Saving and Sociopolitical Instability in Developed and Less-Developed Nations*

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Circumstances under which the effect on saving of uncertainty due to sociopolitical instability is positive or negative are delineated: for specified economic threats, the sign of the relationship depends upon whether consumption relative-risk aversion is above or below a threat-dependent threshold. Cross-national empirical results are presented showing sociopolitical instability significant to saving, with sociopolitical instability affecting saving positively in developed nations and negatively in less developed nations. It is suggested that this cross-sectional result, and an analogous sequential result of Grossman and Shiller, reveal a positive relationship between development and preference for more economic progress.

I. Introduction

If economic development is a process that, while individual to each nation, nevertheless has sufficient statistical regularity for it to be sensible to speak of one or a small number of paradigms of development, then the cross-national study of developing and developed nations, and the course of development in a single nation, are capable of giving insights into one another. One should expect, for example, to see consistent relationships between preference shifts over the course of the

development of the United States and comparatively across nations. This is true in particular of risk-aversion. Shiller and others have argued that the volatility of U.S. stock market prices is too great to be explained by changes in dividends, because dividends do not vary enough over time.[1] Grossman and Shiller have shown that consumption variability in a consumption-based capital asset pricing model is explanatory of the volatility of stock prices if consumption risk-aversion is sufficiently large.[2] Their graphical presentation of actual versus "perfect foresight" prices suggests that in the United States, in the period 1889-1979, relative aversion to the risk of consumption uncertainty must have increased from an initial value of roughly four or less to a final value characterized by them as "implausibly high." If there is indeed a consistent relationship between the temporal and cross-national characteristics of preferences, the simplest form it might take is that relative aversion to the risk of consumption uncertainty increases dramatically as consideration moves from less to more developed nations.

One approach to the cross-national estimation of risk aversion is through examination of the connection between saving and sociopolitical instability. Across nations, sociopolitical instability (SPI), is an important, risk-aversion dependent determinant of saving. Stewart and Venieris (1985) and Venieris and Gupta (1986) have characterized the empirical relationship between SPI and saving as highly significant and negative, in a sample of sixty less-developed countries (LDC's). Additional data presented in this paper generalizes these results to the entire spectrum of development, demonstrating that the relationship between SPI and saving changes sign as one passes from LDC's to developed countries. We address the problem theoretically as well, assuming that SPI is perceived as a threat against future income and previously accumulated wealth, directly through loss of resources to partial expropriation or indirectly through physical and psychological inconevnience, with the possibility of agreement or disagreement between the signs of losses (negative gains) of wealth and income.[3] The cross-national change in the sign of the SPI dependence, we suggest, is

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[1] See Shiller (1981).

[2] See Grossman and Shiller (1981). We accept their facts at face value, but not Shiller's interpretation that investors are irrational, in the economic sense of being non-optimizing with respect to preferences and beliefs.

[3] The individual's perception may originate in statements made by initiators of sociopolitical instability, may be based upon information as to what happened elsewhere under similar circumstances, or may reflect a philosophy or tendency toward optimism or pessimism of some sociopolitical group. Regardless of their detailed nature and origin, all may be translated into perceived modifications of future resources.

plausibly related to the temporal shifts of consumption risk-aversion seen by Grossman and Shiller.

Section II sets forth the theoretical elements. Since the purpose of initiators of actions such as demonstration, riots, armed attacks, and the like can be presupposed to be either a change in government policies or a change in controlling regime, it is plausible that as sociopolitical instability increases, so does the objective and perceived probability that certain feared or hoped for economic consequences will be actualized. We construct a simple model in which, given the political and economic programs of two contesting factions at an initial date, the threat of change in the sociopolitical order at a final date is represented by two mutually exclusive future events, one in which the incumbent regime remains in power and the other in which there is a new regime. Abstracting away influences other than sociopolitical, income and saving are determined by the outcome of the conflict.[4] The perceived uncertainty of this change can be measured by the subjective probability that the incumbent government loses to its opposition, this probability being assumed an increasing function of SPI. This uncertainty is taken to be independent of the more usually considered uncertainties present in financial markets. We demonstrate under specified conditions that the relationship between SPI and saving changes from negative to positive as consideration moves from representative agents whose relative aversion to variations in consumption lies below a certain relative aversion to variations in consumption lies below a certain relative aversion threshold, to those whose relative aversion lies above.

Section III, IV and V present the empirical evidence. As already mentioned, our empirical findings support the hypothesis that the sign of the dependency of saving upon SPI changes from negative among LDC's to positive among developed countries. This and our related theoretical result are interpreted in Section VI. The observed sign change depends upon consumption risk aversion exceeding a consumption risk aversion threshold at some point in the course of development. Consumption risk aversion will indeed become large, plausibly very large, if the object of future preference shifts from wealth, the absolute ability to buy future consumption, to wealth with respect to an expectation proportional to present consumption, or the relative ability to buy future consumption. We suggest that this effect is present in greater degree in the more advanced stages of economic development. Section VII summarizes the work and suggests avenues for further research.

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[4] We have chosen to overlook some interesting strategic aspects of this problem, such as that of the opponents choosing threats opportunistically, rather than ideologically.

II. Theory

For simplicity we assume a basic two-period model with one first-period state and two second-period states, the *status mutatus,* Δ, in which there is a new sociopolitical order, and its logical complement, the *status quo,* Δ^c, in which the present regime continues in power. Both incumbent and challenger have advanced definite sociopolitical and economic programs, the one currently in application and the other to be carried out if and when the political opposition is empowered. We assume that these usually antithetical programs reflect the ideological orientations of the opponents, and that the parties are more likely to devote their energies to the resolution of the conflict between themselves than to continuing programmatic revisions, i.e., that these programs are likely to remain stable, so that they represent well-defined threats. Sociopoliticl instabilty is the condition in which different regimes are possible in successive periods, so that the probability of status mutatus is nonzero, and, conversely, sociopolitical stability is the condition in which status mutatus is impossible. For convenience in representing variables that are random with respect to the second-period states, we introduce two random indicator functions: 1_Δ, which assumes the value one in the status mutatus and zero otherwise, and 1_{Δ^c}, which assumes the value one in the status quo and zero otherwise.

Let w_1 denote initial wealth and y income in the second period, given the status quo; let c and s represent consumption and saving in the first period; and let R be the total second-period return to a dollar of first period saving. The portion of saving lost due to a change of regime is the random variable

$$\tilde{\theta} = \theta 1_\Delta = \begin{cases} \theta \text{ in status mutatus} \\ 0 \text{ in status quo} \end{cases},$$

where θ, referred to as the "threat to saving," represents the anticipated fraction of saving lost in the status mutatus (a net gain if $\theta < 0$). Analogously, the portion of second-period wages lost due to a change of regime is the random variable

$$\tilde{\mu} = \mu 1_\Delta = \begin{cases} \mu \text{ in status mutatus} \\ 0 \text{ in status quo} \end{cases},$$

where μ, referred to as the "threat to income or wages," represents the anticipated fraction of income lost in status mutatus (a net gain if $\mu < 0$).

Note well that by the term *saving* we mean the *total stock* of financial assets by means of which agents transfer claims to consumption across time.

With this interpretation in mind, it is apparent that second-period wealth is a random variable of the form

(1) $w(\tilde{\theta}, \tilde{\mu}) = (1 - \tilde{\theta})Rs + (1 - \tilde{\mu})y.$

With any positive total return, $R > 0$, even with a negative *rate of return*, $R < 1$, agents can get more wealth in both second-period states by saving than by borrowing as long as expropriation is strictly bounded by saving. If an agent has probability belief π that the sociopolitical order will change (belief $1 - \pi$ that the incumbent regime will continue in power), the agent's sociopolitical anticipations can be represented by a 3-vector $\Psi = [\pi, \theta, \mu]$. We exclude the complete expropriation, or worse, of saving or second-period income in the event of sociopolitical change, but not the possibility of gains over the status quo.

Given that the representative agent has initial wealth w_1, second-period income y, an opportunity to earn an anticipated total return of R on saving, an assessment of the environment Ψ, and a transitive, reflexive, and complete preference ordering, $\preccurlyeq$, of consumption-wealth pairs, (c, w), representable as a real-valued, state-independent, von Neumann-Morgenstern utility function, $u(c, w)$,

$$(c, w) \preccurlyeq (c', w') \Leftrightarrow u(c, w) \leqslant u(c', w'),$$

the agent's indirect valuation of saving in environment $[\pi, \theta, \mu]$ is given by the expected utility function

(2) $V(s | \pi, \theta, \mu) = E_\pi[u(w_1 - s, w(\tilde{\theta}, \tilde{\mu}))].$

We assume that μ is smooth, monotonic increasing in both arguments, and strictly concave, at least within a region of the domain of preferences sufficiently large to contain the consumption-saving problem. Thus marginal value is well-defined, consumers prefer more to less, and savers are risk averse.

The decision-maker's problem is to choose saving s subject to the constraints that first-period consumption and second-period wealth are nonnegative, so as to maximize the valuation function $V(s |, \pi, \theta, \mu)$. The nonnegativity constraint on second-period wealth is interpretable as a no-international-bankruptcy condition in the event of borrowing; or,

alternatively, the agent may be thought of as consuming all his remaining wealth in the second period, with the usual restriction to nonnegative consumption.[5] To ensure that the agent can get more second-period wealth by saving than by borrowing, we assume a strictly positive total return to saving, $R > 0$. With monotonic increasing, strictly concave preferences, and in the absence of redundant securities, the decision maker's problem has a unique solution

$$(3) \qquad s^*(\pi, \theta, \mu) = \arg \max V(s \mid \pi, \theta, \mu),$$

specified by the first-oder condition that the marginal value of saving equals zero,

$$(4) \qquad V_s(s \mid \pi, \theta, \mu)\big|_{s = s^*(\pi, \theta, \mu)} = 0.$$

(In the sequel we follow the convention used here, that subscripted functions denote partial derivatives, and asterisks denote the evaluation of functions at decision-optimal arguments.)

Consider the two-period consumption-saving problem under *certainty*. The agent attempts to maximize utility, $u(c, w)$, under budget constraint

$$p_1 c + p_2 w = Rc + w = Rw_1 + y,$$

with the price of second-period wealth, p_2, taken to be the numeraire, and the ratio of the price of consumption, p_1, to the price of second-period wealth the total return, R. The agent attains an allocation vector consistent with the budget and first-order conditions

$$(5) \qquad \begin{cases} u_1 = \lambda p_1 = \lambda R \\[2mm] u_2 = \lambda p_2 = \lambda \quad, \end{cases}$$

λ a Lagrange multiplier representing the shadow price of total resources. The second-order condition for the existence of a constrained maxi-

mum is the positivity of the bordered Hessian determinant, $H = 2Ru_{12} - u_{12} - R^2u_{22}$, which is assumed. It is also assumed that first-period consumption and second-period wealth are normal goods, which, from the Slutsky equation, means that $Ru_{22} - u_{12}$ and $u_{11} - Ru_{12}$ are negative.

Differentiating saving with respect to total return, and noting that saving is a function of quantities and prices in both parties, one obtains the saving response to changes in R,

$$(6) \qquad \frac{\partial s}{\partial R} = 2\frac{u_2}{H} + 2\frac{w \cdot (u_{22} - R^{-1}u_{12})}{H};$$

the first term of is the sum of a positive demand effect — increased saving in response to a higher price of consumption — and a positive substitution effect — increased saving in response to a lower price of second-period wealth; the second term of which is a negative income effect. Analogously, the income response is

$$(7) \qquad \frac{\partial s}{\partial y} = R\frac{u_{22} - R^{-1}u_{12}}{H},$$

which is negative. The response of saving to a change of any parameter that simultaneously affects both wages and the return to saving, $\eta = \eta(y,R)$, will be a linear composition of a demand-substitution effect with an income effect,

$$(8) \qquad \frac{\partial s}{\partial \eta} = \frac{\partial s}{\partial y}\frac{\partial y}{\partial \eta} + \frac{\partial s}{\partial R}\frac{\partial R}{\partial \eta} = \alpha\frac{u_2}{H} + \beta\frac{u_{22} - R^{-1}u_{12}}{H},$$

where the composition coefficients, α and β, are functions of wages and return to saving. Such a parameter will play an important role along certain paths defined in the sequel, on which the variation of μ and θ effectively determines y and R, and therefore the ratio of the behavioral parameters multiplying α and β,

$$(9) \qquad \frac{u_{22} - R^{-1}u_{12}}{u_2},$$

will be an important determinant of the effect of threats on saving. In the case in which the representative agent under consideration has simple temporally-additive utility,

$$(10) \qquad u(c,w) = g(c) + h(w),$$

the cross-derivative, u_{12}, vanishes, and this ratio, (9), is the agent's Pratt-Arrow risk aversion, $RA = -u_{22}/u_2$. In the sequel, under uncertainty, the appropriate analogue is the parameter

$$(11) \qquad \rho(\Delta) = -w(\Delta) \; \frac{u_{22}(\Delta) - \dfrac{u_{12}}{(1-\theta)R}}{u_2(\Delta)}$$

where Δ represents the (consumption, second-period wealth) vector in status mutatus. For want of a better term, and in the interest of not proliferating terminology, we refer to ρ as the consumption relative-risk aversion (CRRA).[6]

To incorporate SPI we associate the linear ordering of enviornments of increasing sociopolitical instability with that of increasing probabilities of sociopolitical change. Under this assumption, the probability of sociopolitical change π is a monotonic function of SPI, the relationship between saving and sociopolitical instability corresponding to the relationship between s* and π with threats θ and μ held constant. But since, as will be shown, any alteration of SPI under a constant threat is equivalent in its effect on saving to some alteration of threats under constant SPI, there is a dual problem in which one investigates the relationship between saving and threats, holding SPI, or equivalently π, constant. To admit the tools of calculus we need to consider infinitesimal changes accumulated sequentially along smooth paths, and the formulation of the problem in terms of increasing-threat paths turns out to be more tractable than the formulation in terms of increasing sociopolitical instability.

Paths are formed by permitting coordinates to vary as functions of a path parameter, considering arbitrary, pathwise differentiable mappings, $\Psi(t)$, from a closed interval τ into the space of admissible environments $\Psi = \{\psi = \Psi\}$. (We italicize parameters when we regard them as variable along the path, and not when we regard them as constant.) At every point of the space of admissible environments, for every agent representing a wealth-income group in a given nation, there is an optimal value of saving, so that an environmental path can also be characterized as an optimal saving path, $s*(t) \equiv s*(\psi(t))$.

Several classes of paths may be distinguished, corresponding to the signs of the rates of change of risk and saving along them. Let the probability of political transition have a monotonic dependence on the path

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6 The mixed derivative term is not crucial to our argument.

parameter, $t \in \tau$. An environmental path will be said to be (strictly) SPI-increasing if $\dot{\pi} \geqslant 0$ ($\dot{\pi} > 0$), and SPI-neutral if $\dot{\pi} = 0$, where $\dot{\pi} = \partial_t \pi$. The velocity of saving, $v(t) \equiv \partial_t s^*(t)$, along the path is determined by the incremental response of saving to increments in the environmental parameters over small parametric intervals, through the chain-rule relation

$$(12) \qquad v(t) = [\partial_\pi s^*, \partial_\theta s^*, \partial_\mu s^*] \begin{bmatrix} \dot{\pi} \\ \dot{\theta} \\ \dot{\mu} \end{bmatrix}.$$

Consider path segments over which saving varies monotonically. An environmental path will be said to be locally (strictly) saving-increasing if $v(t) \geqslant 0$ ($v(t) > 0$), and saving-neutral if $v(t) = 0$, in the environmental detertermined by t. Analogous terminology applies globally when saving is globally monotonic.

Two kinds of paths are of particular interest. First, paths that represent the smooth introduction of political risk into a stable environment in such a way that θ and μ remain fixed at the nonthreat value 0, while π increases to its terminal value. For these "initiation paths" let the parameter domain be $\tau = [-1, 0]$ and the path endpoints

$$\begin{cases} \psi(-1) = [\pi, \ \theta, \ \mu](-1) = [0, 0, 0] \\ \psi(0) = [\pi, \ \theta, \ \mu](0) = [\pi, 0, 0]. \end{cases}$$

Second, paths for which the probability of transition remains fixed at its terminal value π, while θ and μ increase or decrease monotonically to their terminal values. For these "completion paths" let the parameter domain be $\tau = [0, 1]$ and the path endpoints

$$\begin{cases} \psi(-1) = [\pi, \ \theta, \ \mu](-1) = [0, 0, 0] \\ \psi(+1) = [\pi, \ \theta, \ \mu](+1) = [\pi, \theta, \mu]. \end{cases}$$

Probabilities vary along initiation paths and threats along completion paths. Obviously, initiation paths are globally saving-neutral and completion paths (globally) SPI-neutral. The two can be pieced together to form compound paths. Saving, a function of a path's final endpoint, $\psi(1) = [\pi, \ \theta, \ \mu]$, is pathwise independent. Therefore only completion paths need be characterized as saving increasing or decreasing.

Path independence allows us to confine our attention to simple completion paths. At arbitrary points, the important characteristics of such paths are the ratio $(1 - \mu_t)/(1 - \theta_t)$, representing the conflict inherent in

the division of incentives to investment versus labor, and the ratio of threat increments, or derivative, $\partial \mu_t / \partial \theta_t$. For our purposes, the most useful parametric form is that in which the ratio of these parameters, or, equivalently, the threat-elasticity parameter

$$(13) \qquad \eta_t \equiv \frac{\partial \ln(1 - \mu_t)}{\partial \ln(1 - \theta_t)} = \frac{(1 - \theta_t)}{(1 - \mu_t)} \frac{\partial \mu_t}{\partial \theta_t},$$

remain globally constant along the path. Defining path-constants

$$(14a) \qquad \eta = \frac{\log(1 - \mu)}{\log(1 - \theta)} > 0$$

and

$$(14b) \qquad \zeta = \frac{1 - \mu}{1 - \theta} > 0,$$

and excluding the trivial case $\eta = 1$, it is easily verified that the parameterization

$$(15) \qquad \theta_t = 1 - \zeta^{\frac{t}{\eta - 1}}$$
$$\mu_t = 1 - \zeta^{\frac{\eta t}{\eta - 1}}$$

represents a path whose threats are zero at the initial point $t = 0$ and (θ, μ) at the final point $t = 1$, with threat-elasticity η_t everywhere equal to a constant η.

Consider a constant threat-elasticity path specified by η, ζ. Define the *saved-wealth to income ratio*,

$$(16) \qquad \sigma = \frac{Rs^*}{y}.$$

The following Lemma states conditions for saving along a completion subpath to be increasing or decreasing. These conditions are local, but the parametric subscripts are dropped for readability.

Lemma 1.   If agents ar net savers and consumption is a normal good, the velocity of saving is **negatively** related to the threat to saving, i.e., is negative if the threat to saving is positive and positive if the threat to saving is negative, **unless** the following condition is obeyed:

(17) $(\sigma + \eta^t)\, \rho(\Delta) > \sigma + \zeta^t,$

in which case the velocity is **positively** related to the threat. In the exceptional case $\dot\theta = 0$, when condition (17) is not well-defined, the sign of the velocity of saving is positively related to the threat to second-period income. ■

Proof. See Appendix A.

Assume that agents are risk averse in the sense $\rho > 0$. Then condition (17) can written

(18) $\rho^{-1} < \dfrac{\sigma + \eta\zeta^t}{\sigma + \zeta^t} = 1 + \dfrac{\eta - 1}{1 + \sigma\zeta^{-t}}.$

The right-hand side of inequality (18) can be negative, but ρ, by assumption, cannot. When the right-hand side is negative, it might as well be zero, for in neither case can the inequality be obeyed; and, therefore, we can substitute the following for (17):

(19) $\rho^{-1} < \max\ (0.1 + \dfrac{\eta - 1}{1 + \sigma\zeta^{-t}}).$

When the right-hand side of inequality (19) is zero, inequality (19) is equivalent to the condition $\rho > \infty$, which cannot be obeyed. Therefore, condition (17) can be written

(20) $\rho > R_t = [\max(0,\ 1\dfrac{\eta - 1}{1 + \sigma\zeta^{-t}})]^{-1},$

where 0^{-1} is taken the be ∞. We refer to the nonnegative, possibly infinite, path parameter R_t as the local *CRRA* threshold. We have proved the following theorem.

Theorem 1. If risk avese agents are net savers and consumption is a normal good, the velocity of saving is **negatively** related to the threat to saving **unless** CRRA exceeds the local CRRA threshold R_t,

(21) $\rho(\Delta_t) > R_t,$

in which case the velocity is **positively** related to the threat. In the exceptional case that there is no threat to saving, when condition (17) is not well-defined, the sign of the velocity of saving is positively related to the threat to second-period income. ■

On a constant threat-elasticity completion subpath, Theorem 1 establishes the relationship between threats and saving locally, while the saving response to sociopolitical instability can be established only by the global response to threats. The following trivial corollary gives sufficient global conditions for the sign of the velocity of saving.

Corollary 1. If risk averse agents are net savers and consumption is a normal good, saving is **negatively** related to the threat to saving if

$$(22a) \quad \max_t \; \rho(\Delta_t) < R_{min} = \min_t \; R_t,$$

and saving is **positively** related to the threat to saving if,

$$(22b) \quad \min_t \; \rho(\Delta_t) > R_{max} = \max_t \; R_t.$$

In the exceptional case that there is no threat to saving, when condition (17) is not well-defined, the sign of the velocity of saving is positively related to the threat to second-period income. ∎

Despite a region of ambiguity in which the sign of the velocity of saving is indeterminate, the essential point of Theorem 1 and its corollary is that a sufficiently high degree of consumption relative-risk aversion can reverse the marginal behavior of saving with respect to incremental threats. Since it is a matter of indifference whether one considers threats to saving and income or the sociopolitical instability that put these threats into effect, these results may be interpreted as stating that the response of saving to sociopolitical instability is negative if consumption relative-risk aversion is below some threshold, and positive if above. With any finite threshold, there is some value of *CRRA* beyond which the return available to saving inadequately compensates agents for the financial risk concomitant with political transition.

III. A Statistical Model Consistent with the Theory

In studying the effects of sociopolitical instability on saving, we saw that the stage of development plays a decisive role. Accordingly denoting the stage of development by $d \in \{\{DC's\}, \{LDC's\}\}$, and making explicit the income, initial wealth, and return-to-saving arguments, y, w_1, and R, the saving function, equation (3), becomes

$$(23) \quad s_d^* = \text{argmax} \; V_d(s|y, w_1, R, \pi, \theta, \mu).$$

We assume the probability of sociopolitical change (π) to be a

monotonic function of SPI, with an upper asymptote of one; i.e., $\pi = \phi(\text{SPI})$, $\phi'(\text{SPI}) > 0$, and $\lim_{SPI \to \infty} \phi(\text{SPI}) \to 1$. Substituting ϕ for π yields a new function, $\hat{V}_d(\text{SPI}) \equiv V_d(\pi)$.

If an agent's first-period wealth is, like second-period wealth, the sum of initial saved-wealth, w_0, and income, y_1, so that $w_1 = w_0 + y_1$, saving can be decomposed into saved-wealth plus incremental saving from first-period income, $s = w_0 + {}^\Delta s$. Thus,

$$(24) \qquad {}^\Delta s_d{}^* = f(y_2, w_0 + y_1, \text{SPI}, d, R, \zeta) - w_0$$
$$= {}^\Delta s_d{}^*(w_0, y_1, y_2, \text{SPI}, d, R, \zeta),$$

where, to be notationally consistent, second-period income is now written y_2 instead of y, and ζ is a list variable consisting of the (unobservable, subjectively estimated) threats to saving and income as well as other variables our model may have omitted.

At this juncture, it is useful to remind the reader that, given our definitions, ${}^\Delta s_d$ denotes the level of annual saving as commonly understood.

Since equation (24) makes saving depend on present and future income, the appropriate framework for empirical work is either the Friedman permanent income, or Modigliani-Brumberg life-style, model. Although their theoretical bases are similar, their empirical formulations differ. We have used Friedman's model because of data limitations. Moreover, since consumption data is invariably superior to saving data, we have opted to test the effect of SPI on consumption, with the understanding that it must be opposite in sign from the effect on saving.

Defining c_{it}^* and y_{it}^* ($i \in d$), to be the permanent consumption and income of the i-th country at date t, we assume $c_{it}^* = \psi(y_{it}^*)$, and, more explicitly, that ψ is a linear function, $\alpha_i + \beta_i y_{it}^*$. Allowing for a transitory component, u_{it}, measured consumption will be, $c_{it} = \alpha_i + \beta_i y_{it}^* + u_{it}$. Approximating permanent income by $y_{it}^* = (1 - \lambda_i) \sum_{0 \leqslant l < \infty} \lambda_i^l y_{i,t-1} (0 < \lambda_i^l < 1)$, and applying the usual transformations,

$$(25) \qquad c_{it} = \alpha_i^* + \beta_i^* y_{it} + \lambda_i c_{i,t-1} + \epsilon_{it},$$

where $\alpha_i^* \equiv (1 - \lambda_i)\alpha_i$, $\beta_i^* \equiv (1 - \lambda_i)\beta_i$, and $\epsilon_{it} \equiv u_{it} - \lambda_i u_{i,t-1}$.

If the model is well-specified, the u_{it}'s will not display much serial correlation, but the ϵ_{it}'s will. Also, it is likely that the u_{it}'s (and, hence, the ϵ_{it}'s) will be heteroskedastic. If the standard deviations of the u_{it}'s grow proportionally with income, a more efficient estimating equation

would be

$$(26) \qquad \frac{c_{it}}{y_{it}} = \beta_i^* + \alpha_i^* \frac{1}{y_{it}} + \lambda_i \frac{c_{i,t-1}}{y_{it}} + \eta_{it},$$

Equation (26) requires regression of the average propensity to consume, c_{it}/y_{it}, against inverse income, and lagged propensity to consume adjusted for the rate of income growth, i.e., $c_{i,t-1}/y_{it} \equiv (c_{i,t-1}/y_{i,t-1})(y_{i,t-1}/y_{it})$.

Regarding SPI, we assume its effect on average propensity to consume is manifested through a lagged structure,

$$(27) \qquad \frac{c_{it}}{y_{it}} = \beta_i^* + \alpha_i^* \frac{1}{y_{it}} + \Sigma_{0 \leqslant l < \infty} \alpha_{il} \mathrm{SPI}_{i,t-l} + \lambda_i \frac{c_{i,t-1}}{y_{it}} + \eta'_{it},$$

Tests of the lag shape over time showed it to be hyperbolic. We have imposed the hyperbolic form $a_{il} = b_{il}/l(l = 1,2,3,4)$, for simplicity. Substitution yields,

$$(28) \qquad \frac{c_{it}}{y_{it}} = \beta_i^* + \alpha_i^* \frac{1}{y_{it}} + a_{i0} \mathrm{SPI}_{it} + b_{i1} \Sigma_{1 \leqslant l < 4} \frac{\mathrm{SPI}_{i,t-l}}{l} + \lambda_i \frac{c_{i,t-1}}{y_{it}} \eta_{it}''.$$

Estimation of the return to saving requires rates of nominal return and inflation. For example, in temporary equilibria with one-period horizons,

$$r_{it} = R_{it} - 1 = \hat{r}_{it} - E_t(^\Delta p_{it}/p_{it}),$$

where $\hat{r}_{it}$ is the rate of nominal return, and $E_t(^\Delta p_{it}/p_{it}) = E_t p_{i,t+1}/p_{it} - 1$ the rate of inflation. However, nominal return data is not available. Since ignoring the $\hat{r}_{it}$'s altogether would result in a misspecification, we assume a constant interest rate absorbed into the constant term of the regression, and estimate only the inflationary term, $E_t(^\Delta p_{it}/p_{it})$. A commonly used specification of inflationary expectations is the adaptive form,[7]

$$E_t(\frac{^\Delta p_{it}}{p_{it}}) = E_{t-1}(\frac{^\Delta p_{i,t-1}}{p_{i,t-1}}) + v_i [\frac{^\Delta p_{i,t-1}}{p_{i,t-1}} - E_{t-1}(\frac{^\Delta p_{i,t-1}}{p_{i,t-1}})].$$

[7] Expected inflation may also be viewed as correlated with expected future real income and rates of a number of other assets (Stewart and Venieris, *op. cit.*). Recently, it has been argued in different theoretical frameworks, that saving is positively related to expected (Jump, 1980) and unexpected (Deaton, 1977) inflation. Our work should not be viewed as a test of either hypothesis.

Because of the large number of countries in our sample, we have not tried to distinguish adaptive parameters, but have assumed the same value, v, for each. Tests in the range $v \in \{0.0, 0.1,...,1.0\}$ typically showed v close to one, i.e., $E_t(\Delta p_{it}/p_{it}) = \Delta p_{i,t-1}/p_{it,t-1}$.[8] Introducing this term into (28),

$$(29) \quad \frac{c_{it}}{y_{it}} = \beta_i^* + \alpha_i^* \frac{1}{y_{it}} + a_{i0}SPI_{it} + b_{i1}Z_{it} + b_{i1}\frac{\Delta p_{i,t-1}}{p_{i,t-1}} + \lambda_i \frac{c_{i,t-1}}{y_{it}} + \eta_{it}''' ,$$

with the replacement $Z_{it} = \Sigma_{1 \leqslant l < 4}SPI_{i,t-l}/l$. This final version is the model to be tested.[9]

IV. The Data

In estimating equation (29), we have used a sample of 21 developed and 50 less-developed countries, spanning the period of 28 years from 1950 to 1978. The list of countries in each group is given in Appendix B. The sources of data on all economic variables are the recent publications of Summers and Heston (1984, 1988). In particular, c_{it} and y_{it} represent per-capita consumption and GNP, and c_{it}/y_{it} stands for the average propensity to consume of the i-th country at date t, while p_{it} denotes the implicit price deflator of GNP_{it}.

SPI is a multidimensional concept, and does not lend itself to direct measurement. Taylor and Hudson (1976) have recorded variables that reflect dimensions of social and political instability [Taylor and Hudson, (1976), Taylor and Jodice, (1983)]. These are: frequency of executive transfers, irregular executive transfers, government sanctions, riots, political protest demonstrations, political strikes, number of armed attacks, and deaths from domestic violence. From the above variables, the last six are usually chosen in the literature to represent the sociopolitical ambience of a country. Social scientists [e.g., Hibbs (1973)], through the use of factor analysis, have concluded, first, that these variables represent two different dimensions of instability: anomic (riots, protest demonstrations, and political strikes) and violent (assassinations, armed attacks, and deaths from domestic violence); and, second, that the former can be represented by political demonstrations, the latter by deaths.

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[8] Blinder reports similar results in characterizing the effects of income distributions on aggregate consumption (Blinder, 1975).

[9] The reader may wonder why we have not divided all explanatory variables of (29) by y_{it}. One may argue either way; we have therefore tested both versions.

We have used these last two variables in the framework of cluster analysis, to construct an index for SPI. In the process, we also determined that the inclusion of the variable "assassinations" results in a major improvement in the classification of observations. We used these three variables for "clustering" each year of each country, and found that all country-year observations can be classified optimally into six clusters. We used a multinominal logit model to determine the conditional probabilities that each country-year takes the value one to six, and to estimate the contributions of each variable (i.e., demonstrations, deaths, assassinations) to the conditional probabilities in questions. Finally, SPI indices were constructed by taking expected values of conditional probabilities so determined. Appendix C provides explicit details of method and estimated equations.

V. Statistical Analysis

In estimating equation (29), we pooled all DC's and LDC's separately, and used the dummy variable model,[10]

$$(30) \qquad \frac{c_{it}}{y_{it}} = \Sigma_j \beta_{ij} D_{jt} + \Sigma_k \beta_k x_{kit} + e_{it},$$

[10] Before we use equation (29) we have to test two hypothesis. First, whether linear approximation of (24) is satisfactory, and, second, whether use of the dummy variable model is legitimate.

In regard to the first equation, we have estimated the model,

$$\frac{c_{it}}{y_{it}} = x\beta + \epsilon = x_1\beta_1 + x_2\beta_2 + \epsilon$$

where β_1 is the vector of coefficients of equations 1 of Tables 1 and 2, and β_2 is the vector of coefficients of second order terms in Taylor's expansion of equation (24), whose terms are cross-products and squares of the explanatory variables in equation 1. The hypothesis to be tested is $H_0: \beta_2 = 0$. For this, we use the likelihood ratio, $\lambda_{LR} = 2[L(\beta) - L(\beta_1)] \sim \chi^2(k_2)$, where k_2 is the number of restrictions. We have found,

$$L(\beta) = 7.297 \qquad \bar{R}^2_\beta = .928$$
$$L(\beta_1) = 7.292 \qquad \bar{R}^2_{\beta 1} = .923$$

for DC's, and,

$$L(\beta) = 7.790 \qquad \bar{R}^2_\beta = .965$$
$$L(\beta_1) = 7.786 \qquad \bar{R}^2_{\beta 1} = .961$$

for LDC's. The corresponding λ_{LR}'s are .005 and .004, much smaller than the critical value of $\chi^2_{.05}(6) = 12.59$. We therefore accept H_0, concluding that the linear parameterization is a satisfactory approximation.

The second question entails the tests,

$$H_0: \beta_{1i} = \beta_{1j} \qquad i \neq j \; \epsilon(1,...,21) \text{ for DC's}$$
$$H_1: \beta_{1i} \neq \beta_{1j} \qquad i \neq j \; \epsilon(22,...,71) \text{ for LDC's}$$

For this, we compare equations 1 of Tables 1 and 2 with corresponding versions restricting

with $1 \leqslant i$, $j \leqslant 21$ for DC's, $22 \leqslant i$, $j \leqslant 71$ for LDC's, $1 \leqslant t \leqslant 28$, and $D_{jt} = 1(j=i)$.[11] Tables 1 and 2 provide a bird's eye view of various versions we estimated, for DC's and LDC's repsectively.

To conserve space, we do not present the estimates of the short-run marginal propensities to consume for each equation in these tables. In the case of DC's, these are all highly significant, with t-values ranging from 14.1 to 15.6. In the case of LDC's, the majority of the estimates are significant with t-values ranging from 1.22 to 18.53. In all cases the critical value of the t-statistics at 5% significance is 1.645.

All equations in both tables have high explanatory power, and reveal no serious autocorrelation (column of rho's, $\hat{\rho}$); neither do they seem subject to heteroskedasticity. Indeed, application of the White test for heteroskedasticity [White (1980)] to equations 1 of Tables 1 and 2, for example, suggests that the null-hypothesis of homoskedasticity should be accepted.[12] Moreover, in terms of explanatory power, autocorrelation, and standard errors (column of SE's), it seems to make no difference whether explanatory variables are divided by y_{it}, or not. For our purposes, the main characteristics of these tables are, first, that SPI enters significantly into the majority of equations, and, second, that it changes sign between DC's (Table 1) and LDC's (Table 2).[13]

Our preliminary empirical findings are therefore consistent with our

their constants to a single value. We use the F-statistics,

$$F = \frac{(\tilde{e}'\tilde{e} - \hat{e}'\hat{e})/(N-1)}{\hat{e}'\hat{e}/(NT-T-K)},$$

where $\tilde{e}'\tilde{e}$ and $\hat{e}'\hat{e}$ denote the residual sums of squares from the restricted and unrestricted models respectively, and $(N-1)$ and $(NT-T-K)$ the corresponding numbers of degrees of freedom. We find, $F^{DC} = 13.68$, and $F^{LDC} = 12.31$. The corresponding values at 5% significance are $F_{.05}^{DC}(20,509) = 1.57$, and $F_{.05}^{LDC}(49,1164) \cong F_{.05}^{LDC}(49,\infty) = 1.35$. We therefore reject the null hypothesis, concluding that the specification to be used is the unrestricted model, equation (30).

[11] The indicator function, $1(A)$, equals one if A is true, zero if A is false.

[12] To test for heteroskedasticity, we form the White equations based on equations 1 of Tables 1 and 2. The unadjusted $\bar{R}^2$'s of these equations for DC's and LDC's are .352 and .965, respectively. The corresponding degrees of freedom are 504 and 1159. Denoting the degrees of freedom by T_{DC} and T_{LDC}, we find that $T_{DC} \cdot \bar{R}_{DC}^2 = 177.41$ and $T_{LDC} \cdot \bar{R}_{LDC}^2 = 1118.44$. The corresponding $\chi_{.05}^2$ values are 590 and 1289. We therefore accept the null hypothesis that our errors are homoskedastic. The formula used to approximate the $\chi_{.05}^2$ with the above degrees of freedom is,

$$\chi_{.05}^2 = \frac{1}{2}(1.6449 = \sqrt{2T_d - 1})^2, \quad d\epsilon\{\{DC's\},\{LDC's\}\}.$$

On this, see Pearson and Hartley (1966).

[13] Furthermore, the coefficients of SPI are larger in the Table 2.

Table 1

ESTIMATES OF AVERAGE PROPENSITY TO CONSUME: SAMPLE DC'S

Model: $\dfrac{c_{it}}{y_{it}} = \sum_j \beta_{ij} D_{jt} + \sum_k \beta_k x_{kit} + e_{it}$　$D_{jt} = \begin{cases} 1 \text{ if } j=i \\ 0 \text{ if } j \neq i \end{cases}$　$\begin{cases} i,j \in (1,\cdots,21) \\ t \in (1,\cdots,28) \end{cases}$

Eq'n #	$\dfrac{1}{y_{it}}$	SPI_{it}	z_{it}	$\dfrac{\Delta p_{i,t-1}}{p_{i,t-1}}$	$\dfrac{SPI_{it}}{y_{it}}$	$\dfrac{z_{it}}{y_{it}}$	$\dfrac{\Delta p_{i,t-1}}{y_{it}p_{i,t-1}}$	$\dfrac{c_{i,t-1}}{y_{it}}$	$\bar{R}^2$	$\hat{\rho}$	SE	# of OBS.
					Explanatory Variables							
1	60.94 (5.64)	−9.07 E−4 (1.87)						0.48 (17.95)	.86	.110	1.83 E−2	533
2	63.40 (6.01)	−8.48 E−4 (1.83)		−3.44 E−2 (2.43)				0.51 (18.86)	.88	.0822	1.81 E−2	524
3	43.50 (2.32)				−5.19 (1.79)			0.48 (17.80)	.87	.114	1.82 E−2	533
4	43.90 (2.34)				−5.04 (1.81)		−9.24 (2.12)	0.49 (18.91)	.91	.108	1.80 E−2	533
5	76.33 (5.37)	−5.24 E−4 (1.81)	−1.08 E−4 (1.91)	−5.88 E−2 (3.84)				0.53 (18.52)	.92	.0371	1.69 E−2	445
6	77.26 (2.17)				−3.10 (1.83)	−2.74 (1.91)	−15.71 (2.06)	0.52 (16.66)	.92	.0378	1.68 E−2	445

|t|-statistics in parentheses.

Table 2

ESTIMATES OF AVERAGE PROPENSITY TO CONSUME: SAMPLE LDC'S

Model: $\dfrac{c_{it}}{y_{it}} = \sum_j \beta_{ij} D_{jt} + \sum_k \beta_k x_{kit} + e_{it}$ $D_{jt} = \begin{cases} 1 \text{ if } j=i \\ 0 \text{ if } j \neq i \end{cases}$ $\begin{cases} i,j \in (22,\cdots,71) \\ t \in (1,\cdots,28) \end{cases}$

Eq'n #	Explanatory Variables								$\bar{R}^2$	$\hat{\rho}$	SE	# of OBS.
	$\dfrac{1}{y_{it}}$	SPI_{it}	z_{it}	$\dfrac{\Delta p_{i,t-1}}{p_{i,t-1}}$	$\dfrac{SPI_{it}}{y_{it}}$	$\dfrac{z_{it}}{y_{it}}$	$\dfrac{\Delta p_{i,t-1}}{y_{it}\,p_{i,t-1}}$	$\dfrac{c_{i,t-1}}{y_{it}}$				
1	30.90 (5.59)	6.09 E-3 (2.46)						0.42 (21.59)	.87	211	3.34 E-2	1217
2	30.25 (5.47)	4.24 E-3 (2.51)		1.30 E-2 (1.82)				0.43 (21.70)	.90	210	3.04 E-2	1217
3	20.97 (2.42)				2.64 (1.77)			0.42 (21.36)	.88	213	3.35 E-2	1217
4	19.78 (2.28)				2.84 (1.68)		13.09 (1.82)	0.42 (21.57)	.92	209	3.09 E-2	1217
5	27.72 (4.49)	4.20 E-3 (2.51)	1.99 E-2 (1.54)	2.15 E-2 (2.69)				0.47 (23.02)	.94	204	2.98 E-2	1094
6	21.93 (2.15)				3.34 (1.89)	0.87 (1.69)	17.93 (2.32)	0.46 (21.48)	.94	205	3.00 E-2	1094

$|t|$ - statistics in parentheses.

theoretical conclusion that the effect of SPI on saving is negative in LDC's and positive in DC's.

At this point, it is interesting to determine the extent of contribution of SPI. For this, we employ equation (30), but excluding SPI from the set of explanatory variables. More specifically, we use modified versions of equations 1 of Tables 1 and 2. The results are:

$$\text{DC's:} \quad \frac{c_{it}}{y_{it}} = 59.91 \frac{1}{y_{it}} + .51 \frac{c_{i,t-1}}{y_{it}} \qquad \begin{array}{l} \bar{R}^2 = .764 \\ \hat{\rho} = .132 \end{array}$$

$$\qquad\qquad\quad (5.01) \qquad (19.47) \qquad\qquad \text{SE} = .0218$$

$$\text{LDC's:} \quad \frac{c_{it}}{y_{it}} = 30.21 \frac{1}{y_{it}} + .46 \frac{c_{i,t-1}}{y_{it}} \qquad \begin{array}{l} \bar{R}^2 = .781 \\ \hat{\rho} = .273 \end{array}$$

$$\qquad\qquad\quad (4.97) \qquad (20.49) \qquad\qquad \text{SE} = .0438$$

Comparing the above with equations 1 of Tables 1 and 2, we see that inclusion of SPI among the arguments of the latter increases $\bar{R}^2$ by 12 and 11 percent, respectively. Autocorrelation coefficients and standard errors are decreased as well. Finally, regression parameters are uniformly smaller, meaning that exclusion of SPI biases short-run marginal propensities to consume, downward.

We characterized the above results as preliminary because equation (30) restricts all coefficients (other than the constants), forcing them to assume the same value across all countries of the respective samples. Clearly, this assumption can be tested. We regressed the unrestircted model

$$(31) \qquad \frac{c_{it}}{y_{it}} = \Sigma_j \, \beta_{ij} \, D_{jt} + \Sigma_k \, \beta_{k,j} \, D_{it} \, x_{kit} + e_{it},$$

with $1 \leqslant i$, $j \leqslant 21$ for DC's, $22 \leqslant i$, $j \leqslant 71$ for LDC's, $1 \leqslant t \leqslant 28$, and $D_{jt} = 1 (j = i)$. Equation (31) is equivalent to one regression per country of our sample. Estimating (31) using the same explanatory variables as equations 1 of Tables 1 and 2, and comparing them with the F-test as before (see footnote 11), we found $F^{DC} = 1.87$ and $F^{LDC} = 2.05$. These values are greater than the critical values at 5% significance,

$$F^{DC}_{.05}(62,536) \cong F^{LDC}_{.05}(62,1155) \cong F^{X}_{.05}(62,\infty) \cong 1.32.$$

We therefore reject the null hypothesis, that all coefficients (other than the constant) are the same across each sample, and accept the alternative, that the coefficients of at least one explanatory variable are not

so restricted.

To determine which should be allowed to vary, we estimated the coefficients of variation ($CV \equiv s/\bar{x}$, where s denotes the standard deviation, and $\bar{x}$ the mean) of the set of coefficients of all variables in equation (31), i.e., 21 for DC's, and 50 for LDC's, and unrestricted the coefficient with the highest CV in equations 1 of Tables 1 and 2. This criterion is both intuitively appealing and simple to use. Table 3 provides the pertinent statistics associated with each set of coefficients.

From Table 3 we see that the largest CV is that of the coefficients of $1/y_{it}$. We removed restrictions on the coefficients of this variable by estimating.

$$(32) \quad \frac{c_{it}}{y_{it}} = \Sigma_j \ \beta_{1j} \ D_{jt} + \Sigma_j \ \beta_{2j} \ \frac{D_{jt}}{y_{it}} + \Sigma_k \ \beta_k \ x_{kit} + e_{it} \begin{cases} \text{DC's: i,j}\epsilon(1,...,21) \\ \text{LC's: i,j}\epsilon(22,...,71) \end{cases}$$

with $t \ \epsilon \ (1,...,28)$. Estimates of this model are presented in Tables 4 and 5.

From Tables 4 and 5 we see that both the qualitative and quantitative effects of the various explanatory variables, on the average propensity to consume, are in line with those reported in Tables 1 and 2. Adjusted correlations have increased, while standard errors have declined, substantially. Unrestricting the coefficients of $1/y_{it}$ also resulted in a uniform reduction of the values of the adjustment coefficients, particularly in LDC's, and a corresponding increase in the constants (short-run marginal propensities to consume). Additionally, a number of previously significant constants, in LDC's, are no longer significant, although the majority do retain their significance. The |t|-values of the constants in the DC's have not been affected.

The |t|-values of the coefficients of $1/y_{it}$ themselves, have generally declined in DC's, but remain significant; whereas in LDC's the number of insignificant coefficients now exceeds that of significant ones.

The effects of unrestricting the coefficients of $1/y_{it}$ have a silver lining. Now the null hypothesis, that the rest of the coefficients can remain restricted has to be accepted. Indeed, we now find that, $F^{DC} = 1.163$, and $F^{LDC} = 1.254$. The corresponding critical value of the F-statistics at 5% significance is $F^{DC}_{.05}(102,536) \doteq F^{LDC}_{.05}(102,1155) \doteq F^{*}_{.05}(102,\infty) \doteq 1.26$.

The empirical results presented in Tables 1, 2, 4, and 5 support our theoretical conclusions, and corroborate the findings of earlier works, in relation to LDC's [Stewart and Venieris, *op.cit.*, Venieris and Gupta (1986), Venieris and Stewart (1987)].

Table 3
Means, Standard Deviations, and Coefficients of Variation of Coefficients of Explanatory Variables

Sample	$1/y_{it}$			SPI_{it}			$c_{i,t-1}/y_{it}$		
	$\bar{x}$	s	CV	$\bar{x}$	s	CV	$\bar{x}$	s	CV
DC's	17.80	1.86 E2	10.45	-7.97 E-2	.00673	-8.48	.50	.218	.44
LDC's	24.87	2.37 E2	9.53	2.08 E2	.142	6.83	.36	.325	.90

Notation: $\bar{x} \equiv$ mean of the coefficient

$s \equiv$ standard deviation

$CV \equiv s/\bar{x}$

Table 4

ESTIMATES OF AVERAGE PROPENSITY TO CONSUME: SAMPLE DC'S

Model: $\dfrac{c_{it}}{y_{it}} = \sum\limits_{j=1}^{21} \beta_{1j} D_{jt} + \sum\limits_{j=1}^{21} \beta_{2j} \dfrac{D_{jt}}{y_{it}} + \sum\limits_{k} \beta_k x_{kit} + e_{it}$ $\quad \begin{cases} i,j \in (1,\cdots,21) \\ t \in (1,\cdots,28) \end{cases}$

Eq'n #	SPI_{it}	$SPI_{i,t-1}$	z_{it}	$\dfrac{\Delta p_{i,t-1}}{p_{i,t-1}}$	$\dfrac{SPI_{it}}{y_{it}}$	$\dfrac{SPI_{i,t-1}}{y_{it}}$	$\dfrac{z_{it}}{y_{it}}$	$\dfrac{\Delta p_{i,t-1}}{y_{it}\,p_{i,t-1}}$	$\dfrac{c_{i,t-1}}{y_{it}}$	$\bar{R}^2$	$\hat{\rho}$	SE	# of OBS.
						Explanatory Variables							
1	−6.75 E−4 (1.97)								0.36 (13.64)	.90	.0809	1.62 E−2	533
2	−5.32 E−4 (1.78)	−2.76 E−4 (1.69)							0.38 (12.11)	.91	.0706	1.60 E−2	512
3	−5.42 E−4 (1.92)			−2.87 E−4 (2.57)					0.41 (13.71)	.94	.0403	1.37 E−2	533
4					−6.11 (1.81)				0.36 (13.60)	.90	.0902	1.61 E−2	533
5					−4.98 (1.79)	−3.57 (1.54)			0.37 (12.57)	.90	.0803	1.62 E−2	512
6					−6.04 (1.80)			−8.04 (2.11)	0.39 (12.35)	.94	.0501	1.43 E−2	533

Table 4 (Continued)

7	-3.03 E-4 (1.81)	-1.21 E-4 (1.78)					0.40 (12.63)	.92	.0603	1.59 E-2	445
8	-4.01 E-4 (1.77)	-1.19 E-4 (1.54)	-3.19 E-4 (2.32)				0.42 (12.56)	.94	.0371	1.26 E-2	445
9				-3.17 (1.83)	-1.78 (1.62)	-12.56 (2.17)	0.41 (12.31)	.94	.0389	1.32 E-2	445
10		-7.01 E-4 (1.58)	-3.25 (2.16)				0.43 (12.76)	.89	.0912	1.71 E-2	445

$|t|$ - statistics in parentheses.

Table 5

ESTIMATES OF AVERAGE PROPENSITY TO CONSUME: SAMPLE DC'S

Model: $\dfrac{c_{it}}{y_{it}} = \sum_{j=22}^{71} \beta_{1j} D_{ji} + \sum_{j=22}^{71} \beta_{2j} \dfrac{D_{ji}}{y_{it}} + \sum_k \beta_k x_{kit} + e_{it}$ $\quad \begin{cases} i,j \in (22,\cdots,71) \\ t \in (1,\cdots,28) \end{cases}$

Eq'n #	Explanatory Variables									$\bar{R}^2$	$\hat{\rho}$	SE	# of OBS.
	SPI_{it}	$SPI_{i,t-1}$	z_{it}	$\dfrac{{}^{\Delta}p_{i,t-1}}{p_{i,t-1}}$	$\dfrac{SPI_{it}}{y_{it}}$	$\dfrac{SPI_{i,t-1}}{y_{it}}$	$\dfrac{z_{it}}{y_{it}}$	$\dfrac{{}^{\Delta}p_{i,t-1}}{y_{it}p_{i,t-1}}$	$\dfrac{c_{i,t-1}}{y_{it}}$				
1	6.97 E-3 (2.01)								0.24 (11.38)	.91	.267	2.96 E-2	1217
2	4.14 E-3 (2.46)	1.01 E-3 (1.78)							0.25 (12.01)	.92	.261	2.78 E-2	1167
3	4.21 E-3 (1.98)			2.01 E-2 (2.11)					0.25 (11.97)	.94	.259	2.51 E-2	1217
4					3.91 (1.87)				0.24 (12.11)	.90	.261	2.95 E-2	1217
5					2.44 (1.79)	1.76 (1.64)			0.24 (11.81)	.91	.261	2.82 E-2	1167
6					2.76 (1.83)			15.01 (1.96)	0.25 (11.89)	.93	.260	2.67 E-2	1217

Table 5 (Continued)

7	4.85 E-3 (2.17)	1.58 E-3 (1.86)					0.24 (11.53)	.93	.258	2.76 E-2	1094
8	4.38 E-3 (1.99)	1.69 E-3 (1.62)	2.23 E-2 (2.01)				0.24 (12.38)	.97	.231	1.86 E-2	1094
9				3.73 (1.77)	1.11 (1.68)	19.31 (2.04)	0.23 (11.54)	.97	.229	1.89 E-2	1094
10		1.85 E-3 (1.69)	2.11 (2.65)				0.24 (11.56)	.87	.262	2.86 E-2	1094

$|t|$ - statistics in parentheses.

Table 6 presents some pertinent descriptive statistics of the variables, for both DC's and LDC's. From Table 6, we see that the mean estimated value of SPI in LDC's is approximately 60% higher than in DC's. This might not appear to be a significant increase, but looking at it's effect on average propensity to consume, it becomes another story. To appreciate the magnitude of the effects, consider, for instance, equation 1 of Table 5. At the mean-value of SPI, the effect on c/y is .0244($= 6.97E - 3 \times 3.22$). For example, Tait, Gratz, and Eichengreen (1979, Table 15), report average tax revenue in LDC's of approximately 18% of GNP. Given the average propensities to consume from Table 6, it follows that APS $= 1 - $ APC $- $ ATR $= .17$.[14] Therefore, SPI, evaluated at its mean and one standard deviation, results in a reduction of the saving rate by 13% and 22% respectively, a substantial foregone growth opportunity.

The reader might think the increase of average propensity to save in DC's due to its decrease in LDC's, because of capital flight. While transfer of capital across international borders may be related to SPI, it is erroneous to extend the same reasoning to a cross-national comparison of international relationships between saving and SPI. The observed movements of average propensities to save in DC's and LDC's are independent of one another. Indeed, this is one reason why we chose average propensity to consume as our dependent variable in the first place.

VI. Interpretation

The important functional dependencies expected on the basis of our two-period model are the following: [1] An increase in second-period income, y_2, will reduce saving, by the normality of second period wealth and equation (7). Since initial saved wealth, w_0, is fixed exogenously, this is equivalent to a reduction in incremental saving, $\triangle s^*$. [2] An increase in first-period income, y_1, being equivalent to an increase in first-period wealth, w_1,

$$(33) \qquad \frac{\partial s}{\partial y_1} = \frac{\partial s}{\partial w_1} = \frac{\partial(w_1 - c)}{\partial y_1} = 1 + R^2\left(\frac{u_{22} - R^{-1}u_{12}}{H}\right)$$

which, by the normality of first-period consumption and second-period

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[14] APS $=$ average propensity to save, APC $=$ average propensity to consume, ATR $=$ average tax rate.

Table 6

DESCRIPTIVE STATISTICS OF THE SAMPLES

Variables	Parameters							
	Mean		Min		Max		Standard Deviation	
	DC's	LDC's	DC's	LDC's	DC's	LDC's	DC's	LDC's
c/y	.61	.65	.50	.71	.84	.99	.0618	.174
y	5.80 E3	2.54 E3	1.15 E3	2.38 E2	1.08 E4	*4.74 E4	1.97 E3	4.46 E3
SPI	2.01	3.22	1.005	1.005	5.98	6.00	1.26	2.16
$\Delta p/p_{t-1}$	.0207	.00513	-.49	-.95	.25	1.36	.0589	.12

*The high values of per capita income are due to Kuwait and Saudi Arabia..

wealth, is positive but less than one. Thus incremental saving is a normal good with respect to first-period income. [3] The relationship between saving and SPI, therefore the relationship between incremental saving and SPI, is negative for CRRA below the CRRA threshold, and positive above. Our empirical evidence is incapable of illuminating the first point, but is in agreement with the second. With regard to the third, it indicates that the difference between CRRA and its threshold is a function of development.

Let us first consider the CRRA threshold and then CRRA itself. Environments in which the CRRA threshold is less than one are somewhat special. When there are positive threats to both income and saving, they demand that the threat to saving be smaller than that to income. That the CRRA threshold is less than one when the threat elasticity, η, is greater than one, is a straightforward deduction from condition (20). It is equivalent to $\mu > \theta > 0$ when a loss of saving is anticipated in status mutatus; when a gain of saving is anticipated in status mutatus, it is equivalent to $\mu < \theta < 0$. In other words, the CRRA threshold is less than one when both threats have the same sign and the magnitude of the threat to income is greater than the magnitude of the threat to saving. The CRRA threshold is infinite when

(34) $\qquad \eta \leqslant -\sigma\zeta^{-t} \leqslant 0,$

which under the exclusion of the complete expropriation of saving implies that the threat to saving is opposite in sign to the threat to income, a situation we refer to as "explicit conflict" to distinguish it from the case of "implicit conflict," in which redistribution occurs with threats to saving and income of the same sign. Not all explicit conflict yields an infinite threshold, of course, because the threat elasticity must be sufficiently negative.[15]

These dependencies are illustrated in figures 1 through 5, which depict the relationship of the CRRA threshold R, to the threat to saving θ, for various values of the threat to income μ and (constant) saved-wealth to income ratio σ. In each figure, the left-hand side, a, characterizes the beginning of the completion path and the right-hand side, b, the end. (R_t behaves monotonically in between.) Abscissas are linear, representing loss threats to saving from two to ninety-eight percent. Ordinates are logarithmic and values of R are truncated to the maximum ordinate, 100. Each figure, except figures 4b and 5b, contains

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[15] Pareto optimality need not apply to the sociopolitical situation, in which externalities and non-traded assets are heavily at stake.

graphs for values of σ equal to .25, .50, 1.00, 2.00, and 4.00. In figure 4b, the graph for $\sigma = .25$ exceeds the maximum ordinate at every point, and is therefore omitted; in figure 5b, the graphs for $\sigma = .25$ and $\sigma = .50$ exceed the maximum ordinate at every point, and are omitted. The progression is sufficiently regular in σ that only extremes, e.g., .25 and 4.00, are labeled.

Threat of loss to saving appears to be the case relevant to most of our sample, not because a change of regime would necessarily not be favorable to capital formation, but because the sociopolitically unstable route to a change of regime will generally involve losses to be subtracted from any gains. The horizon for "sociopolitical planning" tends to be further away than the horizon for economic planning. Nevertheless, the

Figures
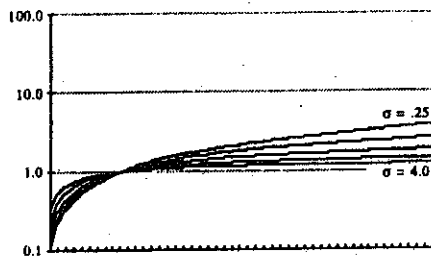
Figure 1a. Implicit conflict. Income loss 20% at $t = 0$.
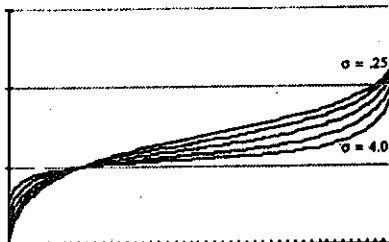
Figure 1b. Implicit conflict. Income loss 20% at $t = 1$.
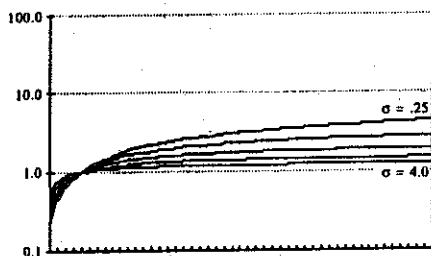
Figure 2a. Implicit conflict. Income loss 10% at $t = 0$.
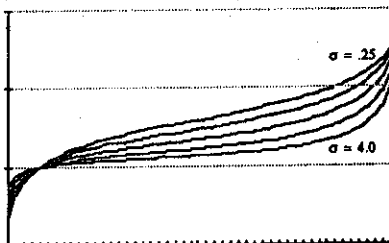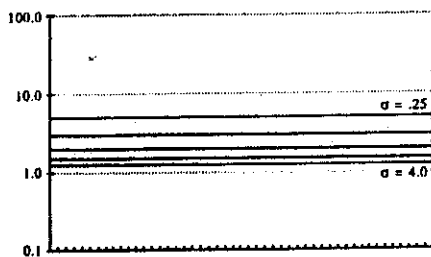
Figure 2b. Implicit conflict. Income loss 10% at $t = 1$.

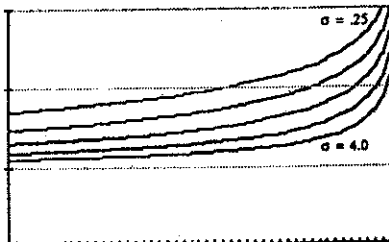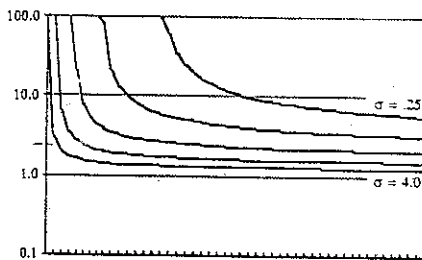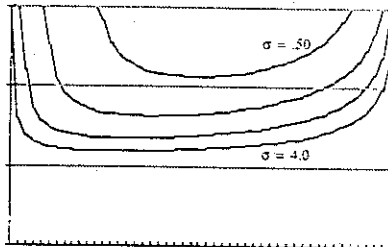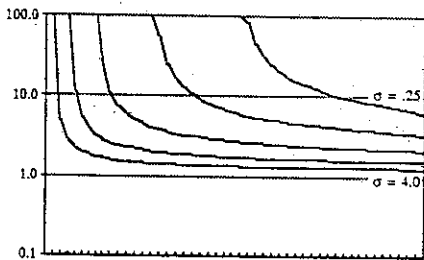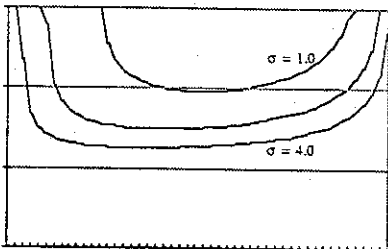Figure 3a. No loss or gain of income at $t = 0$.

Figure 3b. No loss or gain of income at $t = 1$.

Figure 4a. Explicit conflict. Income gain 10% at $t = 0$.

Figure 4b. Explicit conflict. Income gain 10% at $t = 1$.

Figure 5a. Explicit conflict. Income gain 20% at $t = 0$.

Figure 5b. Explicit conflict. Income gain 20% at $t = 1$.

corresponding graphs for the case in which there are gains to saving are qualitatively similar, with the following adjustments: the cases of implicit conflict, figures 1 and 2, become the cases of explicit conflict, and conversely the cases of explicit conflict, figures 4 and 5, become the cases of implicit conflict; and the large CRRA thresholds seen for large loss threats to saving in figures 4b and 5b do not occur, because a ninety-eight percent loss of saving is a reduction of saving by a factor of fifty, while a ninety-eight percent gain is an increase of less than a factor of two.

As for CRRA itself, it has been noted, originally by Arrow (1971), that if one assumes boundedness of the utility function, relative-risk aversion must be less than logarithmic as the quantity of the object of preference approaches zero and more than logarithmic as it approaches infinity. In the present case, this is consistent with the empirical evidence. But, since the two-period problem involves multiple objects of preference, i.e., first-period consumption and second-period wealth, the concept of risk aversion appropriate to it is subject to the definitional ambiguity common to multiple-good problems. Any such definition suffers from ambiguity in at least two respects: First, there is no standard measure of the second-period good, and therefore no standard measure of its variance when multiplicatively disturbed. Second, for similar reasons there is no standard measure of the premium to be paid for insuring away the risk. To clarify this, we give an example.

For agents having simple temporally additive utility as expressed by equation (10), aversion to an infinitesimal multiplicative mean-preserving random deviation from a certain quantity of second-period wealth, w, is characterized by the Pratt-Arrow coefficient

$$(35) \qquad RRA = -w\frac{u_{22}(c,w)}{u_2(c,w)} = -w\frac{h''(w)}{h'(w)}.$$

For example, if h is logarithmic,

$$(36) \qquad h(w) = \ln w, \ h'(w) = w^{-1}, \ h''(w) = -w^{-2},$$

then $RRA = 1$. But suppose the second-period object of preference is wealth with respect to a baseline reference proportional to first-period consumption, $x = w - \gamma c$, where γ is a nonnegative constant. (Assume, for simplicity, that x is positive, $w > \gamma c$.) The utility function becomes

$$(37) \qquad u(c,x) = g(c) + h(x) = g(c) + \ln x,$$

and the measures of aversion to an infinitesimal multiplicative mean-preserving random deviation from a certain quantity of the second-period good should be analogy bé

$$(38) \qquad RRA = -x\frac{u_{22}(c,x)}{u_2(c,x)} = -x\frac{h''(x)}{h'(x)}$$

which again equals one. If one attempts to use the original definition, the first part of equations (35), one must use the chain rule relations,

$$(39) \qquad u_w(c,x) = \frac{\partial x}{\partial w} u_2(c,x) = u_2(c,x)$$

$$u_{ww}(c,x) = \frac{\partial x}{\partial w} \frac{\partial}{\partial x} [\frac{\partial x}{\partial w} u_2(c,x)] = u_{22}(c,x),$$

from which relative-risk aversion, rather than equaling one, has the anomalously large value

$$(40) \qquad RRA_w = \frac{w}{w - \gamma c}.$$

It is worth noting that if we drop the assumption of temporal additivity, the cross-derivative u_{cw} becomes

$$(41) \quad u_{cw}(c,x) = \frac{\partial u_w(c,x)}{\partial c} = \frac{\partial u_2(c,x)}{\partial c} = u_{12}(c,x) + \frac{\partial x}{\partial c} u_{22}(c,x)$$

$$= -\gamma u_{22}(c,x) = \gamma x^{-2},$$

so that in the absence of any threat to saving CRRA assumes the value

$$(42) \quad \rho = -w \frac{u_{ww} - R^{-1} u_{cw}}{u_w} - w \frac{(1 + \gamma R^{-1}) u_{22}}{u_2}$$

$$= (1 + \gamma R^{-1}) RRA_w = (1 + \gamma R^{-1}) \left(\frac{w}{w - \gamma c}\right)$$

If one thinks of γ and R as having values roughly equal to one, the presence of the cross-derivative term in the definition of ρ inflates its value by roughly a factor of two, with respect to the value of RRA_w. But, since it may be the case that second-period wealth does not greatly exceed the baseline reference γc, RRA_w can exceed $RRA_x = 1$ by an order of magnitude or more. Thus, CRRA may plausibly attain very large values, even when utility is temporally additive.

Now if the difference between CRRA and its threshold is a function of development, it may be the case that both CRRA and the CRRA threshold are functions of development, or that the effect predominantly involves one term or the other. If CRRA remains fairly constant over the range of development, to obtain agreement with the empirical evidence the CRRA threshold must fall. Indeed, as Figures 1 through 5 illustrate, when there is a positive threat to saving the CRRA threshold is inversely dependent upon σ. We would expect development to be correlated with increased capital and a higher saved-wealth to income ratio, exactly as required. However, the dependence is relatively slow when there is a threat to saving and conflict is implicit, at least compared with the potential for variation of the baseline reference CRRA. It appears more probable, particularly if CRRA is less than one at very low consumption levels, that both effects are present, with the latter dominant. Sociopolitical conditions will of course differ across nations at any given time, and there will be a random distribution of economic threats associated with these different sociopolitical environments. Thus, there will be a random distribution of CRRA thresholds. If CRRA has a value near one in LDC's, it is unlikely that an arbitrary LDC will demonstrate a positive relationship between SPI and saving. If CRRA has a value well in excess of one in developed countries, it is unlikely that a DC will demonstrate a negative relationship between SPI and saving.

What is the empirical status of this conjecture? There are, to our knowledge, no other empirical studies characterizing relative-risk aversion across nations as a function of development, but Grossman and Shiller's well-known study characterizes it across time. Using price and dividend data in time-series from 1890 to 1979, they compared the volatility of U.S. stock market prices with "perfect foresight" stock prices computed by valuing discounted dividend payout streams under specified values of consumption relative-risk aversion in an additive utility model. Their result, a graph comparing the actual Standard and Poor's price average with a perfect-foresight price average computed on the basis of relative-risk aversion equal to a constant value of four over their time interval, shows about the right amount of volatility from 1889 to the early 1930's, somewhat too little volatility (given CRRA) from then to the early 1950's, and strikingly too little since. Since the volatility of perfect foresight prices increases with relative-risk aversion, this means that for the United States the value of relative-risk aversion has been not only anomalously large but, more importantly, increasing over time. Mehra and Prescott (1985) studied a superset of Grossman and Shiller's data over the same time interval using a more analytical method, and obtained the result that relative-risk aversion, if constant over time, was in excess of ten. Thus we may conclude that Grossman and Shiller saw relative-risk aversion rising from a value substantially below ten to a value substantially above.[16]

Measured values of relative-risk aversion, even when confined to recent years in the United States, are widely dispersed. Various studies support the notion that individuals are risk-lovers, at least over certain domains [Kahneman and Tversky (1979), Schoemaker (1980)], risk-averse but more risk-loving than logarithmic [Hansen and Singleton (1982,83)], logarithmic [Brown and Gibbons (1985)], slightly more risk-averse than logarithmic [Hansen and Singleton (1983), Friend and Blume (1975)], and anomalously more risk-averse than logarithmic [Grossman and Shiller (1981), Mehra and Prescott (1985)]. It must be remembered, however, that here we are dealing with representative individuals in a market setting, so that studies performed with a single-subject psychological methodology, such as those of Kahneman and Tversky, and Schoemaker, may be inappropriate. Furthermore, relative-risk aversion to variations in consumption must be distinguished from relative-risk aversion to variations in wealth,

$$(43) \qquad -w_1 . V''(w_1) \big/ V'(w_1), \text{ where } V(w_1) = V(s^* | \pi, \theta, \mu) |_{w_1}.$$

16 We reiterate that we are taking these studies at face value, without Shiller's interpretation that investors are irrational.

Of the studies cited, only those of Grossman and Shiller, and Mehra and Prescott are of this CRRA kind.

VII. Summary and Comments on Further Research

We have investigated the effects on saving of uncertainty due to sociopolitical instability. Our theoretical treatment delineates the circumstances under which the effect of SPI on saving will be negative or positive, demonstrating that for given threats to saving and income, the relationship between SPI and saving will change from positive to negative as a function of whether or not consumption relative-risk aversion is above or below a threat-dependent threshold. Our empirical results show that sociopolitical instability, a significant factor in understanding saving across nations, affects saving negatively in the case of less developed countries and positively in the case of developed ones. Taken together, these two approaches to the one problem reveal that the difference between consumption relative-risk aversion and its threshold is a function of development.

Consistency between theory and empirical results demands anomalously large values of relative-risk aversion for developed nations. Agents can be highly rish averse to variations in consumption if their object of preference is second-period wealth relative to a baseline proportional to present consumption. The most important aspect of this effect for understanding development is its apparently dynamical behavior. Consumption relative-risk aversion will increase, even to very large values, if the baseline increases to a level close to the level of sustainable future consumption, a plausible scenario when a national economy is perceived as having reached its "limits of growth." Our empirical results do not tell us the values of relative-risk aversion in the countries studied, but anomalously high values of consumption relative-risk aversion have been observed in studies of stock markets in the United States.

From a slightly different perspective, our work adds to a body of literature pointing toward the necessity of enlarging the set of variables traditionally considered the concern of economists. While saving and income were investigate here in relation to SPI, the sociopolitical environment, we believe, is capable of reasonably objective characterization beyond the single SPI variable. Political scientists have developed methods of content analysis for quantifying news reports, e.g., for predicting the outcomes of national elections [Budge and Farlie (1983), for example], and it is conceivable that these could be applied to the

estimation of threats and threat-relevant parameters such as shocks to
the probability of economic change. Since the sign of incremental saving
under marginal shocks to SPI is a function of the disparity between
CRRA and its threshold, which in turn is a function of the threat en-
vironment, these methodologies could reduce a significant source of
randomness in the estimation of CRRA and its dependence upon
development. More accurate data could also be a spur to theory, hope-
fully demanding refinement in the representation of the relationship be-
tween development and preferences, and ultimately between socio-
political and economic change.

Appendix A

Proof of Lemma 1

Proof. Differentiating equation (2) along an arbitrary saving-
optimal path segment, using the chain rule, and dividing through by
$\partial_{ss}V^*$, a strictly negative variable under the strict concavity of utility
assumption, gives the normalized "value-velocity" condition,

$$(A1) \quad \frac{\partial_{ts}V^*(s^*)}{\partial_{ss}V^*(s^*)} = 0 = [\partial_\pi s^* + \frac{\partial_{\pi s}V^*}{\partial_{ss}V^*}]\dot{\pi} + [\partial_\theta s^* + \frac{\partial_{\theta s}V^*}{\partial_{ss}V^*}]\dot{\theta} + [\partial_\mu s^* + \frac{\partial_{\mu s}V^*}{\partial_{ss}V^*}]\dot{\mu}.$$

Since this must hold for an arbitrary path segment, i.e., for arbitrary
time derivatives of π, θ, and μ, it must be the case that

$$\partial_\pi s^* = -\frac{\partial_{\pi s}V^*}{\partial_{ss}V^*}$$

$$(A2) \quad \partial_\theta s^* = -\frac{\partial_{\theta s}V^*}{\partial_{ss}V^*}$$

$$\partial_\mu s^* = -\frac{\partial_{\mu s}V^*}{\partial_{ss}V^*}$$

everywhere within the region under consideration. This region was itself
arbitrary. Therefore condition (A2) is generally applicable within the
space of admissible environments.

Consider a continuation path at an arbitrary point $[\pi, \theta, \mu] \equiv \psi(t)$.
Since the parameterized environment determines saving velocity through
the chain-rule relation

$$(12) \qquad v(t) = [\partial_\pi s^*, \ \partial_\theta s^*, \ \partial_\mu s^*] \begin{bmatrix} \dot{\pi} \\ \dot{\theta} \\ \dot{\mu} \end{bmatrix}$$

equations (A2) and the negativity of $\partial_{ss}V^*$, equation (12) imply

$$(A3) \qquad \text{sign } v(t) = \text{sign}([\partial_{\pi s}V^*, \ \partial_{\theta s}V^*, \ \partial_{\mu s}V^*] \begin{bmatrix} \dot{\pi} \\ \dot{\theta} \\ \dot{\mu} \end{bmatrix})$$

The right-hand side of equation (A3) has the form of a dot-product between a coefficient vector determined by preferences and beliefs and a path velocity vector. The explicit form of the coefficient vector may be obtained by twice differentiating saving velocity at $s = s^*(\psi)$. As before letting the state symbol Δ represent the consumption-wealth pair in the status mutatus, $[c,(1-\theta)Rs^* + (1-\mu)y]$, and now also Δ^c in the status quo, $[c, Rs^* + y]$, we obtain

$$\partial_{\pi s}V^* = R[(1-\theta)u_2(\Delta) - u_2(\Delta^c)] - [u_1(\Delta) - u_1(\Delta^c)]$$

$$(A4a,b,c) \quad \partial_{\theta s}V^* = -\pi R u_2(\Delta) - \pi R s^*[(1-\theta)Ru_{22}(\Delta) - u_{12}(\Delta)]$$

$$\partial_{\mu s}V^* = -\pi y[(1-\theta)Ru_{22}(\Delta) - u_{12}(\Delta)]$$

as the coefficients.

In general by the normality of consumption, $\partial_{\mu s}V^*$ is positive, but $\partial_{\theta s}V^*$ can be positive or negative. However, $\partial_{\theta s}V^*$ is always larger than $\partial_{\mu s}V^*$ times Rs^*/y. Recalling the definition of the saved-wealth to income ratio,

$$(16) \qquad \sigma = \frac{Rs^*}{y},$$

and defining the marginal utility parameter,

$$(A5) \qquad Q = \pi R u_2(\Delta).$$

We may relate $\partial_{\theta s}V^*$ to $\partial_{\mu s}V^*$ by

$$(A6) \qquad \partial_{\theta s}V^* = \sigma \partial_{\mu s}V^* - Q.$$

This is useful because it is the difference between positive factors, and

permits the sign of the velocity of saving to be expressed as

(A7) $\text{sign } v(t) = \text{sign}(\partial_{\mu s}V^*(\mu + \sigma\dot{\theta}) - Q\dot{\theta}).$

Since $\partial_{\mu s}V^*$ and Q are both positive, which effect predominates depends upon their ratio, which, using equation (A4c) and the definition of the CRRA parameter ρ, equation (11), can be written

(A8) $\dfrac{\partial_{\mu s}V^*}{Q} = -\dfrac{y[(1-\theta)Ru_{22}(\Delta) - u_{12}(\Delta)]}{Ru_2(\Delta)} = (1-\theta)\dfrac{y}{w^*}\rho(\Delta),$

or, using equations (1), (14b), (15), and (16) to put wealth in the form

$$w^* = (1-\theta)Rs^* + (1-\mu)y$$
$$= [(1-\theta)\sigma + (1-\mu)]y$$
$$= (1-\theta)(\sigma + \zeta')y$$

and defining the (strictly positive) parameter Γ,

(A9) $\Gamma = \sigma + \zeta',$

more simply as

(A10) $\dfrac{\partial_{\mu s}V^*}{Q} = \dfrac{\rho(\Delta)}{\Gamma}.$

We may complete the argument relating the velocity of saving to relative-risk aversion on the basis of equation (A10).

Let us distinguish the cases $\dot{\theta} = 0$ and $\dot{\theta} \neq 0$. According to equation (A7), if $\dot{\theta} = 0$, the positively of $\partial_{\mu s}V^*$ implies that saving is a positive function of the threat to second-period income,

(A11) $\text{sign } v(t) = \text{sign}(\dot{\mu}\partial_{\mu s}V^*) = \text{sign } \dot{\mu}.$

If $\dot{\theta} \neq 0$, we can factor the argument of the right-hand side of equation (A7) into $Q\dot{\theta}$ and the argument divided through by $Q\dot{\theta}$, to obtain the velocity of saving in the form

(A12) $\text{sign } v(t) = \text{sign}[(Q\dot{\theta})(\dfrac{\partial_{\mu s}V^*}{Q\dot{\theta}}(\dot{\mu} + \sigma\dot{\theta}) - 1)]$

$$= \text{sign } (Q\dot{\theta}) \text{ sign } [\frac{\partial_{\mu s} V^*}{Q}(\sigma + \frac{\partial \mu}{\partial \theta}) - 1],$$

or, using equation (A10) and the (strict) positively of Γ and Q, in the form

(A13) $\text{sign } v(t) = -\text{sign}\dot{\theta} \text{ sign}[\Gamma - (\sigma + \frac{\partial \mu}{\partial \theta})\wp(\Delta)].$

From equations (13) and (15) it follows that

$$\frac{\partial \mu}{\partial \theta} = \eta \zeta^t,$$

so that equation (A13) can be written as

(A14) $\text{sign } v(t) = -\text{sign } \dot{\theta} \text{ sign } [\Gamma - (\sigma + \eta \zeta^t)\wp(\Delta)].$

The second sign function on the right-hand side of (A14) is positive unless inequality (17) is obeyed, in which case it is negative. Using the definition of Γ, equation (A9), Lemma 1 follows immediately. ■

Appendix B

Countries in Sample

DC's: United States, Canada, United Kingdom, Netherlands, Belgium, Lexembourg, France, Switzerland, Spain, Portugal, West Germany, Austria, Italy, Finland, Sweden, Norway, Denmark, Iceland, Japan, Australia, and New Zealand.

LDC's: Haiti, Dominican Republic, Jamaica, Trinidad and Tobago, Barbados, Mexico, Guatemala, Honduras, El Salvador, Nicaragua, Costa Rica, Panama, Columbia, Venezuela, Guinea, Ecuador, Peru, Brazil, Bolivia, Paraguay, Chile, Argentina, Uruguay, Ireland, Malta, Greece, Cypress, Iran, Turkey, Iraq, Syria, Jordan, Israel, Saudi Arabia, Kuwait, Afghanistan, Taiwan, Hong Kong, South Korea, India, Pakistan, Burma, Sri Lanka, Nepal, Thailand, Malaysia, Singapore, Philippines, Indonesia, Papua New Guinea.

Appendix C

Construction of SPI Index

We used cluster analysis, a method of grouping like data into non-overlapping clusters by a Euclidean metric criterion, to synthesize information embedded in "event" variables into a single index of sociopolitical instability. The computation proceeds in two phases. In the first, "nearest component sorting," cluster components are computed, and each variable is assigned to the component with which it has the highest squared correlation. In the second, a search procedure tests whether classifying variables into different clusters improves the explained variance. The criterion for terminating this procedure associates to each cluster a single eigenvalue, greater than one.

Initially, following the literature [e.g., Hibbs, *op. cit.*], we used protest demonstrations and the log of deaths (plus one) due to domestic violence, to characterize the sociopolitical environment. Experimentation indicated, however, that inclusion of assassinations increased the "total explained variation" substantially, from .7934 to .9898. This gave an optimum number of clusters equal to five, with an additional grouping of unclustered observations corresponding to the most sociopolitically stable environments. These groups were assigned indices 1,...,6, starting from the most sociopolitically stable country-year observations (1), and ending with the most unstable.

By this procedure, we classified 1988 observations, for 71 countries over 28 years, as follows:

Group	Observations
1	833
2	408
3	89
4	55
5	59
6	544

As the reader can see, DC's and LDC's were analyzed together. One may object that the characterization of sociopolitical instability by our variables, should be different in the two. This objection makes a great deal of sense. We therefore tried to cluster observations of DC's and LDC's separately; but no significant difference with respect to c/y was

obtained.

This classification scheme does not allow for differentiation among observations belonging to the same cluster. Moreover, it yields a discontinuous variable for SPI. To allow for such differentiation, and to produce a continuous SPI index, we used a multinominal logit model,

$$(C1) \quad \Pr(Y = i) = \frac{e^{s_i}}{\Sigma_{j=1}^{6} e^{s_j}}$$

where Y denotes the cluster containing a given country-year observation, to predict the probability that the observation, characterized by certain variables (e.g., number of political protest demonstrations), belongs to a given cluster. The s_i's are calculated according to the model,

$$(C2) \quad \begin{aligned} s_1 &= X_1\beta_{11} + X_2\beta_{12} + X_3\beta_{13} \\ &\vdots \\ s_6 &= X_1\beta_{61} + X_2\beta_{62} + X_3\beta_{63}, \end{aligned}$$

where the X's are the regressors, and the β's the coefficients to be estimated. Equations (C2) apply to single observations. The input data consists of the value of the dependent variable (1,...,6) and the X_i, $i = 1,2,3$.

The multinominal logit model provides us with several sets of coefficients, so that a separate s_i can be computed for each potential value of the dependent variables, six in our case. Since one of the s_i can be normalized to zero, implying that one of the sets of coefficients will also be zero, only five sets of coefficients need be estimated. In PROC MLOGIT [Steinberg (1987)], it is the score of the highest value of the dependent variable, corresponding to cluster 6, that is normalized to zero.

The estimates of (C2) are:

$$s_1 = \begin{array}{cccc} -10.24 & +8.92 \text{ PPD} & +4.45 \text{ Assas} & +17.25 \log(D+1) \\ (16.74) & (10.23) & (3.38) & (12.94), \end{array}$$

$$s_2 = \begin{array}{cccc} -9.51 & +9.40 \text{PPD} & +4.15 \text{ Assas} & +14.95 \log(D+1) \\ (16.07) & (10.80) & (3.12) & (11.32) \end{array}$$

$$s_3 = \begin{array}{cccc} -9.31 & +9.43 \text{ PPD} & +3.93 \text{ Assas} & +14.63 \log(D+1) \\ (15.09) & (10.83) & (2.91) & (11.09) \end{array}$$

$$s_4 = \begin{array}{cccc} -10.18 & +9.28 \text{ PPD} & +4.13 \text{ Assas} & +16.03 \log(D+1) \\ (16.71) & (10.65) & (3.12) & (12.07) \end{array}$$

$$s_5 = \begin{matrix} -5.38 \\ (10.74) \end{matrix} \quad \begin{matrix} +9.52 \text{ PPD} \\ (10.94) \end{matrix} \quad \begin{matrix} +4.67 \text{ Assas} \\ (3.56) \end{matrix} \quad \begin{matrix} +10.65 \log(D+1) \\ (8.34), \end{matrix}$$

$|t|$-statistics in parentheses; log likelihood -610.34,

where PPD, Assas, and $\log(D+1)$ denote political protest demonstrations, assassinations, and the log of deaths (plus one) due to domestic violence.[17]

Given the predicted value of s_j, the probability of its cluster score can be derived from equation (C1). The expected value of SPI is therefore determined, as $\text{SPI}_{kt} = \Sigma_{i\epsilon[0,6]} i \cdot \text{Pr}(i)$. This value was subsequently used in estimating the c/y equations.

[17] A large number of country-year observations of D are zero.

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