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by

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Two of the principal implications of the permanent income hypothesis (PIH) as originally developed by Friedman (1957) were that the "short run" marginal propensity to consume (MPC) should be less than the "long run" marginal propensity to consume and that the short run MPC should be lower for households or groups for whom a large fraction of the variation of their income is transitory than for households or groups experiencing more permanent variations in income. Consumption was assumed proportional to permanent income, proxied as a distributed lag of past incomes, with arbitrary restrictions imposed on the lag weights in order to identify the long run MPC. The coefficient of current income was assumed less than unity, constraining the short run MPC to be less than the long run MPC.

With the advent of rational expectations theory in macroeconomics, in particular Lucas' (1976) seminal "Critique," came the recognition that distributed lags derived from expectations should depend on the parameters of the stochastic processes governing the variables about which expectations are being formed.[1] One implication that emerges from the rational expectations version of the permanent income hypothesis (RE-PIH) as developed by Hall (1978) is that permanent income, and hence consumption, should (like asset prices) possess the martingale property.[2] A second implication of the RE-PIH is that the contemporaneous effect on consumption (β) of an innovation to current income should be positively related to the revision in permanent income (V) implied by the innovation. This implication is similar in spirit to Friedman's.

However, under the RE-PIH, the measure of the "permanence" or "persistence" of an income innovation, V, is derived specifically from the parameters of the stochastic process governing income. In this paper we use annual post-war data from thirty countries to test these two implications of the RE-PIH.

Hall's now well-known test of the first implication requires that lags of income convey no information about current consumption when consumption lagged once is included among the explanatory variables. We adapt a version of this test to our post-war data for thirty countries and find results generally favorable to the RE-PIH. Flavin (1981) describes the theoretical relation between the effect of an income innovation on consumption and the stochastic process governing income, but does not test the validity of the relation within the single country she investigates (the US), arguing that the relation between β and V is exact only under extremely restrictive assumptions. Since one would not expect the β - V relation to hold exactly in any single regime (country), we construct "cross-regime" tests which are designed to determine whether a positive relation between β and V holds across our panel of thirty countries. Our results reveal the positive relation predicted by the RE-PIH.

Our results also relate directly to the sensitivity of consumption to income innovations. Flavin, using detrended US data, argues that consumption is excessively sensitive to income innovations. More recently, Mankiw and Shapiro [1985] have argued that if income contains a unit root, findings of excess sensitivity derived from detrended data may be spurious. Deaton [1986] and Campbell and Deaton [1987] showed that, for the US, when one allows for a unit root in the income process, consumption in fact appears under-sensitive. The results in this paper produce similar findings over the post-war data of thirty countries. In the notation of this paper, we find that β is significantly less from V for all thirty countries in our panel. Given the increasing preponderance of evidence in favor of a unit root in GNP--by Nelson and Plosser (1982), Campbell and Mankiw (1987), Schwart (1987) and others for the US and by Campbell and Mankiw (1987) and Kormendi and Meguire (1987) for more than thirty countries spanning both century long and postwar sample periods--it appears likely that consumption is not excessively sensitive to income innovations. We argue further that when one accounts for a less than unit long run propensity to consume and imperfect information, consumption appears generally appropriately sensitive to income. When combined with the strong positive relation across countries between β and V predicted by the RE-PIH, the evidence in this paper is quite favorable to the RE-PIH.

I. TESTABLE IMPLICATIONS OF THE RE-PIH

1. Hall-Flavin Tests

The basic model, following Flavin's formulation, is described by:

(1) $y_t = rw_{t-1} + x_t$: definition of current income

(2) $y_t^p \equiv rw_{t-1} + (1 = \alpha)\Sigma a^j E_t x_{t+j}$: definition of permanent income

(3) $w_t = (1 + r)w_{t-1} + x_t - c_t$: budget constraint

(4) $c_t = ky_t^p$: consumption function

where $y_t \equiv$ current income

$y_t^p \equiv$ permanent income

$r \equiv$ interest rate (assumed constant)

$\alpha \equiv 1/(1 + r)$

$w_t \equiv$ end of period non-human wealth

$x_t \equiv$ labor income (assumed exogenous)

$c_t \equiv$ consumption

$k \equiv$ constant marginal propensity to consume out of permanent income

Combining the definitions (1) and (2) implies that

$$y_t = y_t^p + \omega_t$$

where $\omega_t = (1 - \alpha)\ \Sigma\alpha^j(x_t - E_t x_{t+j})$. Note the ω_t, the difference between current and permanent income, is not necessarily white noise or even mean zero if, for example, the labor income process is nonstationary.

Combining (2) and (3) produces

$$(5) \quad y_t^p = (1 + r)y_{t-1}^p - rc_{t-1} + \theta_t \ ,$$

where $\theta_t = y_t^p - E_{t-1}y^p{}_t = (1 - \alpha)\Sigma\alpha^j(E_t - E_{t-1})\ x_{t+j}$ is defined as the innovation to permanent income. The above description of the permanent income process holds for any consumption function and any labor income process. When consumption behavior is given by (4), however, we can substitute (4) into (5) to obtain

$$(6) \quad y_t^p = (1 + r(1-k))y_{t-1}^p + \theta_t \ ,$$

or alternatively

$$(7) \quad c_t = (1 + r(1-k))c_{t-1} + k\theta_t \ .$$

Equation (7) is the consumption process tested by Hall. Setting $k = 1$ in (7) yields the consumption process tested by Flavin. Under rational expectations, θ_t should be orthogonal to all information dated $t - 1$ and earlier. A finding that the coefficient on a variable added to (7) and dated $t - 1$ or earlier is significantly different from zero would then be intrepreted as rejecting the RE-PIH as embodied in the model (1) - (4). Hall's empirical results rejected the null hypothesis marginally, while Flavin obtained strong rejections.

We now consider some reasons why rejections of the null might arise in Hall-Flavin tests. First, if expectations of permanent income are not formed rationally, or if (4) is not the correct form of the consumption function (due, for example, to significant liquidity constraints) the null hypothesis may be rejected. Rejections may also arise from violations of the simplifying assumptions embodied in the basic model which are not essential to the RE-PIH. In particular, any of (1) stochastic interest rates, (2) stochastic disturbances to the consumption function, or (3) adjustment lags induced by adjustment costs could cause rejections as well. With the Flavin specification ($k = 1$), a rejection of the null might arise simply because $k \neq 1$. Moreover, if consumption depends only upon current income innovations, but with a magnitude unrelated to the permanence of these innovations, the Hall-Flavin tests would fail to reject the RE-PIH, even though it would be false in this case.

2. A Test Based on the Effect of Income Innovations on Consumption

From the basic model, another test can be derived that lays the foundation for the tests that are the focus of this paper. Returning to (5), we see that the innovation to consumption is directly related to the innovation to permanent income which in turn depends on the revision of expectations of future labor income. Suppose expectations of future labor income were derived from an integrated autoregressive representation (of order p) of the labor income process, so that

$$(8) \quad \Delta x_t = \sum_{j=1}^{p} a_j \Delta x_{t-j} + \varepsilon_t ,$$

where $\Delta x_t = x_t - x_{t-1}$, the a_j are fixed autoregressive parameters, and ε_t is a white noise residual. The innovation to permanent income resulting from an innovation to labor would then be

$$(9) \quad \theta_t = (1 - \alpha)\, \Sigma \alpha^s (E_t - E_{t-1})\, x_{t+s}$$

$$= [1 - \sum_{j=1}^{p} \alpha^j a_j]^{-1}\, \varepsilon_t \equiv V \cdot \varepsilon_t \, .$$

V can be interpreted as the permanent income flow associated with the revision in expected future labor income induced by a unit innovation in x_t.

Assuming an interest rate, the system of equations (7) and (8) could be jointly estimated subject to the restriction (9), to obtain estimates of the parameters $a_j (j - 1, \ldots, p)$ and k. A less restricted version of the model would use in lieu of (7) the consumption equation,

$$(10) \quad c_t = (1 + r(1-k))c_{t-1} + \beta\varepsilon_t \, ,$$

to produce estimates of the parameters a_j, k and β. Using the relation $\theta_t = V\varepsilon_t$ from (9) to compare (10) with (7), one can see that the restriction on the model implied by the RE-PIH is simply $\beta = kV$. In the less restricted case, β is simply the propensity to consume out of a current period income innovation. Thus the test is whether the propensity to consume (β) out of a current income innovation is equal to the marginal propensity to consume (k) out of permanent income times the revision in permanent income (V) induced by a unit innovation to labor income.[3]

3. Two Measurement Issues

Although the exact restriction $\beta = kV$ is testable in principle, there are two potentially important measurement problems that may lead to a rejection of the RE-PIH even if it is true. In the remainder of this section, we will discuss these two complications and their implications for developing a cross-regime test of a generalized relation between β and kV. In order to simplify the discussion, we will initially assume that $k = 1$. In section 1.4, we will relax this assumption.

a. The Information Set

If the estimated model does not include all of the information available to agents when they make revisions in the estimates of permanent income, then the estimates of unanticipated income will be noisy, biasing estimates of β downward. Suppose, for example, that agents use the labor income forecasting equation

$$(11) \quad \Delta x_t = a\Delta x_{t-1} + bz_{t-1} + \varepsilon_t$$

where z_{t-1} is a white noise stochastic variable observed at $t - 1$ and orthogonal to both Δx_{t-1} and ε_t. Given this specification of the labor income process, the innovation to permanent income is given by

$$(12) \quad \theta_t = \left(\frac{1}{1-\alpha a}\right)\varepsilon_t + \frac{\alpha b}{(1-\alpha a)} z_t ,$$

and the consumption equation is

$$(13) \quad \Delta c_t = V\varepsilon_t + \mu_t ,$$

where $V = \left(\frac{1}{1-\alpha a}\right)$. We have added a consumption disturbance μ_t, a white noise residual assumed to be orthogonal to ε_t.

Now suppose that instead of estimating (11), one estimates

$$(14) \quad \Delta x_t = a\Delta x_{t-1} + \varepsilon_t' ,$$

where $\varepsilon_t' = bz_{t-1} + \varepsilon_1$ is orthogonal to Δx_{t-1}. Jointly estimating (14) with

$$(15) \quad \Delta c_t = \beta\varepsilon_t' + \mu_t'$$

would result in a violation of the restriction

(16) $\beta = V = \frac{1}{1-\alpha a}$,

since the unrestricted least squares estimate of β would be given by

(17) $\text{COV}\,(\Delta c_t,\ \varepsilon_t')/\text{VAR}(\varepsilon_t') = \frac{1}{(1-\alpha a)} \cdot \frac{\sigma_\varepsilon^2}{\sigma_\varepsilon^2 + b^2\sigma_z^2}$

which would deviate from V by the usual bias factor $\phi = \sigma_\varepsilon^2/[\sigma_\varepsilon^2 + \beta^2\sigma_z^2]$. Thus, the above analysis predicts the β - V relation to be

(18) $\beta = \phi V$.

Since ϕ is just identified, this is not a testable restriction on the model, at least within a single economic regime. On the other hand, if the bias factor ϕ is the same across countries, or at least uncorrelated with the V's, (18) would predict a positive relation between the β's and V's from different economic regimes which would be testable across regimes.[4]

b. Total Income and Endogenous Capital Accumulation

Flavin (1981) discusses the conceptual difference between the present value of a household's expected future total income and the present value of expected future (exogenous) labor income. In particular, since the former includes the future return on the endogenous accumulation of assets, it does not correctly summarize the constraint on consumption decisions. Although recognizing this problem, Flavin used total income in her empirical work, implicitly assuming that the differences are of not first-order importance. In an attempt to address this problem directly, Hayashi (1982) used data on labor income and non-labor assets in his tests of the RE-PIH. Unfortunately, data on labor income are not generally available across the panel of countries used in this paper.

Labor income, moreover, is not a completely satisfactory concept for a number of reasons. First, labor income is itself the result of endogenous human capital accumulation and labor supply decisions. Second, physical capital accumulation will affect labor income in the aggregate if labor and physical capital are complementary inputs, so that the time series properties of labor income will reflect the effects of savings decisions as well as endogenous disturbances. Furthermore, non-labor income will generally include exogenous unanticipated capital gains. Thus, the important distinction is not between total and labor income, but between the time series behavior of a measure of income which solely reflects exogenous disturbances and one which includes the effects of endogenous savings and labor supply behavior.

Ideally, one would like to derive an exact relation between the revision to permanent income implied by an innovation to exogenous income, V^*, and the revision to permanent income calculated from the innovation to observed total income, V_y. Although an exact relation would be difficult to obtain, we can deduce some facts about the relation between V^* and V_y that enable us to generate a useful approximation. First, when $V^* < 1$, an increase in exogenous income is partly temporary and would thus result in increased current saving. The latter would then increase total income in future periods making total income appear more persistent than exogenous income, so that $V_y > V^*$ will obtain. Conversey, when $V^* > 1$, $V_y < V^*$ will obtain. Finally, when $V^* = 1$, innovations to exogenous income are permanent and hence will be consumed. Since no endogenous capital accumulation (physical or human) results, $V_y = 1$ as well. From these observations, we can postulate a linear approximation to the relation between V^* and V_y of the form:

$$V^* = -b + (1 + b) V_y$$

with $b > 0$. Given this approximation, the RE-PIH predicts the $\beta - V_y$ relation to be

$$\beta = -b + (1 + b)\, V_y ,$$

which just identifies b and hence is not testable within a single regime. On the other hand, if the parameter b is a constant, or at least uncorrelated with V_y, across countries, then one would expect a positive relation between the β's and V's estimated from different economic regimes, which would be testable across regimes.

If the only departures from the basic model were due to the complications previously discussed (i.e., misspecified information sets and the difference between exogenous and total income), and if ϕ and b are the same across countries,[5] then the RE-PIH would predict

$$\beta_i = \phi V^*_i$$

or

$$\begin{aligned} \beta_i &= -\phi b + \phi(1 + b)\, V_i \\ &= \gamma_0 + \gamma_1 V_i \end{aligned}$$

where V_i now denotes the V_y for country i, $\gamma_0 = -\phi b$ and $\gamma_1 = \phi(1 + b)$. Note that this relation fails to identify both γ_0 and γ_1 (or ϕ and b) within any single regime. However, this relation can be the basis for the cross-regime tests set forth explicitly in the following section once the assumption that $k = 1$ is relaxed.

4. Allowing For a Long Run Propensity to Consume $k < 1$

Flavin (1981) assumed $k = 1$ and dealt with the nonstationarity of income via detrending, resulting in a levels specification for the income process and a first difference specification for consumption. Although detrending is appropriate in the case of deterministic (time trend) nonstationarity, it is not when stochastic unit-root nonstationarity is present.[6] As it turns out, the RE-PIH implies that both consumption and income will be stochastically nonstationary and that the stochastic trend in both consumption and income arises because $k < 1$.

With $k \neq 1$, the consumption equation for country i becomes (see (8))

$$\Delta c_{it} = r(1 - k_i)c_{it-1} + \beta_i \varepsilon_{it} + \mu_{it}$$

where we allow the k_i to vary across countries, but continue to assume that the interest rate is a known constant.[7] Similarly, the first differences of permanent and hence current income will include a term reflecting trend growth relative to lagged consumption,

$$\Delta y_{it} = \Delta y^p_{it} + \Delta \omega_{it}$$

$$= r(1-k_i) \frac{1}{k_i} c_{it-1} + \eta_{it} ,$$

where the properties of $\eta_{it} = \theta_{it} + \delta_{it} - \delta_{it-1}$ depend on the properties of x_{it}, and $E\eta_{it} = 0$ if Δx_{it} is stationary. Finally, in order to correct for heteroskedasticity, we weight the data by the reciprocal of lagged consumption which yields

$$\Delta y'_{it} = r(1-k_i) \frac{1}{k_i} + \eta'_{it} ,$$

where the primes (') indicate weighted data and η'_{it} is homoskedastic and serially correlated.[8] If η'_{it} has the representation $\eta'_{it} = R_i(L)\, \varepsilon'_{it}$, where $R_i(L)^{-1} = 1 - \sum_{j-1}^{p} \rho_{ji} L^j$, we obtain

$$(19)\quad \Delta y'_{it} = \lambda_i + \sum_{j=1}^{p} \rho_{ji} \Delta y'_{it-j} + \varepsilon'_{it}$$

$$\Delta c'_{it} = \delta_i + \beta_i \varepsilon'_{it} + \mu'_{it}$$

$$V_i = [1 - \sum_{j=1}^{p} \alpha^j \rho_{ji}]^{-1}$$

where $\delta_i = r(1 - k_i)$, $\lambda_i = R_i(L)\ \delta_i/k_i$, and ε'_{it} is white noise. The cross-regime restriction on the β_i, of interest when $k_i \neq 1$, becomes

$$(20)\quad \beta_i = k_i(\gamma_0 + \gamma_1 V_i).$$

Imposing (20) as a restriction on system (19) constitutes a test of the "cross-regime" relation between the regime-specific β_i and $k_i V_i$ implied by the RE-PIH and overidentifies γ_0 and γ_1 or, alternatively, the "deeper" parameters ϕ and b. We can test for a general positive relation between the β_i and $k_i V_i$ by testing $\gamma_1 = 0$ against $\gamma_1 > 0$. We can test separately whether there is significant excluded information by testing $\phi = 1$ against $\phi < 1$. Finally, we also have N - 2 overidentifying restrictions with which to test the general validity of the restrictions implied by (20).[9]

II. THE EMPIRICAL RESULTS

1. The Data

All the data for this paper are taken from the International Financial Statistics data base of the International Monetary Fund. Our tests require continuous and consistent time series data on real income per capita and real consumption expenditures per capita. We found thirty countries that had uninterrupted data on nominal consumption expenditures, C_t, nominal gross national product, Y_t, consumer prices, P_t, population, POP_t, and government spending,

G_t, over the 1951-1979 period. For the thirty countries in our panel, these series form the data set for this study. We then calculated per capita real consumption as $c_t = C_t/(P_t \cdot POP_t)$, and per capita real "disposable" income as $y_t = (Y_t - G_t)/(P_t \cdot POP_t)$. This measure of private sector "disposable" income represents the income available for disposition by the private sector as either private consumption or private investment.[10] Note also that c_t includes all consumer expenditures, including durable consumption goods that may more appropriately be considered investment. While this may affect our tests, IFS data do not permit any adjustment for consumer durables.

2. The Hall-Flavin Tests

Before proceding to the β - kV relation across regimes, we undertake Hall-Flavin tests of the RE-PIH for each of the thirty countries in our panel. We do so for three reasons. First, doing so provides a link between the extant literature and our proposed cross-regime evidence. Second, such tests over thirty countries situate the results for the US in a broader context. Third, the Hall-Flavin tests correspond to the standard rational expectations procedure of testing the restricted model (in which only unanticipated income appears in the consumption equation (19)) against an unrestricted alternative (in which current and lagged income appear separately).

Our implementation of the Hall-Flavin tests is based on the following equation:

$$(21) \quad \Delta c'_{it} = b_{0i} + b_{1i}\Delta y'_{it-1} + b_{2i}\Delta y'_{it-2} + v'_{it} \qquad i = 1,\ldots, 30; \; t = 1953,\ldots, 1979.$$

Although both one and three lags of Δy_i were also considered, the results did not differ much from those with two lags. The null hypothesis associated with the RE-PIH is H_0: $b_{1i} = b_{2i} = 0$. We test this via an F-test against the nonspecific alternative H_1: b_{1i} or $b_{2i} \neq 0$.

The results of the Hall-Flavin tests are reported in Table 1. The F statistics for 2 numerator and 22 denominator degrees of freedom are reported in column 2, with the corresponding marginal significance levels (MSL) given in column 3. First note that for the US, the marginal significance level is .08, which is similar to Hall's finding but is not as low as Flavin's. For various reasons, however, our results are not directly comparable with either Hall's or Flavin's: we measure consumption expenditures and disposable income differently, we use annual instead of quarterly data, we use differenced instead of level data. These noncomparabilities notwithstanding, our results for the US seem to fall in the same ballpark as Hall's and Flavin's.

A comparison of the US with the other countries in the panel reveals a number of interesting features. First, only five of the thirty countries have MSL's less than or equal to that of the U.S.--Finland, Ireland and Iceland, which reject the null hypothesis at less than the 5% level, and Peru and South Africa, which reject at the 8% level. Second, there is no apparent correlation between the MSL's for testing lags of Δy in the consumption equation and lags of Δy in the income equation. Thus, our failure to reject lags of Δy, in the consumption equation is not because lag of Δy do not appear in the income equation. Third, the distribution of MSL's across countries appears uniform over the range [0,1]. In fact, under the null hypothesis, the distribution of the MSLs should be U(0,1), whereas under the alternative it should be skewed toward 0. A Kolmogorov-Smirnov (K-S) test of whether the empirical distribution of the MSL deviates significantly from the null distribution U(0,1) yields a statistic (for the values in column 2) of .95 with a corresponding MSL greater than 20%. At the bottom of Table 1, we also display the number of countries (under the null) expected to fall below a given probability along with the actual number.

On the whole, the result of Table 1 are not at gross odds with the RE-PIH. Consequently, we will proceed to the cross-regime tests of the RE-PIH maintaining the assumption that only unanticipated income causes revisions in consumption.

3. Cross Regime Evidence of the RE-PIH

We first estimate system (19), with lag length p set equal to two, over each of the thirty countries without imposing the cross-regime restriction (20).[11] This is referred to as the "unrestricted" system even though the within-country restriction that only unanticipated income (ε'_{it}) appears in the consumption equation is maintained. We first use these unrestricted estimates, to investigate the cross-regime relation between the β_i and V_i, and among β_i, k_i and V_i. We then present the results obtained by estimating (19) subject to the cross-regime restrictions (20) and compare them to the unrestricted results.

The results of estimating (19) (with $\alpha = .9$) for each of the thirty countries in our panel are reported in Table 2.[12] Columns 2 and 4 present the estimated $\hat{\beta}_i$ and and $\hat{V}_i$; columns 3 and 5, their respective (asymptotic) standard errors. The $\hat{\beta}_i$ range between .13 and .91 with a mean of .47. The standard errors of the $\hat{\beta}_i$ sufficiently small that $\beta_i = 0$ is easily rejected for all but a few countries. The $\hat{V}_i$ range between .63 and 1.28 with a mean of .93.[13] The standard errors of the $\hat{V}_i$ are somewhat larger than those for the $\hat{\beta}_i$. Since we wish to draw inferences concerning systematic differences in the β_i and V_i, we ask whether we can reject $\overline{\beta}_i = \beta$ and $V_i = \overline{V}$ for the full panel of thirty countries (sixty equations). The likelihood ratio statistic for this null hypothesis is 171.6, which, given 58 degrees of freedom, is significant at better than the 1% level.

One striking feature of the results in Table 2 concerns the relation between β and V. For each of the 30 counries on our panel, $\hat{\beta} < \hat{V}$ obtains. The means are $\bar{\beta} = .47$ and $\bar{V} = .93$. Thus, to a first approximation we are finding that consumption, far from being excessively sensitive to income innovations, appears to be undersensitive. It is important to recall that the key to this result is our maintained assumption that disposable income is characterized by a unit root and hence must be specified in differenced form. Accounting for a unit root will yield systematically higher values of V in small samples than that obtained from detrended level data. Thus, our finding is fundamentally related to the evidence in favor of a unit root in GNP such as found in Nelson and Plosser (1982), Campbell and Mankiw (1987a, 1987b), Kormendi and Meguire (1987) and Schwert (1987). Our results at this point for thirty countries accord well with those obtained independently by Deaton (1986) and Campbell and Deaton (1987) for US data, who find undersensitivity when they estimate the income process in terms of first differences.

We now present in Figure 1, the scatter diagram for $\hat{\beta}_i$ against $\hat{V}_i$, which reveals a positive relation. The ordinary correlation is .55 with a t-statistic of 3.4, which rejects the hypothesis that the $\hat{\beta}_i$ and $\hat{V}_i$ are uncorrelated at less than the 1% level. An alternative nonparametric test based on the rank correlation of $\hat{\beta}_i$ and $\hat{V}_i$ is also of interest because it is robust both to the (monotonic) functional form of the relation and the distributional properties of the estimates. The rank correlation between $\hat{\beta}_i$ and $\hat{V}_i$, .46, with a corresponding t-statistic of 2.7, also rejects the hypothesis that $\hat{\beta}_i$ and $\hat{V}_i$ uncorrelated in favor of the hypothesis that they are positively related at better than the 1% level.[14]

In order to account for a less than unit long run propensity to consume, and investigate the b - kV relation, estimates of k are required. As shown in the discussion surrounding (19), it is possible to obtain estimates of k directly from the δ, under the maintained assumption that all countries have the same constant discount factor α. In column 1 of Table 3, we report the values of k^{δ} derived from the δ reported in Table 2 under the assumption that $\alpha = .90$. If transitory income has mean zero, $\overline{c}/\overline{y}$ will be a consistent estimator of k, where $\overline{c}$ and $\overline{y}$ are averaged over time. We denote this estimate as $\hat{k}$ and report its values in column 2 of Table 3. There is a positive correlation between k^{δ} and $\hat{k}$ (with Israel as an outlier), and the means of the two are virtually identical, which lends some support to the assumption that $\alpha = .90$. In subsequent empirical results we use the estimate $\hat{k}$, mainly because k^{δ} is very sensitive to the constant interest rate assumption. If r in fact does vary across countries, this will add noise to k^{δ}, whereas $\hat{k}$ would not be affected.[15]

Accounting for $k < 1$ has implication for the issue of the undersensitivity of consumption to income innovations. The correct implication for drawing inferences as to undersensitivity is $\beta < kV$, not $\beta < V$. Table 3 presents estimates of kV for our 30 countries. The mean of kV over all 30 countries is .68. When compared to the mean of β of .47, we still find some undersensitivity, but not nearly as pronounced as comparing β with V alone. If, in fact, β is downward biased due to excluded information as discussed in Section I.3.a, that could easily account for the remaining undersensitivity.

The average undersensitivity that we find of approximately 30% (1 - .44/.68) should be compared to the 70% undersensitivity found by Deaton (1986) and Campbell and Deaton (1987) for the US. Our results for the US showed undersensitivity of about 60% (1 - .44/.93) when $k < 1$ is not taken into account, and 50% when allowing for $k < 1$. Thus, the US appears somewhat more undersensitive than the average country in our sample, which along with allowing for $k < 1$, partly explains Deaton's and Campbell and Deaton's strong results.[16]

The scatter diagram for $\hat{\beta}_i/\hat{k}_i$ and $\hat{V}_i$ presented in Figure 2 suggests a positive relation between the two quantities. In order to for test a positive β - kV relation, as well as to find estimates of γ_0 and γ_1 in (20), we report in (22) the results of a simple OLS projection of $\hat{\beta}_i$ onto $\hat{k}_i$ and $\hat{k}V_i$. Because the β_i and V_i are errors-in-variables estimates of the true β_i and V_i, especially when (19) is estimated over a small sample, an OLS estimate of γ_1 in (22) will in all likelihood be biased downward. Thus we also report in (23) the estimated reverse regression from which an approximate upper bound for γ_1 can be calculated.[17] Equation (23') inverts (23) to show the implied slope coefficient for the β - kV relation. The OLS standard errors are given in parentheses below each estimated coefficients. The implied values for ϕ and b are also reported.

(22) $\hat{\beta}_i = \underset{(.17)}{-.02}\ \hat{k}_i + \underset{(.18)}{.70}\ \bar{k}_i\hat{V}_i + \upsilon_i$ $\qquad \phi = .68 \qquad b = .03$

$R^2 + .93$ s.d.r. = .138 N = 30

(23) $\hat{k}_i\hat{V}_i = \underset{(.08)}{.60}\ \hat{k}_i + \underset{(.13)}{.49}\ \hat{\beta}_i + \tilde{\upsilon}_i$ $\qquad \phi = .82 \qquad b = 1.49$

The results presented thus far indicate that there is a positive relation between the $\hat{\beta}$ and $\hat{k}\hat{V}$ for the thirty countries in our panel, and that this relation is not simply due to a positive relation between the $\hat{\beta}$ and the $\hat{k}$. However, because of sampling errors in the $\hat{\beta}$ and $\hat{V}$ (and $\hat{k}$, for that matter), it is difficult to make precise inferences about the relation between the true β and kV, nor can we test whether the RE-PIH restriction (20) is violated. Thus, we turn to cross-regime results obtained by estimating (19) subject to the restrictions on the β_i given by (20). The estimates of γ_0 and which are identified when (20) is imposed, the corresponding values for ϕ and β, and the likelihood ratio statistic for testing the restrictions are reported in Table 4.

Table 4

RESULTS FOR ESTIMATING SYSTEM (19) WITH (20)

$\hat{\gamma}_0$	$\hat{\gamma}_1$	$\hat{\phi}$	$\hat{b}$	likelihood ratio statistic	MSL
-1.29 (.40)	2.05 (.42)	.76 (.05)	1.68 (.49)	38.27	.093

Several results bear discussion here. First, $\hat{\gamma}_1$, which is similar to that obtained from the auxiliary reverse regressions reported earlier, is significantly different from zero at the 1% level (t = 5.0). Thus, the positive β - kV relation predicted by the RE-PIH is even more strongly supported in this test. Second, $\hat{\phi}$ is significantly different from both zero and one, indicating that the downward bias due to excluded information in the income equation is also important. Moreover, the estimated ϕ of around .75 would fully explain the average undersensitivity of 30% obtained earlier. Taken together, such results seem to indicate that consumption is not nearly as undersensitive as it may seem on first inspection. Finally, the set of overidentifying restrictions not rejected at the 5% level, indicating that the non-stochastic b - kV relation in (20) is not grossly at odds with the data.

4. Concluding Remarks

Our empirical investigation of the behavior of consumption and income in thirty countries reveals a significant positive relation between the marginal propensity to consume out of current income innovations (β) and the persistence of current income innovations (V), which may also be interpreted as a measure of the average ratio of permanent to current income innovations. This positive relation becomes clearer when we also account for variations in the marginal propensity to consume out of permanent income (k) across countries. Neither simple Keynesian models nor arbitrary distributed lag models of consumption spending would predict this positive relation; hence the results lend support to the RE-PIH.

Our finding that on average over 30 countries, $\hat{\beta} < \hat{k}\hat{V}$ suggests that consumption appears somewhat undersensitive to income innovations, a result that accords with those of Deaton (1986) and Campbell and Deaton (1987) for the US. However our results over 30 countries revealed an average undersensitivity of only 30%, whereas Deaton (1986) and Campbell and Deaton (1987) find approximately 70% undersensitivity for the US. Finally, indirect estimates of the effects of excluded information suggest that the remaining 30% undersensitivity may be due to the existence of excluded information. Hence averaging across our panel of 30 countries, consumption is generally appropriately sensitive to income innovations.

FIGURE 1

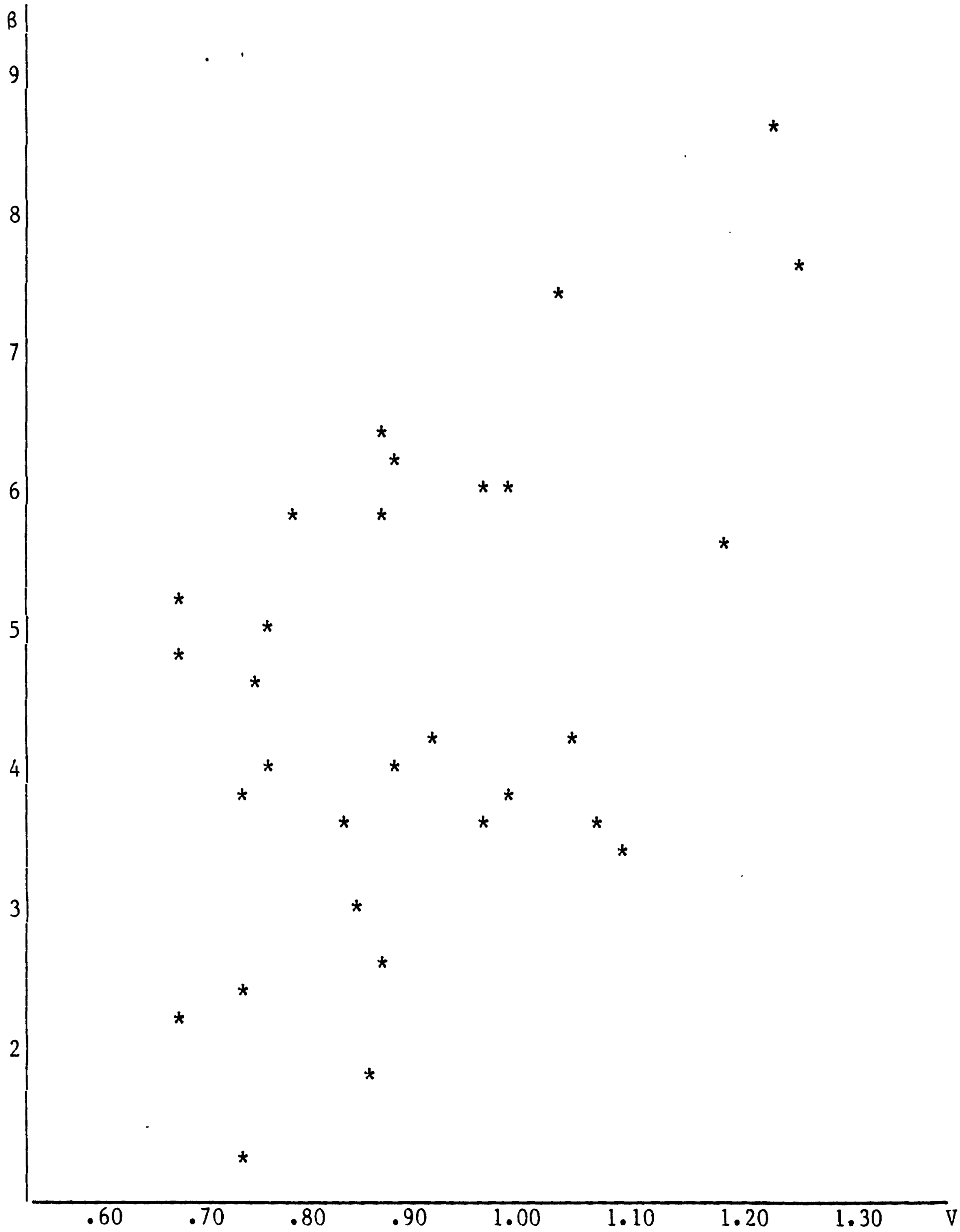

FIGURE 2

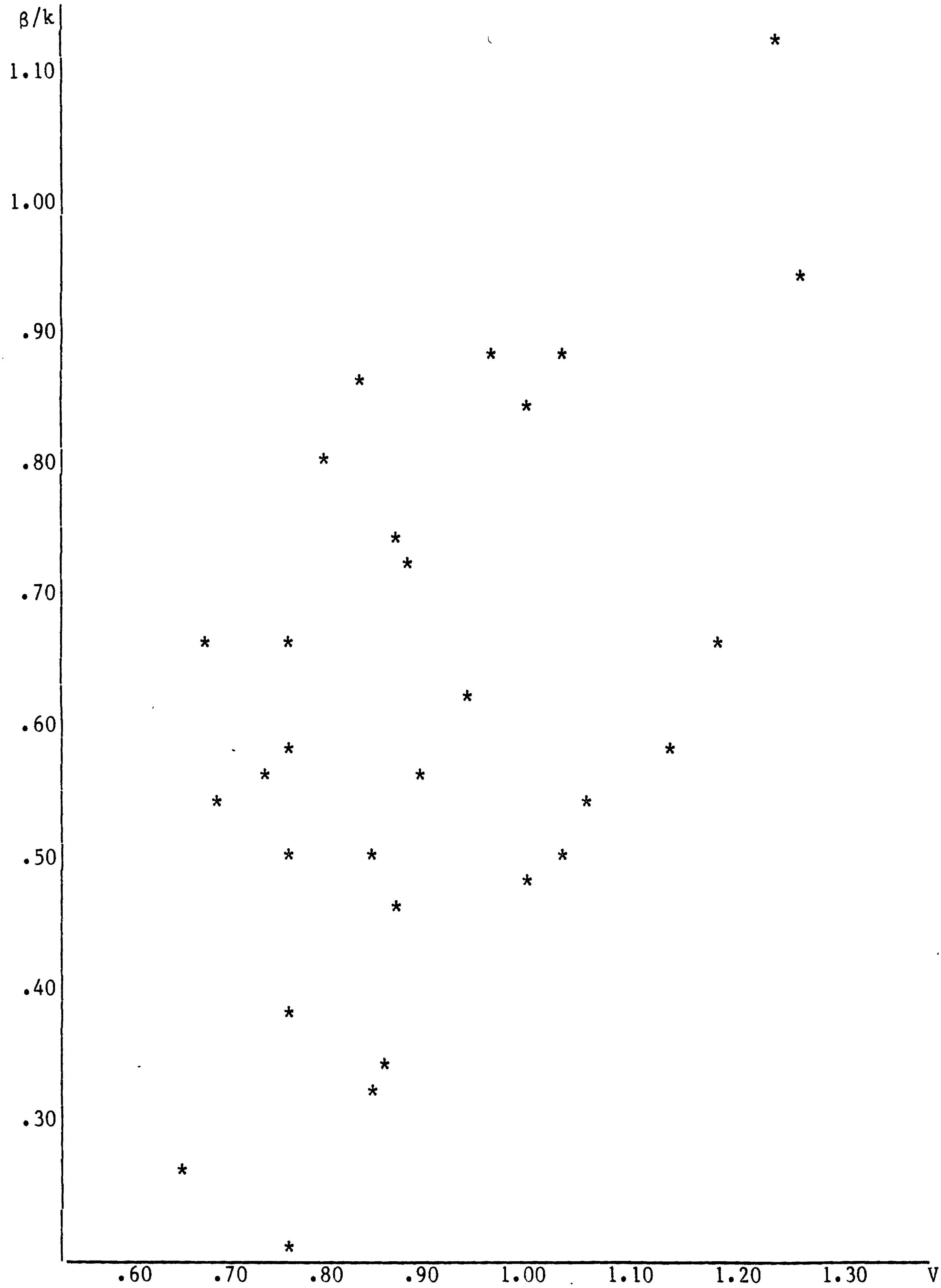

Table 1

RESULTS FOR THE HALL-FLAVIN TESTS (1954-1979)

$$\Delta c'_{it} = b_{0i} + b_{1i}\Delta y'_{it-1} + b_{2i}\Delta y'_{it-2} + v'_{it}$$

$$H_0: \ b_{1i} = b_{2i} = 0'$$

$$H_1: \ b_{1i}, b_{2i} \neq 0$$

Country	F(2, 22)	MSL
1. Bolivia	.38	.69
2. Columbia	1.29	.30
3. Denmark	.81	.46
4. Dom. Republic	.59	.56
5. Finland	7.05	.00
6. France	.09	.92
7. Guatemala	.40	.67
8. New Zealand	1.74	.20
9. Panama	.15	.86
10. Paraguay	.59	.56
11. Sweden	1.85	.18
12. Switzerland	2.58	.10
13. Venezuela	1.10	.35
14. Austria	2.70	.21
15. Australia	.13	.87
16. Canada	.04	.96
17. Costa Rica	1.55	.23
18. Germany	2.68	.09
19. Greece	.08	.92
20. Ireland	4.32	.03
21. Netherlands	1.21	.32
22. Norway	.83	.45
23. Peru	2.89	.08
24. Philippines	.20	.82
25. So. Africa	2.86	.08
26. Sri Lanka	.68	.52
27. United Kingdom	.25	.78
28. United States	2.82	.08
29. Iceland	3.86	.04
30. Israel	1.44	.26

	MSL $\leq$ .10	MSL $\leq$ .30	MSL $\leq$ .50	MSL $\leq$ .70	MSL $\leq$ .90
expected #	3	9	15	21	27
actual #	8	14	18	23	27

K - S = .95 MSL(K-S) > .20

Table 2

RESULTS FOR ESTIMATING SYSTEM (19) (1956-1979)

$$\Delta y'_{it} = \lambda_i + \rho_{1i}\Delta y'_{it-1} + \rho_{21}\Delta y'_{it-2} + \varepsilon'_{it}$$

$$\Delta c'_{it} = \delta_i + \beta_i\varepsilon'_{it} + \mu'_{it}$$

$$V_i = [1 - \alpha\rho_{1i} - \alpha^2\rho_{21}]^{-1}(\alpha = .9)$$

	Country	$\hat{\beta}$	$\hat{\sigma}_\beta$	$\hat{V}$	$\hat{\sigma}_V$	$\hat{\delta}$	$\hat{\sigma}_\delta$	$\hat{\lambda}$	$\hat{\sigma}_\lambda$
1.	Bolivia	.74	.04	1.06	.03	.005	.031	.013	.040
2.	Columbia	.27	.13	.87	.21	.020	.008	.035	.015
3.	Denmark	.37	.09	1.00	.21	.025	.007	.036	.015
4.	Dom. Republic	.63	.12	.85	.14	.024	.013	.031	.016
5.	Finland	.58	.09	.88	.10	.039	.011	.064	.017
6.	France	.62	.11	.98	.17	.039	.006	.059	.013
7.	Guatemala	.47	.12	.67	.09	.016	.005	.035	.009
8.	New Zealand	.61	.13	1.01	.17	.017	.011	.024	.013
9.	Panama	.91	.28	1.26	.33	.028	.014	.030	.011
10.	Paraguay	.80	.14	1.28	.30	.035	.010	.034	.013
11.	Sweden	.35	.07	.96	.13	.022	.004	.032	.009
12.	Switzerland	.34	.04	1.06	.10	.028	.004	.041	.011
13.	Venezuela	.25	.07	.74	.14	.042	.015	.094	.040
14.	Austria	.36	.07	1.03	.17	.043	.005	.065	.015
15.	Australia	.44	.06	.92	.12	.025	.005	.039	.011
16.	Canada	.36	.05	.82	.09	.029	.004	.053	.011
17.	Costa Rica	.62	.09	.87	.11	.033	.010	.044	.015
18.	Germany	.20	.05	.86	.14	.046	.006	.087	.024
19.	Greece	.42	.06	.88	.11	.049	.006	.085	.017
20.	Ireland	.45	.08	.73	.09	.024	.006	.047	.013
21.	Netherlands	.32	.09	.85	.15	.038	.005	.064	.015
22.	Norway	.13	.08	.74	.15	.028	.005	.064	.019
23.	Peru	.53	.12	.63	.18	.017	.008	.039	.012
24.	Philippines	.39	.10	.75	.13	.021	.005	.052	.012
25.	So. Africa	.37	.08	.71	.08	.012	.006	.042	.012
26.	Sri Lanka	.58	.13	1.20	.25	.042	.013	.038	.017
27.	United Kingdom	.50	.08	.75	.09	.021	.004	.042	.009
28.	United States	.44	.03	1.07	.09	.023	.004	.029	.010
29.	Iceland	.58	.07	.78	.07	.044	.014	.087	.023
30.	Israel	.24	.06	.67	.10	.044	.007	.068	.021
Average		.47		.93		.029		.049	

Note: Likelihood Ratio Statistic ($\chi^2(58)$) for H_0: $\beta_i = \bar{\beta}$ and $V_i = \bar{V}$ ($\forall i$) = 171.6 MSL < .010.
The estimates and standard errors of the autoregressive parameter (ρ_j) of the income equations as well as the standard errors of the two equations are reported in Table 2A.

Table 2A

	Country	$\hat{\rho}_1$	$\hat{\sigma}_{\rho_1}$	$\hat{\rho}_2$	$\hat{\sigma}_{\rho_2}$	$\hat{\sigma}_\varepsilon$	$\hat{\sigma}_\mu$
1.	Bolivia	.018	.032	.045	.025	.187	.032
2.	Columbia	-.328	.189	.184	.191	.057	.035
3.	Denmark	-.026	.168	.031	.174	.059	.025
4.	Dom. Republic	-.136	.137	-.066	.148	.071	.039
5.	Finland	.063	.130	-.241	.120	.070	.027
6.	France	-.032	.140	.007	.142	.033	.016
7.	Guatemala	-.222	.152	-.373	.157	.032	.017
8.	New Zealand	.063	.161	-.056	.160	.059	.034
9.	Panama	.161	.175	.076	.180	.040	.052
10.	Paraguay	.115	.144	.145	.151	.045	.030
11.	Sweden	.130	.145	-.202	.145	.037	.012
12.	Switzerland	.157	.106	-.100	.103	.048	.008
13.	Venezuela	-.243	.198	-.154	.215	.169	.057
14.	Austria	.156	.155	-.133	.167	.046	.015
15.	Australia	-.042	.112	-.067	.115	.044	.013
16.	Canada	-.045	.113	-.227	.118	.044	.011
17.	Costa Rica	.122	.131	-.318	.136	.063	.026
18.	Germany	.098	.177	-.315	.182	.088	.022
19.	Greece	-.019	.117	-.145	.111	.060	.017
20.	Ireland	-.355	.157	-.066	.155	.052	.018
21.	Netherlands	-.008	.174	-.209	.183	.041	.017
22.	Norway	-.203	.199	-.204	.195	.063	.024
23.	Peru	-.436	.161	-.228	.154	.049	.027
24.	Philippines	-.293	.167	-.076	.160	.037	.018
25.	So. Africa	-.068	.151	-.434	.145	.052	.018
26.	Sri Lanka	.193	.164	-.012	.172	.072	.044
27.	United Kingdom	-.111	.127	-.287	.132	.031	.011
28.	United States	.077	.062	-.008	.069	.045	.006
29.	Iceland	-.243	.117	-.088	.107	.103	.028
30.	Israel	-.368	.170	-.201	.174	.087	.024

Table 3

	Country	k_i^δ	$\bar{k}_i$	$\hat{\beta}_i/\bar{k}_i$	$\bar{k}_i.\hat{v}_i$
1.	Bolivia	.95	.85	.87	.90
2.	Columbia	.82	.80	.34	.70
3.	Denmark	.78	.74	.50	.74
4.	Dom. Republic	.79	.84	.75	.71
5.	Finland	.65	.68	.85	.60
6.	France	.65	.71	.87	.69
7.	Guatemala	.85	.88	.53	.59
8.	New Zealand	.85	.73	.84	.74
9.	Panama	.75	.80	1.14	1.01
10.	Paraguay	.68	.85	.94	1.09
11.	Sweden	.81	.71	.49	.68
12.	Switzerland	.75	.67	.51	.71
13.	Venezuela	.62	.64	.39	.48
14.	Austria	.61	.68	.53	.70
15.	Australia	.78	.72	.61	.66
16.	Canada	.74	.74	.49	.60
17.	Costa Rica	.70	.85	.73	.74
18.	Germany	.59	.65	.31	.56
19.	Greece	.55	.78	.54	.69
20.	Ireland	.78	.80	.56	.58
21.	Netherlands	.66	.69	.46	.59
22.	Norway	.74	.66	.20	.49
23.	Peru	.84	.82	.65	.52
24.	Philippines	.81	.79	.49	.60
25.	So. Africa	.89	.70	.53	.50
26.	Sri Lanka	.62	.87	.67	1.04
27.	United Kingdom	.81	.77	.65	.58
28.	United States	.79	.79	.56	.84
29.	Iceland	.60	.72	.81	.56
30.	Israel	.60	.95	.25	.64
	Average	.74	.76	.60	.68

FOOTNOTES

[1]See Muth (1960) for the conditions under which exponential lag weights reflect the optimal forecast of future income.

[2]Another important contributor to the RE-PIH is Sargent (1978), who attempted tests based on the systematic part of income rather than on income innovations. See Flavin (1981) for a discussion of the relation between Hall's and Sargent's papers. Hayashi (1982), using data on consumption, labor income and non-human wealth, derives more general tests along the lines of Sargent, Hall and Flavin.

[3]Bilson (1980) and Hall and Mishkin (1982) attempted to test the strict equality $\beta = V$. Bilson used a two-step procedure to estimate β and a range of V's (corresponding to a range of interest rates) and found that, for reasonable values of the interest rate, the hypothesis $\beta = V$ could not be rejected for the three countries in his study. Using household panel data for the U.S., Hall and Mishkin find that the effect of income innovations on consumption (β) is large relative to the effect predicted by the RE-PIH (V), again assuming reasonable interest rates. More recently, Bernanke (1985) explicitly accounts for durable goods and finds evidence that β approximately equals V.

[4]If the excluded information, z_{t-1}, were correlated with Δx_{t-1}, then V would be biased as well, with the bias being upward (downward) if the correlation between z_{t-1} and Δx_{t-1} were positive (negative). If the biases in β and V are constant or uncorrelated across countries, the cross regime tests are unaffected, although cross-country measurement error in β and V will tend to obscure the true β - V relation. If the biases are negatively correlated acros s countries, as one might expect if z_{t-1} is positively correlated with Δx_{t-1} thereby imparting an upward bias to V (and a downward bias to β), a true true positive β - V relation would be further obscured. If the biases are

positively correlated, the possibility of a spurious positive observed $\beta - V$ relation exists. For these reasons, further research aimed at improving the income forecasting equation, to include, for example money supply, stock returns and interest rates, would be of interest. An interesting recent proposal of Campbell's (1987) would include lags of saving in the income equation to proxy for such excluded information, since saving will adjust endogenously to offset the effects of current information on the future income stream. However, exogenous shocks to saving would affect future income in the opposite direction, and therefore interfere with the proxy effect.

[5] See fn 8 for the effects of relaxing the assumption of ϕ and b constant.

[6] Granger and Newbold (1974), Plosser and Schwert (1977, 1978), and Nelson and Kang (1981) discuss some of the problems involved in inappropriately assuming only deterministic non-stationary when stochastic non-stationarity is present. Plosser and Schwert (1977, 1978), for example, suggest that fewer problems are likely to be encountered by overdifferencing than by differencing. Moreover, Plosser and Nelson (1982) and Schwert (1987) have examined a large number of aggregate annual time series for the U.S., up to 100 years long, and found strong evidence of stochastic non-stationarity for virtually all series. Recently, Kormendi and Meguire (1987) have undertaken a variety of new tests for unit roots in GNP over 30 countries using samples spanning up to a century in length and find very strong evidence in favor of a unit root and against trend stationary models. Campbell and Mankiw (1987) found similar results for seven countries.

[7] Note that the disturbances in the consumption equations (μ_{it}) may reflect disturbances in the consumption function (4) in addition to an imperfect measure of income innovations (ε_{it}). If consumption innovations (due to a temporary disturbance to time preference, for example) and income innovations are

positively correlated, as would be predicted by both Keynesian and market clearing (see Barro 1984) models, estimates of β will be upward biased, yielding spurious "excess sensitivity". See also Michener (1984) for market-clearing excess sensitivity.

[8]Hayashi (1982) makes a similar correction for heteroskedasticity.

[9]If ϕ or b are not equal across countries, the relation between the β_i and $k_i V_i$ would not be exact but would include country specific "residuals," producing a relation of the form

$$\beta_i = k_i(\gamma_0 + \gamma_1 V_i) + \nu_i$$

where the ν_i summarize the country-specific factors. If the ν_i are orthogonal to the $k_i V_i$ and k_i, then imposing (20) on (19) will produce consistent, but inefficient estimators of γ_0 and γ_1 since the errors in the consumption equations will be heteroskedastic (conditional on $\Delta y'_{it}$, $\Delta y'_{it-j}$ and ρ_{ji}). Preliminary attempts to correct for such heteroskedasticity had no major effect on either the estimated parameters or their standard errors. A test of the N-2 overidentifying restrictions in (20) is also a test of whether the country-specific factors in the β - kV relation can be ignored; failing to reject these restrictions (see results in Table 4) suggests that the country-specific factors may not be too important.

[10]Results obtained using per capita real GNP differed little from those reported in the paper.

[11]We estimate the 60 equation system using nonlinear system OLS (PROC SYSNLIN OLS in SAS). This is equivalent to estimating (19) for each country separately with no cross-regime restrictions imposed.

[12]Experimentation with interest rates in the range .01 to .25, i.e., $.8 \leq \alpha \leq .99$, had little effect on our results.

[13]Although there is no reason to expect the income processes of different countries to have the same persistence characteristics, it would be of interest to know why V_i varies across countries. To shed light on this issue we examined two potential factors for explaining variation in V_i--the openness of a country to foreign trade as measured by the mean ratio of exports to income (MXY), and the size of the government sector as measured by the mean ratio of government spending to income (MGY). (Friedman (1957) argued that a farm non-farm distinction may be important, which suggests a measure such as the mean ratio of farm income to total income. We did not pursue this.) We found MXY_i and V_i to be negatively related, with $\mathrm{Corr}(MXY_i, V_i) = -.38$ being different from zero at the 4% level. This is consistent with the hypothesis that income from exports is generally more transient than domestic income. We also found MGY_i and V_i to be negatively related with $\mathrm{Corr}(MGY_i, V_i) = .33$ being significantly different from zero at the 8% level. This negative sign is contrary to the usual presumption that a large government sector stabilizes real income.

[14]In addition to the issues already discussed which are expected to cause departures from the exact $\beta = kV$ relation, it is also conceivable that a positive correlation between the sampling errors $(\hat{\beta}_i - \beta_i)$ and $(\hat{V}_i - V_i)$ could produce a spurious positive relationship between $\hat{\beta}$ and $\hat{V}$. We examined the correlations between the sampling errors and found them to be generally small--none are greater than .30, only 6 of the 30 are greater than .20, and 10 are negative. A potentially more serious problem concerns measurement error in income innovation may bias estimates of β towards 0. Likewise, if the measurement error were more transitory than the true income V may also be biased towards 0. Thus the measured correlation between β and V is potentially biased upwards. If this were the case, however, any positive correlation between β and

V observed over the entire sample would be attenuated when computed over subsamples stratified by an ordinal measure of data quality. Summers and Heston (1984) group countries into four categories based on the quality of their national accounts data. We repeated the cross-country correlations controlling for data quality using the categories of Summers and Heston, and found the correlations to be unaffected. Also, if the results in Table 1 were due to measurement error varying systematically across countries, one would expect the correlations to decline upon controlling for the standard error of the income equation. Again, this did not occur.

[15]When k_i^{δ} was used rather than $\hat{k}_i$, the value of q_1 declined somewhat. We also used $\hat{k}_i$ in conjunction with estimates of δ_i to generate estimates of country-specific interest rates $\hat{r}_i = \delta_i/(1 - \hat{k}_i)$. With the exception of Israel, these $\hat{r}_i$ were all between .05 and .30, with a mean of .12 (recall that $\alpha = .9$ corresponds to $r = .11$). Moreover, using these $\hat{r}_i$ in our calculations of V actually improved the β-V and β-kV relations somewhat.

[16]Deaton (1986) and Campbell and Deaton (1987) obtain a V of approximately 1.6 whereas our V is around 1.1. Since our V is obtained with annual data and two years lags, we suspect that allowing eight quarters of lags in their quarterly specification would bring their V down somewhat.

[16]Using the inverse of the slope coefficient of the reverse regression as in estimator for γ_1 is only an approximate upper bound because the covariance of the sampling errors, $E[(\hat{\beta}_i - \beta_i)(\hat{k}_i\hat{V}_i - k_iV_i)]$ cannot be assumed to be zero, and because of the presence of a second regressor, $\hat{k}_i$.

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