports, which enters into the consumer price index. The latter, however, returns to its trend level once the terms of trade return to their trend level. For further details on the consumption rate of interest, see Dornbusch (1983).

[4] See Ostry (1988). The reason is the same as given in the previous footnote. The transitory rise in the relative price of nontradables raises the consumer price index temporarily, making current goods more expensive relative to future goods.

that the larger are either the intertemporal or intratemporal elasticities of substitution, the greater will be the increase (the smaller the decrease) in private saving in response to a temporary adverse movement in the terms of trade. The outcome in any case is an empirical matter that can only be addressed through estimation of these two critical parameters.

The approach taken in this paper involves estimating the "structural" parameters of a representative household's utility function. The basis for such an approach, in preference to the alternative of estimating reduced-form consumption or saving functions, is related to the Lucas critique. As is well recognized, the Lucas critique implies that there may not be anything that could properly be called a consumption or saving function, in the sense of a stable functional relationship that is independent of the wider macroeconomic context.[5] In contrast to previous studies, we employ a disaggregated commodity structure according to which agents consume both traded and nontraded goods. Disaggregation permits estimation of the two parameters of interest: the intertemporal elasticity of substitution and the intratemporal elasticity of substitution between tradables and nontradables. The data set employed is also suitable for comparing our findings to those of previous studies that employed a one-good structure. In contrast to many such studies, we find evidence that the intertemporal elasticity of substitution is significantly different from zero and lies in the 0.3 to 0.8 range, depending on the region considered. Intratemporal substitution elasticities are estimated to be significantly higher, and indicate that this parameter—which to our knowledge has been entirely ignored in previous Euler equation estimations for developing countries— plays a critical role in determining the sign and magnitude of the HLM effect in these countries.

Finally, although the empirical results of this paper can be used to analyze a variety of other issues—including the effects of permanent terms of trade shocks and the impact of trade reforms (which alter the internal terms of trade of the country that undertakes them)—we focus in what follows on temporary terms of trade shocks, mainly because recent empirical evidence relating to the developing countries suggests that the transitory component of such shocks is quantitatively important.[6] For instance, Cuddington and Urzua (1989) found that fully 60 percent of all shocks to commodity prices were of a temporary nature, and Mendoza (1992) reported a similar result relating to the terms of trade of developing countries.

[5] See Hall (1988, p. 340), for an elegant restatement of this view.

[6] On the usefulness of our estimates to the issue of permanent terms of trade shocks, see, for example, Ostry (1988), Gavin (1990), and Edwards and Ostry (1992); on their applicability to trade reform issues, see Calvo (1987), Ostry (1990, 1991, 1992), Edwards and Ostry (1990), and Ostry and Rose (1992).

The remainder of this paper is organized as follows. Section I illustrates the role of preference parameters in the HLM effect in the context of a simple two-period model that admits closed-form solutions. For the purposes of empirical implementation, however, Section II considers the stochastic, infinite-horizon version of this model and presents the optimality conditions for an intertemporal equilibrium model in which households consume both traded and nontraded goods. Section III describes the approach to estimation and presents the empirical results. The main conclusions are contained in Section IV.

I. A Simple Model of the HLM Effect

Consider a small open exchange economy where the representative household derives utility, C, in each period according to the following constant elasticity of substitution (CES) function:[7]

$$C = (am^{1-1/\epsilon} + n^{1-1/\epsilon})^{\frac{\epsilon}{\epsilon-1}}, \qquad a, \epsilon > 0, \tag{1}$$

where m (n) denotes consumption of importables (nontradables). Agents are assumed to live for two periods.[8] Intertemporal consumption decisions maximize the following CES utility function subject to constraints specified below:

$$U = (C_1^{1-1/\sigma} + \beta C_2^{1-1/\sigma})^{\frac{\sigma}{\sigma-1}}, \qquad \beta, \sigma > 0, \beta < 1, \tag{2}$$

where the subscripts 1 and 2 denote periods 1 and 2, respectively, and where β denotes the subjective discount factor. In equation (1) the parameter ϵ denotes the intratemporal elasticity of substitution between tradables (importables) and nontradables. Larger values of this parameter imply greater responsiveness to relative price (real exchange rate) changes. A value of unity corresponds to the logarithmic utility case, while values above (below) unity imply gross substitutability (complementarity). In equation (2) the parameter σ denotes the intertemporal elasticity of substitution. Larger values of this parameter imply greater

[7] The model of this section is a simplified version of the one developed in Ostry (1988). A stripped-down version is presented here only for the purposes of illustrating the role of preference parameters in the HLM effect. The model to be estimated empirically is presented in Section II.

[8] As is well known, the two-period assumption is not restrictive here, since the second period may represent the aggregation of a large (possibly infinite) number of future periods. The motivation for the two-period structure is that it allows us to obtain closed-form solutions for the response of private saving to terms of trade shocks, something that is precluded in the infinite-horizon version of the model developed later.

responsiveness to movements in intertemporal relative prices (consumption rates of interest). Equation (2) collapses to a logarithmic utility function when $\sigma = 1$.

Perfect capital mobility is assumed, and therefore the country faces a given (in terms of the numeraire) world interest rate.[9] All debts are required to be repaid by the end of the second period. These assumptions imply that the representative household maximizes equation (2) (given equation (1)), subject to the constraint that the present value of expenditures not exceed the present value of resources. The latter, which we refer to below as lifetime wealth, is assumed to take the form of a stream of endowments of tradable (importable and exportable) and nontradable goods. The solution to this optimization problem yields demands of the form

$$m_1 = m_1[p_1, q_1, P_1 C_1(R, W)] \tag{3a}$$

$$n_1 = n_1[p_1, q_1, P_1 C_1(R, W)] \tag{3b}$$

$$m_2 = m_2[p_2, q_2, P_2 C_2(R, W)] \tag{3c}$$

$$n_2 = n_2[p_2, q_2, P_2 C_2(R, W)], \tag{3d}$$

where p_i and q_i denote, respectively, the relative price of importables and nontradables, and P_i denotes the consumer price index in period i; R is the consumption discount factor, which is given by $R = 1/(1 + r)$, where r is the consumption rate of interest; and W is real wealth.[10] The consumption discount factor, R, is related to the world discount factor, R^*, according to

$$R = R^* P_2 / P_1. \tag{4}$$

Thus, the consumption discount factor takes into account that the relevant interest rate for intertemporal consumption decisions depends on the evolution of the relative price structure through time. Since the consumer price index in any period depends on the relative prices of

[9] Without loss of generality, the numeraire is taken to be the exportable good. For recent evidence supporting the view that developing countries, in general, can be characterized as financially open economies, see Haque and Montiel (1991).

[10] Although it is straightforward to obtain explicit solutions for the demand functions in this case, there is no particular interest in doing so. In equations (3a)–(3d), we have made use of the fact that the optimization problem as specified satisfies the assumptions necessary for two-stage budgeting (Goldman and Uzawa (1964)). Accordingly, demands in a given period depend only on relative prices in that period and aggregate spending in that period. The real value of aggregate spending, in turn, depends only on lifetime wealth and on the intertemporal relative price, R (the consumption discount factor, which is equal to 1 over 1 plus the consumption rate of interest).

importables and nontradables, the consumption and world discount factors will differ from one another whenever the terms of trade (the relative price of importables) or the real exchange rate (the reciprocal of the relative price of nontradables) is not expected to remain constant through time. To close the model, market-clearing conditions for nontradable goods are specified:

$$n_1[p_1, q_1, P_1 C_1(R, W)] = \bar{n}_1 \tag{5a}$$

$$n_2[p_2, q_2, P_2 C_2(R, W)] = \bar{n}_2, \tag{5b}$$

where $\bar{n}_i$ represents the endowment of nontradable goods in period i. Finally, we can define the ratio of private saving to GDP (gross domestic product) (s) as follows:[11]

$$s = \frac{\bar{x}_1 - p_1(m_1 - \bar{m}_1)}{\bar{x}_1 + p_1\bar{m}_1 + q_1\bar{n}_1}, \tag{6}$$

where we have used the market-clearing conditions for home goods and where $\bar{x}_1, \bar{m}_1$ represent the endowments of the exportable and importable goods, respectively, in the first period.

Consider now the effect of a transitory deterioration in the terms of trade—that is, a rise in p_1, with p_2 constant. To simplify the analysis, it is convenient to consider an initially stationary equilibrium in which all prices and quantities are constant over time. Differentiating equation (6) around an initial equilibrium with $s = 0$, and using (5a) and (5b) to solve for the effects on q_1 and q_2 gives

$$\frac{ds}{d \log p_1} = \frac{b(1-k)\epsilon\sigma}{b\epsilon + (1-b)\sigma} - b(1-k)\chi, \tag{7}$$

where b is the initial expenditure share on importables (a positive fraction), k is the ratio of current to lifetime spending or wealth, and χ is the ratio of exports to production of tradables.[12] The first term on the right-hand side of equation (7) represents the intertemporal substitution effect, which is equal to (minus) the product of the elasticity of current spending with respect to the consumption discount factor, $(1 - k)\sigma$, and the change in the discount factor, $b\epsilon/(b\epsilon + (1 - b)\sigma)$. This expression is increasing in both ϵ and σ, which shows that saving rises by more, the

[11] Under the assumption of no historical debt commitments, this ratio is also equal to the ratio of the current account balance to GDP, since there is no investment or government saving in the model.

[12] Clearly, both k and χ are positive fractions. If the horizon of households were infinite, a good proxy for k would be the real interest rate. It should also be noted that if there is no domestic production (or endowment) of import substitutes, χ is equal to unity.

larger are either the intratemporal or intertemporal elasticities of substitution. For a given rise in the consumption interest rate, larger values of σ imply larger increases in saving. For a given elasticity of saving, larger values of ϵ imply larger increases in the consumption rate of interest and, hence, larger increases in saving. The second term on the right-hand side of equation (7) represents the consumption-smoothing effect, which depends on the initial volume of exports. With real income falling below its trend level, the consumption-smoothing effect tends to reduce private saving. Equation (7) summarizes the main result of this section, namely that private saving will increase by more (fall by less) in response to a temporary deterioration in the terms of trade, the larger are either intertemporal or intratemporal elasticities of substitution.

II. The Stochastic Euler Equations

The model of Section I was presented in order to illustrate the role of preference parameters in the HLM effect. With a view toward empirical implementation, however, we need to generalize that model by allowing for uncertainty and more than two periods. Accordingly, consider an economy with an infinitely lived representative household whose objective is to choose a consumption stream that maximizes

$$[\sigma/(\sigma - 1)]\, E_0 \sum_{t=0}^{\infty} \beta^t (am_t^{1-1/\epsilon} + n_t^{1-1/\epsilon})^{\frac{1-1/\sigma}{1-1/\epsilon}},$$

$$a, \beta, \epsilon, \sigma > 0,\ \beta < 1, \tag{8}$$

subject to the series of budget constraints

$$p_t m_t + q_t n_t = p_t \bar{m}_t + q_t \bar{n}_t + \bar{x}_t + A_t - (1/R^*_{t-1}) A_{t-1}, \qquad \forall_t \geq 0, \tag{9}$$

and the transversality condition[13]

$$\lim_{t\to\infty} \prod_{i=0}^{t} (1/R_i^*) A_t = 0, \tag{10}$$

where E_0 is the expectations operator conditional on information available at time 0; A_t denotes the real level of debt carried from period t to period $t+1$;[14] $(1/R_t^*) - 1 = r_t^*$ is the real interest rate (in terms of

[13] See Chamberlain and Wilson (1984) for a fuller discussion of the appropriate no ponzi-game constraint in an infinite-horizon consumption model under uncertainty.

[14] We assume that the inherited level of debt, A_{-1}, is given and, for convenience, set equal to zero.

the numeraire) on the debt;[15] and remaining notation is as specified in Section I.

The problem of the consumer, then, is to choose an optimal sequence (m_t, n_t, A_t) that maximizes equation (8), subject to equations (9) and (10). The first-order necessary conditions for an optimum are

$$E_t\left\{\frac{p_t}{R_t^* p_{t+1}}\left[\frac{am_{t+1}^{1-1/\epsilon}+n_{t+1}^{1-1/\epsilon}}{am_t^{1-1/\epsilon}+n_t^{1-1/\epsilon}}\right]^{\frac{\sigma-\epsilon}{\sigma(\epsilon-1)}}\left[\frac{m_{t+1}}{m_t}\right]^{-\frac{1}{\epsilon}}\right\}=\frac{1}{\beta} \tag{11}$$

$$E_t\left\{\frac{q_t}{R_t^* q_{t+1}}\left[\frac{am_{t+1}^{1-1/\epsilon}+n_{t+1}^{1-1/\epsilon}}{am_t^{1-1/\epsilon}+n_t^{1-1/\epsilon}}\right]^{\frac{\sigma-\epsilon}{\sigma(\epsilon-1)}}\left[\frac{n_{t+1}}{n_t}\right]^{-\frac{1}{\epsilon}}\right\}=\frac{1}{\beta} \tag{12}$$

$$a\,(n_t/m_t)^{1/\epsilon} = p_t/q_t. \tag{13}$$

Equation (11) 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 (12) 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 (13) 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 (11)–(13) are not independent. Specifically, combining (13) with either of the two remaining equations yields the third. Therefore, given that the nonstochastic first-order condition holds, equations (11) and (12) do not provide independent restrictions on the evolution of consumption through time.

It is perhaps worth emphasizing the differences between equations (11) and (12) and the corresponding condition in a one-good consumption model.[16] In such a model, the relative price ratio that is relevant for transforming present into future consumption is the real interest rate—that is, the nominal rate deflated by the rate of change of the aggregate price index. When relative prices are not constant, however, as when there are terms of trade shocks, the appropriate intertemporal relative price ratio needs to take account of such changes. This is why, for example, in equation (11) the price ratio that premultiplies the marginal rate of substitution inside the expectation sign is the real rate of interest

[15] Clearly, then, R_t^* is the associated world real discount factor.
[16] This is particularly relevant, since estimation of consumption Euler equations for developing countries has been confined to environments in which there is a single consumption good; see, for example, Giovannini (1985) and Rossi (1988).

in terms of the numeraire, $1/R_t^*$, adjusted for the rate of change of the terms of trade over time, p_t/p_{t+1} (that is, the "own" rate of interest). If the relative price of imports is expected to decline through time, current importables consumption is expensive relative to future importables consumption. In consequence, offsetting changes in the marginal rate of substitution are required in precisely the same direction as would occur if the world rate of interest were to rise (R^* were to fall). An analogous interpretation carries over to equation (12), wherein the appropriate relative price for the purpose of determining the marginal rate of substitution between nontradables consumption in consecutive periods involves the real exchange rate ratio, q_t/q_{t+1}.

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 (11)–(13) and recover the main parameters of interest. Since (13) must hold identically (in the absence of measurement error), and since (11) and (12) are not independent, given that (13) holds, it is sufficient in the estimation to consider equation (11) alone. 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 (8), subject to the constraints given in (9) and (10), are summarized in equation (11). In addition, given the assumption of rational expectations, we can use equation (11) to define the disturbance

$$u_t = \left\{ \frac{p_t}{R_t^* p_{t+1}} \left[\frac{a m_{t+1}^{1-1/\epsilon} + n_{t+1}^{1-1/\epsilon}}{a m_t^{1-1/\epsilon} + n_t^{1-1/\epsilon}} \right]^{\frac{\sigma-\epsilon}{\sigma(\epsilon-1)}} \left[\frac{m_{t+1}}{m_t} \right]^{-\frac{1}{\epsilon}} \right\} - \frac{1}{\beta}, \tag{14}$$

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

III. Empirical Results

The parameters of the representative household's utility function outlined in the previous sections were estimated using annual pooled time-series, cross-section data for 13 developing countries. The countries examined in the analysis include four African countries—Egypt, Ghana, Côte d'Ivoire, and Morocco; five Asian countries—Sri Lanka, India, Korea, Pakistan, and the Philippines; and four Latin American countries—Brazil, Colombia, Costa Rica, and Mexico.

Data Issues

Data coverage for each country begins in 1968 and ends anywhere between 1983 and 1987; see the Appendix for a list of the sample period for each country and the sources of the data.

As equation (14) highlights, estimation of the intertemporal and intratemporal elasticities of substitution requires data on household consumption of traded and nontraded goods and the terms of trade. While time series on the terms of trade are readily available (see Appendix), consumption data are generally not disaggregated into traded and nontraded components. Guided by the theoretical framework, these series were constructed using data from a variety of sources.[17]

The time series for consumption of importables was constructed as follows. The agricultural, mining, and industrial sectors produce traded goods; GDP originating in these sectors thus defines domestic production of traded goods. Private and public services comprise the nontraded goods sector. Domestic production of import substitutes is calculated as domestic production of traded goods less exports, on the assumption that exportables are not consumed at home.[18] If markets clear, all domestic production of import substitutes is consumed at home. Consumption of import substitutes plus consumer goods imports, which are total imports less imports of intermediate and capital goods, make up the series of interest—consumption of importables. Nontraded goods consumption is residually calculated as total private consumption less consumption of importables.[19]

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, and, in their absence, a money market rate. All consumption data are converted to a per capita basis by dividing the aggregates by the existing population.

Methodology

We estimate the parameter vector $\mu = [\beta, \epsilon, \sigma]$ by fitting the first-order condition defined in equation (14) to the panel data using Hansen's

[17] All series are available upon request.

[18] This may admittedly be a restrictive assumption in the case of some countries, but unfortunately, the data do not permit us to disaggregate consumption further.

[19] To ensure consistency, all the series used to disaggregate consumption into its traded and nontraded components (GDP by sector, private consumption, exports, and imports) are on a national income accounts (NIA) basis.

(1982) generalized method of moments (GMM).[20] The residuals in the estimated equation are partly forecast errors, which, by the assumption of rational expectations, are uncorrelated with any variable in the agent's information set at time t; in technical terms, those errors are orthogonal to any chosen instrument known to agents at time t. The assumption that all available information is used in forecasting future consumption and prices (that is, m_{t+1}, n_{t+1}, q_{t+1}, and p_{t+1}) allows us to use a large number of instruments to estimate a smaller number of parameters. That excess of instruments over estimable parameters yields a testable set of over-identifying restrictions. In reality, however, the error term may also include measurement error. Any systematic part of that noise—say, owing to serial correlation—should be allowed for in the estimation. Simply, it may be more efficient to fit the orthogonality condition less tightly in those periods when it is known measurment error swells the composite residual.

Understanding the complex nature of the disturbances—that is, the u_t—is critical to the estimation strategy. Serial correlation among the u_t may arise for a variety of reasons. First, as illustrated by Hayashi and Sims (1983), current values of consumption, m_t and n_t, may not be observed before expectations of future consumption (m_{t+1}, n_{t+1}) are formed, implying some lagged values of u_t are not in the information set; this may make today's forecast error correlated with last period's yet unobserved error.[21]

Second, the nature of the measurement error may make the residuals serially correlated; time aggregation problems in annual consumption data, as discussed in Hall (1988), introduce a first-order moving average process with a known parameter in the error term. Since the moving

[20] The parameter a is some positive number that denotes the weight attached to the imported good in the period utility function. In the analysis that follows, a (which is not of immediate interest), is not jointly estimated with the remaining parameters. Instead, the following values were used: 0.85 for Africa, 1.14 for Asia, and 0.58 for Latin America. These values were obtained by estimating the nonstochastic first-order condition (equation (13)) using ordinary least squares (OLS). Since we tested for and found cointegration among relative prices and consumption of importables and nontraded goods, we know that OLS provides consistent parameter estimates for a. By imposing in the subsequent estimation the values of a, we increase the efficiency of the estimates of the remaining parameters. The estimates of ϵ obtained by applying OLS to (13) were also used as the starting values in the subsequent GMM estimation.

[21] This problem is not likely to arise with prices and interest rates, which are generally available monthly with little or no lag. However, for the consumer making two-period forecasts of consumption, it is not unlikely that overestimating (underestimating) today's consumption level leads to a similar error in the subsequent period, making the two correlated.

average parameter is known, the constraint that the disturbances follow a first-order moving average process is taken into account by quasi-differencing the relevant series.[22] In addition, because of the diversity of countries included in our sample, as well as the fairly long period of coverage considered, we allow for the presence of more general forms of heteroscedasticity in the disturbances.

The estimation proceeds under the assumption that the parameters that characterize household preferences are identical across countries and regions. Although, as we will later show, homogeneity of tastes may be a restrictive assumption, it economizes on the number of parameters to be estimated and allows for the maximum degrees of freedom.[23] Two different sets of instruments are employed. Neither instrument set includes variables measured at time t, since the moving average process in the error term would result in these variables being correlated with the residual, u_t. The selection of instruments is not trivial, since the use of instruments that are correlated with the residual would result in inconsistent estimates. The most recent permissible instrument is one lagged two periods. The first vector of instrumental variables

$$z1_t' = [\text{constant}, m_{t-1}/m_{t-2}, n_{t-1}/n_{t-2}, p_{t-2}/(R_{t-2}^{*}p_{t-1}), m_{t-1}, n_{t-1}],$$

uses six instruments. This implies that there are six orthogonality conditions; with three parameters to be estimated, there are three overidentifying restrictions. The second instrument set replaces the levels of consumption of importables and nontraded with their ratio

$$z2_t' = [\text{constant}, m_{t-1}/m_{t-2}, n_{t-1}/n_{t-2}, p_{t-2}/(R_{t-2}^{*}p_{t-1}), m_{t-1}/n_{t-1}].$$

In the latter case there are five instruments, three free parameters (as before), and, therefore, two overidentifying restrictions. While the variation in the instrument set is slight, comparison of the estimates produced by each set sheds light on which parameters are most sensitive to the choice of instruments or, in other words, which parameter estimates are more robust.[24]

[22] For a complete discussion of how the moving average parameter is calculated, see Working (1960) and Hall (1988).

[23] This assumption will be relaxed later when regional estimates of the preference parameters are estimated.

[24] A common procedure in the existing literature on estimation of Euler equations is to allow the instrument set to vary by introducing more lags, considering instrument sets such as $z3_t' = [z1_t', z1_{t-1}']$. If one is working with time series for a single country, the added lag involves the loss of 1 degree of freedom. However, in the present analysis the added lag would entail the loss of 13 degrees of freedom (1 for each country). For this reason, we consider only the most parsimonious instrument sets.

Estimation Results

Table 1 reports the parameter estimates for each instrument set and the minimized value of the objective function, J, which Hansen and Singleton (1982) showed to be a test statistic for the validity of the overidentifying restrictions.[25] The parameter estimates for β, ϵ, and σ are similar for both instrument sets and are economically meaningful. The discount factor, β, falls in the 0.96–0.99 range. The intertemporal elasticity of substitution, σ, is in the 0.38–0.50 range but is large relative to its standard errors.[26] The intratemporal elasticity of substitution lies in the 1.22–1.27 range, indicating that importables are gross substitutes for nontraded goods.[27] The J-statistics are small relative to the degrees of freedom (for either instrument set), indicating that the overidentifying restrictions imposed by the model are not rejected by the data; that is, the three parameters estimated do a good job of satisfying either the five or six orthogonality conditions that depend on the instrument set. The quasi-differencing of the data and the correction for heteroscedasticity produced regression residuals that are white noise.

Of notable interest is the fact that, in contrast to previous work—including Giovannini (1985) and, to a lesser degree, Rossi (1988) for the developing countries, and Hall (1988) for the United States—the intertemporal elasticity of substitution is estimated to be significantly different from zero. This means that, in response to shifts in real (consumption-based) rates of interest, households would be expected to alter the time profile of their consumption, increasing the growth rate of the latter in response to an increase in real rates of return.

A possible explanation of our finding of a statistically significant degree of intertemporal substitution relates to the restrictive assumptions employed by previous researchers.[28] Specifically, for the most part, these studies either assumed the existence of a single consumption good—making no distinction between consumption of tradable and nontradable goods; or they assumed that standard consumption and price series were

[25] The J-statistic is distributed as $\chi^2(n)$ under the null hypothesis. The degrees of freedom, n, are equal to the number of overidentifying restrictions.

[26] Interestingly, these estimates are consistent with values in the 2.5–3.0 range for the coefficient of relative risk aversion (the reciprocal of the intertemporal elasticity of substitution) used in calibrating real business cycle models: see, for example, Stockman and Tesar (1990).

[27] This is slightly higher than the estimates obtained by Backus, Kehoe, and Kydland (1991) for the United States.

[28] Of related interest are recent empirical papers that have included money in modeling the consumer's choice problem: see, for example, Arrau (1990) and Eckstein and Leiderman (1992).

Table 1. *Estimates of the Model Pooling All Regions for Alternative Instrument Sets*

Parameters	All Countries	
	Instrument set I	Instrument set II
ϵ	1.279 (0.154)	1.223 (0.351)
σ	0.383 (0.087)	0.504 (0.228)
β	0.955 (0.033)	0.991 (0.041)
Memorandum items:		
J-statistic	5.707 (0.127)	2.590 (0.274)
Number of observations	208	208

Note: The data and sample periods covered are described in detail in the Appendix. Standard errors are shown in parentheses. For instrument set I the 0.95 critical value of the J-statistic, which is distributed as $\chi^2(3)$ under the null hypothesis, is 7.815. For instrument set II the relevant critical value is 5.991. The probability values of the J-statistic appear in parentheses.

reasonable proxies for the "true" utility-based indices.[29] Either assumption is likely to prove too restrictive in the case of developing countries, which are frequently subjected to terms of trade shifts and which commonly experience large movements in real exchange rates that alter the relative price between importables and home goods. The practice in previous literature of computing the real interest rate that is relevant for consumption decisions as the nominal rate divided by (one plus) the rate of change of a standard aggregate price deflator—that is, a deflator for which the correct utility-based weights have not been used—may thus potentially imply a serious misspecification, especially when the profile of relative prices (terms of trade and real exchange rates) is not constant through time.

For the most part, researchers in the past have used a linearized version of the Euler equations considered here for the particular case of a single consumption good:[30]

$$\Delta c_t = \alpha + \sigma r_t + e_t, \tag{15}$$

[29] Correct aggregation would apply utility-based weights to the various types of goods consumed. However, available aggregate price indices do not employ such a methodology.

[30] Notice that the assumption of linearity itself involves a number of additional restrictions (particularly on the joint distribution of consumption and rates of return), relative to the model estimated in this paper.

where c_t is (the natural logarithm of) aggregate consumption, r_t is the (conventionally measured) real rate of interest, and e_t is a random disturbance. The coefficient on the real interest rate, σ, is the intertemporal elasticity of substitution. Giovannini (1985) found no systematic relationship between changes in consumption and the real interest rate. Rossi (1988), who allowed for liquidity constraints, also failed to detect a relationship in many of the regions considered. Using the countries in our sample, we estimated the more restrictive version of the model given in equation (15). As in Giovannini (1985), the estimates obtained by applying instrumental variables techniques yielded no systematic relationship between consumption changes and the real interest rate. This, of course, highlights that our finding, summarized in Table 1, of a statistically significant intertemporal elasticity of substitution is not a product of the choice of countries or period covered in our sample. It rather suggests that, in estimating the parameters of consumer preferences, it is important to relax some of the assumptions underlying a specification such as (15). Specifically, in our case, it indicates the importance of disaggregating between traded and nontraded goods. A future line of research, particularly relevant for developing countries, would retain the multigood setting employed here, but would also relax the assumption of a perfect capital market and allow for the existence of liquidity constraints, as in Rossi (1988).

Table 2. *Estimates of the Model Using Panel Data for 13 Developing Countries: Instrument Set I*

Parameters	Africa	Asia	Latin America
ϵ	1.279	0.655	0.760
	(0.474)	(0.105)	(0.172)
σ	0.451	0.800	0.373
	(0.159)	(0.201)	(0.111)
β	0.945	0.995	0.995
Memorandum items:			
J-statistic	6.492	6.928	8.333
	(0.165)	(0.140)	(0.080)
SSR	2.857	7.451	1.234
Number of observations	62	81	65

Note: See note to Table 1. The value of β chosen is that which minimizes the sum of squared residuals (SSR); J is the value of the criterion quadratic function. The 0.95 critical value of the J-statistic, which is distributed as $\chi^2(4)$ under the null hypothesis, is 9.488. The probability values of the J-statistic appear in parentheses.

Thus far, we have imposed the restriction that preference parameters are identical across the three regions in our sample, a restriction that we feel is unlikely to be satisfied in practice. We now present a set of results that relax this assumption by allowing for possible regional variation in taste parameters. To offset the loss of degrees of freedom when the sample is broken up, we economize on the number of parameters to be estimated. Rather than estimate the parameter vector $\mu = [\beta, \epsilon, \sigma]$, as before, we confine our estimation instead to the parameters ϵ and σ. This is in keeping with our overall objective of shedding light on the HLM effect, since the parameter β will not play a critical role in this context (see Section I). Using the same estimation technique as before, we estimate ϵ and σ over a range of feasible values for β. Given the estimates of β presented in Table 1, the search was conducted over the range 0.900–0.995 at intervals of 0.005. The value of β presented in Tables 2 and 3 is that which minimized the sum of squared residuals (SSR). This search procedure not only allows us to pinpoint β for each region, but by imposing its value (as well as imposing the relevant value for a) in the estimation of ϵ and σ, it increases the efficiency of these estimates.

The results for instrument sets I and II are summarized in Tables 2 and 3, respectively. In general, the parameters are estimated with precision in all regions, and the overidentifying restrictions imposed by the model are not rejected by the data. However, interesting regional differ-

Table 3. *Estimates of the Model Using Panel Data for 13 Developing Countries: Instrument Set II*

Parameters	Africa	Asia	Latin America
ϵ	1.441	1.152	1.107
	(0.771)	(0.270)	(0.383)
σ	0.443	0.803	0.430
	(0.178)	(0.235)	(0.135)
β	0.940	0.990	0.995
Memorandum items:			
J-statistic	5.019	3.679	3.731
	(0.170)	(0.298)	(0.292)
SSR	3.506	1.661	1.658
Number of observations	62	81	65

Note: See note to Table 1. The value of β chosen is the one that minimizes the sum of squared residuals (SSR); J is the value of the criterion quadratic function. The 0.95 critical value of the J-statistic, which is distributed as $\chi^2(3)$ under the null hypothesis, is 7.815. The probability values of the J-statistic appear in parentheses.

ences in preferences emerge. Irrespective of the instrument set used, the intertemporal elasticity of substitution is estimated to be about 0.80 for Asia, and to be roughly half as large for Africa and Latin America.[31] In effect, estimates of σ do not appear to be very sensitive to the choice of instruments in any of the regions considered. The value of β that minimizes the sum of squared residuals is around 0.94 for Africa and around 0.995 for Asia and Latin America, indicating that future consumption is discounted more heavily in the African countries considered.

Tables 2 and 3 also reveal that importables and nontradables are closer substitutes in Africa than in Asia or Latin America. Part of these regional differences may be accounted for by regional differences in the commodity composition of tradables and nontradables. In particular, the share of durables in importables is lower in Africa than in other regions, and since nontradables are overwhelmingly nondurable (that is, services), this may account for the higher degree of substitutability in Africa. For instrument set I (Table 2), we find gross substitutability between importables and nontradables for the African countries only, whereas for the second instrument set (Table 3), gross substitutability is obtained in all three regions.

IV. Conclusions

The traditional explanation of the relationship between the terms of trade and the external current account balance has, for many years, rested on the Harberger-Laursen-Metzler hypothesis. According to this hypothesis, an improvement in the terms of trade raises a country's real income level, and since part of that increase in real income will be devoted to saving, the improvement in the terms of trade improves the current account.

This paper has presented a first attempt to obtain quantitative estimates of the main parameters that determine the response of private saving to transitory terms of trade shocks for a cross-section of developing countries in the context of a fully articulated intertemporal optimizing model. The main results of our study are as follows.

First, the estimated parameters that describe consumer behavior in a

[31] Rossi (1988) argued that estimates of the intertemporal elasticity of substitution are biased downward if liquidity constraints are not taken into account. The regional differences in estimates of σ may reflect this omission, since empirical evidence (see Haque and Montiel (1989)) indicates that the Asian countries in our sample are less liquidity constrained than their African counterparts. Unfortunately, the Haque-Montiel sample does not include any of the Latin American countries covered by this study.

simple three-good setting (the intertemporal and intratemporal elasticities of substitution, and the discount factor) are all economically meaningful, irrespective of the choice of instruments employed and/or the region considered. Disaggregation of the panel data allowed us to detect interesting regional differences. The overidentifying restrictions imposed by the model are not rejected by the data.

Second, the estimates of the intertemporal elasticity of substitution are significantly different from zero in all the regions. This finding contrasts with previous work, including Giovannini (1985) and, to a lesser degree, Rossi (1988) for the developing countries, and Hall (1988) for the United States. The implication of our finding is that, in response to shifts in real rates of interest, households in developing countries will generally alter the time profile of their consumption, increasing the growth rate of the latter in response to an expected increase in real rates of return.

Third, our estimates of the intratemporal elasticity of substitution suggest that substitution between tradables and nontradables is an important channel through which terms of trade shocks are transmitted to private saving and the current account. In particular, our results are consistent with the view that terms of trade shocks are likely to generate substantial fluctuations in real exchange rates, which in turn alter consumption rates of interest, thereby affecting saving behavior and the allocation of total expenditure between traded and nontraded goods.

Fourth, the estimates for all the regions considered cast doubt on the view that consumption smoothing is the only relevant factor governing the response of households to transitory terms of trade shocks. An important implication of our estimates is that transitory terms of trade shocks should give rise to intertemporal shifts in consumption both directly and through the movements in real exchange rates that they induce. Calibration of a dynamic stochastic equilibrium model using the econometric estimates of this paper (see, for example, Mendoza (1992)) should enable one to obtain reasonable quantitative estimates of the effects of transitory terms of trade shocks on private saving. Preliminary evidence in this regard suggests that, although private saving is likely to decline in response to transitory adverse terms of trade shocks, the magnitude of this decline is likely to be much smaller than would have been predicted on the basis of previous estimates of the intertemporal elasticity of substitution. A policy implication is that the need to "finance" transitory adverse movements in the terms of trade may be smaller than previously believed. Given the estimated parameter values, this conclusion is likely to be especially true for the Asian countries in our sample and less so for the Latin American and African countries.

Finally, while the paper has focused exclusively on the effects of transitory terms of trade shocks, the parameter estimates obtained here should prove useful in a variety of other contexts, including the assessment of the effects of trade reforms and fiscal policies.

APPENDIX

Description and Sources of Data

This Appendix provides a description of the data analyzed in Section III and lists the sources used.

Description of Data

Country	*Sample Period*	*Number of Observations*
Africa		74
Egypt	1968–87	20
Ghana	1968–83	16
Côte d'Ivoire	1968–85	18
Morocco	1968–87	20
Asia		96
Sri Lanka	1968–86	19
India	1968–85	18
Korea	1968–87	20
Pakistan	1968–87	20
Philippines	1968–86	19
Latin America		77
Brazil	1968–86	19
Colombia	1968–87	20
Costa Rica	1968–85	18
Mexico	1968–87	20

Series and Sources

International Financial Statistics (International Monetary Fund)
- Gross domestic product (GDP)
- Private consumption (national income accounts (NIA))
- Exports (NIA)
- Imports (NIA)
- Interest rate
- Exchange rate
- Population

World Economic Outlook (International Monetary Fund)
- Import unit values
- Export unit values

World Tables, 1988–89 (World Bank)
  GDP by sector of origin
  Deflator for services

Trade Data System (United Nations)
  Imports of consumer goods

Provided by the authors
  Domestic production of traded goods
  Domestic production of import substitutes
  Consumption of traded goods (importables)
  Consumption of nontraded goods

Description of Series Constructed by the Authors

Domestic production of traded goods (NIA basis)
= GDP originating in the agricultural, mining, and industrial sectors.

Domestic production of import substitutes (NIA basis)
= domestic production of traded goods (NIA basis) − exports (NIA basis).

Consumption of traded goods on an NIA basis (m)
= imports of consumer goods (converted to NIA basis) + domestic production of import substitutes (described above). This definition assumes exports are not domestically consumed and all domestic import substitutes produced are consumed. Import prices are used as the relevant deflator.

Consumption of nontraded goods on an NIA basis (n)
= personal consumption expenditures (NIA basis) − consumption of traded goods. The deflator for services is used as a proxy for the nontraded goods deflator.

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