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ABSTRACT

What idiosyncratic consumption risks can countries trade away on international asset markets? This paper develops an empirical methodology for answering the question. The tests are based on the proposition that in an integrated world asset market with representative national agents, the ex post difference between two countries' intertemporal marginal rates of substitution in consumption is uncorrelated with any random variable on which contractual payoffs can be conditioned. This result is applied to annual time-series data for the seven largest industrial countries over 1950-88. Of these countries, Germany seems to have been most successful at internationally diversifying its consumption risks.

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Introduction

This paper develops consumption-based tests of alternative hypotheses about countries' participation in world financial markets. The underlying methodology in principle can throw light both on the efficiency of international trade in noncontingent assets, and on the range of contingent assets countries use to diversify idiosyncratic national risks.

My empirical analysis of aggregate consumption behavior in the seven largest industrial countries is consistent with the hypothesis that most became increasingly integrated into world markets for risk sharing over the 1973-1988 portion of the post-World War II era. In this group, Germany stands out as having reduced idiosyncratic consumption risk through trade to an exceptional degree. For France, Italy, Japan, and the United Kingdom, the interpretation of aggregate consumption behavior after 1973 is more ambiguous, as one might expect given these countries' comparatively late capital-account liberalization drives. Country size makes the United States record difficult to assess, while Canada throws up a puzzle, an apparent reduction after 1973 in the global diversification of its consumption risks.

Empirical studies of international trade in consumption risk generally reach the conclusion that markets for risk function imperfectly at the international level, and certainly less efficiently than do domestic markets. Atkeson and Bayoumi (1992), for example, argue that the national diversification of regional incomes within the United States is significantly greater than the international diversification of European national incomes.[1] French and Poterba (1991), Golub (1991), and Tesar and Werner

(1992) document what appears to be a domestic-asset bias in the security portfolios of major industrial countries.

Yet another approach, proposed by Leme (1984) and Scheinkman (1984), starts from the observation that national consumption levels should move in a synchronized fashion when aggregate preferences are stable and mutual insurance against idiosyncratic risks is feasible. Consumption-based analysis of international risk sharing has been refined and extended by Stockman and Tesar (1990), Devereux, Gregory, and Smith (1992), Backus, Kehoe, and Kydland (1992), Backus and Smith (1992), and Baxter and Crucini (1993), among others. The basic message of this work is that correlations among international consumption movements are too low to be fully explained within a setting of free international asset trade and complete markets.

This paper draws on the consumption-based approach to develop an empirical framework for evaluating international financial integration. My framework recognizes explicitly that the ex post covariation in national consumptions depends not only on the freedom residents of different countries have to transact in securities markets, but also on the completeness of those markets, that is, on the range of contingencies on which cross-border contracts can be written. The empirical methodology can accommodate the two extreme cases of noncontingent asset trade and complete markets, as well as the broad middle ground where only a subset of national consumption risks is insurable. Once identifying assumptions are made, the consumption effects of noninsurable risks can be separated empirically from those of restrictions on international asset trade. (The necessary identifying assumptions are unlikely to be innocuous, however.)

Section 1 below sets out a model of international consumption comovements under possibly incomplete asset markets. Section 2 develops an econometric framework for testing the predictions of this model. The framework generalizes the one based on free international trade in bonds that has been applied to industrial countries by myself (1989) and by Kollmann (1992).

The data and estimation strategy are discussed in section 3, where a central question is the treatment of country-specific preference shocks. Section 4 tests successively less restrictive versions of the model. One test in section 4 is inspired by Feldstein and Horioka's (1980) analysis of economies' saving and investment rates, but leads to a different perspective on the postwar evolution of world capital markets. Section 5 concludes.

1. Market completeness and international consumption correlation

This section develops a general method for analyzing international consumption comovements when there is cross-border trade in assets. The approach illustrates that empirical predictions about consumption comovements depend not only on the opportunity to trade assets freely, but also on the range of events on which assets' payoffs can be conditioned. In the model I develop it is approximately true (and under one set of assumptions, exactly true) that ex post consumption-growth differentials between countries are uncorrelated with any random variable on which contingent contracts can be written. The model thus yields a potentially powerful method for discriminating empirically among different hypotheses about market completeness.

To simplify matters I assume that there is a single tradable consumption good and that each country i, $i = 1, \ldots, N$, is inhabited by a representative infinitely-lived individual.

Modeling the evolution of uncertainty is critical to the developments that follow. For every date t there is a set of possible states of nature $\mathcal{S}_t$, a generic element of which is denoted s_t. Transitions between states obey a Markovian probability law: the probability that state s_t occurs depends on the realized value s_{t-1} and possibly on calendar time. Conditional expectations are thus straightforward to compute.

Let C_{it} be the date-t consumption of the country-i individual. This individual's objective function at $t = 0$ is

$$(1)\quad U_0 = \mathrm{E}\left\{ \sum_{t=0}^{\infty} \beta_i^t u(C_{it}, \theta_{it}) \mid s_0 \right\}, \qquad 0 < \beta_i < 1,$$

where $\mathrm{E}\left\{\cdot \mid s_t\right\}$, given the Markov structure assumed, is an expectation conditional on the information observed up to time t. In (1), θ_{it} is a preference shock the realized value of which is one element determining the world economy's state.[2]

Let V_t be a minimal countable partition of $\mathcal{S}_t$ into *verifiable* events.[3] For any $v_t \in V_t$, contracts can be written on the event $s_t \in v_t$, but not on the event that s_t lies in some element of a partition of $\mathcal{S}_t$ strictly finer than V_t. I make no attempt in this paper to model the nonverifiability of some events. The notation $q(v_t \mid s_{t-j})$ will be used to denote the price on date $t-j$ of the asset that pays 1 consumption unit in the event $s_t \in v_t$ and 0 in the event $s_t \notin v_t$.

Predictions about consumption dynamics are derived from the Euler equations associated with transactions in these state-contingent assets. Let $C_{it} = C_i(s_t)$ be country i's per capita consumption level contingent on event s_t. The stochastic Euler equation associated with the asset described in the last paragraph is

$$(2) \quad q(v_{t+1}|s_t)u'[C_i(s_t),\theta_{it}] = \beta_i \mathrm{E}\left\{u'[C_i(s_{t+1}),\theta_{it+1}] \mid v_{t+1},s_t\right\}\cdot\pi(v_{t+1}|s_t),$$

where $\pi(v_{t+1}|s_t)$ is the date-t conditional probability that event v_{t+1} occurs. The left-hand side of this equation is the current utility cost of buying the state-contingent asset, the right-hand side the discounted expected utility value of its payoff.

If people in different countries i and j face the same asset prices and have rational expectations, then equation (2) implies that for all $s_t \in \mathcal{S}_t$, $v_{t+1} \in \mathcal{V}_{t+1}$,

$$(3) \quad \mathrm{E}\left\{\frac{\beta_i u'[C_i(s_{t+1}),\theta_{it+1}]}{u'[C_i(s_t),\theta_{it}]} - \frac{\beta_j u'[C_j(s_{t+1}),\theta_{jt+1}]}{u'[C_j(s_t),\theta_{jt}]} \mid v_{t+1},s_t\right\} = 0.$$

Equation (3) provides the central link between national intertemporal rates of substitution and insurable risks.

The main prediction of (3) is that ex post differences in individuals' marginal rates of intertemporal substitution are statistically uncorrelated with variables on which contractual payoffs can be conditioned, and with variables known as of date t. To set the stage for a proof, let $D_{ij}(s_{t+1},s_t)$ denote the ex post difference in marginal rates of intertemporal substitution between representative agents of countries i and j, so that (3) becomes:

$$\text{(4)}\quad \mathrm{E}\left\{D_{ij}(s_{t+1},s_t) \mid v_{t+1},s_t\right\} = 0 \qquad (\forall s_t \in \mathcal{S}_t,\ v_{t+1} \in V_{t+1}).$$

Since only events in V_{t+1} or countable unions thereof are verifiable, contingent contracts payable on date $t+1$ can be written only on random vectors $f:\mathcal{S}_{t+1} \to \mathbb{R}^n$ that are *measurable* with respect to V^*_{t+1}, the smallest set containing the null set $\varnothing$ and all countable unions of members of V_{t+1}. Measurability of f means that the inverse image $f^{-1}(I)$ of any product of half-open intervals $I \subseteq \mathbb{R}^n$ is a member of V^*_{t+1}, i.e., that the event $f(s_{t+1}) \in I$ is verifiable for all I. Measurability implies that $f(s_{t+1})$ is constant on every $v_{t+1} \in V_{t+1}$, since $f^{-1}(z) \in V^*_{t+1}$ for every point $z \in \mathbb{R}^n$. Similarly, variables known as of date t are functions on $\mathcal{S}_t$ that are measurable with respect to $\mathcal{S}^*_t$, which is defined analogously to V^*_t.[4]

To prove the main result, notice that for any V^*_{t+1}-measurable function $f:\mathcal{S}_{t+1} \to \mathbb{R}^n$,

$$\mathrm{E}\left\{D_{ij}(s_{t+1},s_t)f(s_{t+1})\right\} = \mathrm{E}\left\{\mathrm{E}\left\{D_{ij}(s_{t+1},s_t)f(s_{t+1}) \mid v_{t+1},s_t\right\}\right\}$$

$$= \mathrm{E}\left\{f(s_{t+1})\,\mathrm{E}\left\{D_{ij}(s_{t+1},s_t) \mid v_{t+1},s_t\right\}\right\} = 0.$$

The first equality follows from the law of iterated conditional expectations, the second from the constancy of $f(s_{t+1})$ on each element $v_{t+1} \in V_{t+1}$, and the third from (4). A similar argument shows that $D_{ij}(s_{t+1}, s_t)$ is uncorrelated with date-t information.

The discussion can be summarized by the following:

Theorem. *The date-t+1 ex post marginal rate of intertemporal substitution difference between any two countries i and j,*

$$(5) \quad D_{ij}(s_{t+1}, s_t) = \frac{\beta_i u'[C_i(s_{t+1}), \theta_{it+1}]}{u'[C_i(s_t), \theta_{it}]} - \frac{\beta_j u'[C_j(s_{t+1}), \theta_{jt+1}]}{u'[C_j(s_t), \theta_{jt}]},$$

is statistically uncorrelated with any random variable on which date-t+1 contracts can be written, as well as with any variables realized on date t or before.

Two simple examples will help to clarify this theorem's meaning:

1. As a first example, suppose that $V_t = \mathcal{S}_t$ on all dates, which is the case of *complete* markets. In this case contracts can be made contingent on *any* state of nature, so the theorem states that the random variable $D_{ij}(s_{t+1}, s_t)$ must be uncorrelated with any random variable realized at time t+1. This can be true, however, only if $D_{ij}(s_{t+1}, s_t)$ is a constant and, by (4), that constant must be zero. So by (5), for all states of nature,

$$(6) \quad \frac{\beta_i u'[C_i(s_{t+1}), \theta_{it+1}]}{u'[C_i(s_t), \theta_{it}]} = \frac{\beta_j u'[C_j(s_{t+1}), \theta_{jt+1}]}{u'[C_j(s_t), \theta_{jt}]}$$

in the case of complete markets. Marginal rates of intertemporal substitution must be equalized after the fact.[5]

2. As a second example consider the opposite extreme in which $V_t = \{\mathscr{S}_t\}$, i.e., in which the minimal verifiable partition of $\mathscr{S}_t$ consists of $\mathscr{S}_t$ alone. Now only noncontingent contracts can be written, so that the only assets traded are indexed bonds. The theorem above implies the weaker result that $D_{ij}(s_{t+1}, s_t)$ is uncorrelated with information available as of time t, or that

$$(7) \quad E\left\{\frac{\beta_i u'[C_i(s_{t+1}), \theta_{it+1}]}{u'[C_i(s_t), \theta_{it}]} - \frac{\beta_j u'[C_j(s_{t+1}), \theta_{jt+1}]}{u'[C_j(s_t), \theta_{jt}]} \mid s_t\right\} = 0$$

[compare with equation (3)]. Hence, the ex post rate-of-substitution differential is uncorrelated with information known as of date t. But its correlation with date-t+1 variables is unrestricted. Relations similar in spirit to (7) have been tested empirically by me (1989) and by Kollmann (1992)[6].

Intermediate between these two extreme possibilities is a range of cases in which partial insurance renders $D_{ij}(s_{t+1}, s_t)$ uncorrelated with some, but not all, date-t+1 variables. The intuition for the theorem is easy. Any idiosyncratic consumption risk systematically related to some verifiable random event is traded, leaving ex post differentials in marginal intertemporal substitution rates as functions of nonverifiable events only.

2. Econometric implications of the model

Empirical testing of the models presented in the last section requires additional identifying assumptions. Here I describe how restrictions on utility functions and on the distributions of preference shocks lead to simple econometric specifications of the models. These specifications, which I will apply to time-series data below, are related to specifications tested against panel microdata by Townsend (1989) and Mace (1991) and against cross-sectional microdata by Cochrane (1991).

Complete asset markets

Assume tentatively that there is free international trade in a *complete* set of Arrow-Debreu securities.

As a first possiblility, assume that the period utility function takes the isoelastic form

$$(8) \quad u(C_i, \theta_i) = \frac{1}{1-\rho} (C_i)^{1-\rho} \exp(\theta_i) \qquad (\rho > 0).$$

Let $t = 0$ be the initial period and let θ_{i0} be normalized, for all countries i, so that $\theta_{i0} = 0$. Under complete markets equation (6) holds true; it implies that $\forall t \geq 0$,

$$(9) \quad \frac{\beta_i^t \exp(\theta_{it}) C_{it}^{-\rho}}{C_{i0}^{-\rho}} = \frac{\beta_j^t \exp(\theta_{jt}) C_{jt}^{-\rho}}{C_{j0}^{-\rho}}.$$

The assumption that countries share a common risk-aversion coefficient ρ is not innocuous, but is a central maintained hypothesis in the analysis and tests that follow. In my 1989 paper I found little evidence against this hypothesis in quarterly 1973-1985 data for Germany, Japan, and the United States.[7]

Taking natural logarithms in (9) yields the time-series model

$$(10)\quad \log C_{it} = \log C_{jt} + \log(C_{i0}/C_{j0}) + \log(\beta_i/\beta_j)(t/\rho) + \frac{1}{\rho}(\theta_{it} - \theta_{jt}).$$

In (10), $\log(\beta_i/\beta_j)$ measures the extent to which country i's residents are more patient than those of country j, while $\log(C_{i0}/C_{j0})$ reflects relative impatience as well as the initial wealth of i relative to j.

A main implication of (10) is that when national time-preference rates coincide ($\beta_i = \beta_j$) and there are no differential preference shocks across countries ($\theta_{it} - \theta_{jt} = 0$, $\forall t$), national per capita consumption levels display equal proportional ex post comovements. In the analysis below, however, country-specific taste shocks will play a role. Equation (10) then makes the weaker prediction that *other things the same,* $\log C_{it}$ *and* $\log C_{jt}$ *should move by equal amounts.* Equation (10) also implies that *no date-t variable that is uncorrelated with* $\theta_{it} - \theta_{jt}$ *will be correlated with* $\log C_{it} - \log C_{jt}$.

In the many-country framework of this paper, an alternative estimation strategy has some potential advantages that section 3 will discuss in detail. Let n_{it} be country i's share in world population and let C_{Wt} be world per capita consumption, so that

$$(11)\quad C_{Wt} \equiv \sum_{j=1}^{N} n_{jt}C_{jt}.$$

Let μ_t be the common value of the marginal rates of substitution in equation (9). Using (9) and (11), one finds that

$$\mu_t = \left[\frac{C_{Wt}}{\sum_j \beta_j^{t/\rho} \exp(\theta_{jt}/\rho) n_{jt} C_{j0}}\right]^{-\rho},$$

from which it follows that

$$(12) \quad \log C_{it} = \log C_{Wt} + \log C_{i0} + (\log\beta_i)(t/\rho)$$
$$+ \left\{\theta_{it}/\rho - \log[\textstyle\sum_j \beta_j^{t/\rho} \exp(\theta_{jt}/\rho) n_{jt} C_{j0}]\right\}.$$

(Notice that the consumption time trend is zero when $\beta_i = \beta_j$, $\forall i,j$, and when countries' population shares contain no time trend.) Equation (12), which provides an alternative and more compact mode of summarizing the main message of equation (10), likewise implies proportional movements between each country's consumption and world consumption, all else equal. It also implies that date-t variables independent of preference and population shocks will also be uncorrelated with $\log C_{it} - \log C_{Wt}$.

The isoelastic-utility specification (8) is an appropriate one in a context of ongoing economic growth, and the log-consumption specification in equations (10) and (12) will therefore be the basis for the tests carried out below. An alternative, exponential form of the period utility function helps simplify some of the econometric arguments I make in section 3 in favor of a testing strategy based on (12) rather than (10). It is therefore useful to develop briefly the empirical implications of complete asset markets under exponential utility,

$$(13) \quad u(C_i, \theta_i) = -\exp(-\rho C_i + \theta_i)/\rho \qquad (\rho > 0).$$

If we assume that $\theta_{i0} = 0$, $\forall i$, equation (6) now implies that

$$\frac{\beta_i^t \exp(-\rho C_{it} + \theta_{it})}{\exp(-\rho C_{i0})} = \frac{\beta_j^t \exp(-\rho C_{jt} + \theta_{jt})}{\exp(-\rho C_{j0})}$$

$\forall t \geq 0$. Taking natural logarithms yields

$$(14) \quad C_{it} = C_{jt} + (C_{i0} - C_{j0}) + \log(\beta_i/\beta_j)(t/\rho) + \frac{1}{\rho}(\theta_{it} - \theta_{jt}),$$

the "levels" version of (10). The analogue of (12) is

$$(15) \quad C_{it} = C_{Wt} + (C_{i0} - \Sigma_j\, n_{jt}C_{j0}) + (\log\beta_i - \Sigma_j\, n_{jt}\log\beta_j)(t/\rho) + (\theta_{it} - \Sigma_j\, n_{jt}\theta_{jt}).$$

As noted above, the results I report below are for equations involving consumption logs, not consumption levels. Equations estimated in levels, however, led to very similar results.

Incomplete asset markets

This section has proceeded under the tentative assumption of complete asset markets. If consumption depends on idiosyncratic uninsured risks, however, equation (4) shows that equations like (10), (12), (14), and (15) must be modified by the addition of extra error terms reflecting those risks.

In the extreme case that only noncontingent assets are traded on date $t-1$, the extra error term can reflect any new date-t information relevant to current consumption decisions. Generally, however, at least some state-contingent assets are traded. Their payoffs are functions of events that do not generate ex post international differences in the growth of the discounted marginal utility of consumption.

The econometric implications of market incompleteness are discussed further in context below [equations (19) and (20)].

3. Specification and data

There are two preliminary specification issues to be settled before estimation. First, should one investigate pairwise regressions such as (10) and (14), or are there advantages to working with a world consumption measure as in (12) and (15)? Second, should consumption consumption data be differenced prior to regression? This section discusses these two issues and then describes the data.

Reducing least-squares bias through use of world consumption data

Several studies have attempted to test relations like (10) and (14). Generally these studies assume that $\beta_i = \beta_j$ and that preference shocks are zero, and then proceed to examinine pairwise correlations between C_{it} and C_{jt} (or between transforms of those variables). These correlations turn out to be low in many cases--generally lower, for industrialized countries, than the correlations between national output levels.[8] The finding of low pairwise consumption correlation is often taken as evidence of imperfect international financial-market integration or of missing markets.

Such low international consumption correlations are not surprising, even under complete markets, when there are significant country-specific preference shocks. Yet regression equations such as (10) or (14) can mask the possibility that international consumption changes due to factors other than preference shifts are closely synchronized. One way to think of this problem is as an *endogenous-regressor* problem: country-j consumption in (10) or (14) is likely to be positively correlated with θ_{jt}--a high realization of θ_{jt} raises the marginal utility of country j's time-t consumption--and so least-squares estimates will tend to produce downward-biased estimates of slope coefficients.

One can reduce this bias by estimating equations of form (12) or (15), in which world consumption is the independent variable explaining country i's consumption. Particularly if country i is small, the degree to which its taste shock θ_i diverges from an average world taste shock should be approximately uncorrelated with world consumption. The composite errors $\theta_{it}/\rho - \log[\sum_j \beta_j^{t/\rho}\exp(\theta_{jt}/\rho)n_{jt}C_{j0}]$ in (12) and $\theta_{it} - \sum_j n_{it}\theta_{it}$ in (15) are thus more plausibly weakly correlated or uncorrelated with their respective regressors than is the error $(\theta_{it} - \theta_{jt})/\rho$ in (10) and (14).

An example based on exponential utility clarifies this intuition. Imagine a pure exchange economy in which world per capita output on date t, Y_{Wt}, is an exogenous random variable. Under complete markets either of (14) or (15) describes the equilibrium consumption allocation; if we simplify the notation by assuming that $\rho = 1$, that $C_{i0} = C_{j0}$ and $\beta_i = \beta_j$ $\forall i,j$, and that $n_{it} = 1/N$ $\forall i,t$, then these two equations become, respectively,

$$(16) \quad C_{it} = C_{jt} + \theta_{it} - \theta_{jt},$$

$$(17) \quad C_{it} = C_{Wt} + \theta_{it} - (1/N)\sum_j \theta_{jt}.$$

Since in equilibrium $(1/N)\sum_j C_{jt} = C_{Wt} = Y_{Wt}$, country i's equilibrium consumption level is, by (17),

$$(18) \quad C_{it} = Y_{Wt} + \theta_{it} - (1/N)\sum_j \theta_{jt}.$$

Let $\hat{\alpha}_{ij}$ be the slope estimate derived from applying ordinary least squares to the pairwise regression equation (16). I make the further simplifying assumptions that the taste shocks θ_{it} are distributed independently of Y_{Wt} and that $\forall i,j$, θ_{it} and θ_{jt} have identical but independent distributions. By (18), it follows that

$$\text{plim } \hat{\alpha}_{ij} = 1 - (\sigma_\theta^2/\sigma_C^2),$$

where σ_θ^2 is the variance of preference shocks and σ_C^2 the variance of national consumption levels. If preference shocks account for part of the overall variance of national consumption, least-squares regressions of country-i on country-j consumption can produce slope estimates that are asymptotically biased below the true value of 1.

Consider next the least-squares slope estimate $\hat{\alpha}_{iW}$ from (17). Under the distributional assumptions just made, plim $\hat{\alpha}_{iW} = 1$. Least-squares estimation of (17) thus gives an asymptotically accurate picture of how national consumptions and world consumption covary holding preferences constant. Furthermore, date-t variables uncorrelated with the error $\theta_{it} - (1/N)\sum_j\theta_{jt}$, other than C_{Wt}, should be insignificant in (17) when it is estimated by least squares.

In section 4 I will try to reduce least-squares bias by relying on a specification like (17), in which world per capita consumption is a regressor. In situations other than the simple one-good pure-exchange economy of my example, however, some bias can remain. Preference shocks may alter the division of world output between consumption and saving, so in principle a nonzero correlation between the regressor and error term in (17) is possible in an economy with investment. This possibility is most important when country i in (17) is large, a point I will revisit in analyzing results for the United States below.

Should the data be differenced?

As mentioned above, the estimates reported below will be derived from the logarithmic specification (12). Tests could in principle be based on equations such as (12) itself or on the implied equations in log-differences. In their microdata studies, Mace (1991) and Cochrane (1991) use differenced specifications to remove fixed household effects corresponding to the terms in period-0 consumption in (12).[9] In the present time-series context another reason for considering a differenced model is the danger of spurious correlations and asymptotically invalid inferences.

Per capita consumption data are well known to be generated by integrated or near-integrated processes, a feature rationalized by forward-looking consumption theories. A regression of country i consumption on world consumption, as in equation (12), might give a misleading impression of close correlation if these series are not cointegrated. Such a spurious relationship could also result in erroneous statistical inferences.

For most of the countries examined below, logarithmic regressions of national on world consumption do give rise to R^2 statistics that are above Durbin-Watson statistics--the informal diagnostic indicator of spurious regression suggested by Granger and Newbold (1974). More formally, for a typical country i, $\log C_{it}$, $\log C_{Wt}$, and $\log C_{it} - \log C_{Wt}$ all appear to be nonstationary processes; indeed, $\log C_{it}$ and $\log C_{Wt}$ are often not cointegrated.[10] These findings suggest that a specification in log-differences will be more informative than one in log-levels.

Taking a linear approximation to (12) (which assumes complete markets) and differencing yields

$$(19) \quad \Delta\log C_{it} = \delta + \Delta\log C_{Wt} + \varepsilon_{it};$$

the disturbance ε_{it}, which is assumed to follow a stationary process, is a function of taste shocks and, possibly, errors in measuring consumption.[11] I will generally assume that ε_{it} and $\Delta\log C_{Wt}$ are approximately uncorrelated, but will also remark on cases where some correlation seems likely.[12]

Equation (19) may lead to unbiased least-squares slope estimates even when asset markets are incomplete. To understand this possibility, notice that in the present context and under incomplete markets, equation (19) would become

$$(20)\quad \Delta\log C_{it} = \delta + \Delta\log C_{Wt} + \varepsilon_{it} + \eta_{it},$$

where η_{it} is a function of date-t innovations that are not verifiable and thus cannot be insured. If such innovations are uncorrelated with $\Delta\log C_{Wt}$, least-squares estimation of (20) gives a consistent slope estimate provided $E(\Delta\log C_{Wt}\varepsilon_{it}) = 0$. The converse of this implication should also be noted, however: if markets are incomplete and the uninsurable factors η_{it} are correlated with $\Delta\log C_{Wt}$, then the least-squares slope estimator is not consistent for (20).

Equation (20) will be the workhorse for the empirical analysis in section 4. A test of financial integration asks if the coefficient of $\Delta\log C_{Wt}$ in (20) is 1 once the uninsurable risk factors underlying η_{it} are added as regressors (assuming these, like $\Delta\log C_{Wt}$, are uncorrelated with ε_{it}). Conversely, random variables uncorrelated with ε_{it}, and on which contracts cannot be written, should enter significantly into (20). It is thus possible in principle to identify the uninsured factors contributing to idiosyncratic national consumption fluctuations.

Data description

The annual national income and product account and population data used in this study come from the Penn World Table (Mark 5), as described by Summers and Heston (1991). The national-account components studied below--gross domestic product (*GDP*), consumption (*C*), private plus public investment (*I*), and government consumption (*G*)--are all measured in real per capita terms at 1985 international prices.[13]

World per capita consumption is as defined in equation (11). For my purposes the "world" consists of the 47 Penn World Table countries with data extending over the entire 1950-1988 sample, and awarded a quality grade of at least C– by Summers and Heston.[14] The major oil exporters are not part of this group.

To allow convenient comparison with the findings in other studies of international consumption comovements, I report in table 1 correlation coefficients for changes in the logarithms of annual national per capita consumption rates. Each box in the table contains two estimated correlation coefficients one (above the diagonal) for the period 1951-1972 and a second (below the diagonal) for the period 1973-1988. The sample split is motivated by independent evidence that the first subperiod was on the whole an era of considerably lower global asset-market integration than the second. The individual-country sample is the Group of Seven (G-7), consisting of the largest industrial nations. Obviously it is feasible and desirable to apply tests such as those done here to additional countries.

The first row of Table 1 shows the correlation coefficient between the change in each G-7 country's log consumption per capita and the change in the rest of the world's log consumption per capita.[15]

Three main facts are apparent from the table. First, as the recent empirical literature has shown, pairwise correlations between national consumption growth rates, as well as correlations between national growth rates and rest-of-world growth rates, are typically far below the unit correlation that would characterize a world with costless asset and commodity trade, complete markets, no preference shocks, and no errors in measuring real consumption per head. Second, in most cases the country-to-country correlation coefficients for the second sample subperiod are higher than those for the first (most of the exceptions involve Canada). Third, for all countries except Canada, the correlation of domestic with world consumption growth rises in the second sample period--dramatically so in the cases of Germany and Japan.

Taken as a whole, Table 1 is consistent with the hypothesis that increased international trade in a broader range of financial assets took place after 1973. Table 2 offers some additional evidence pertinent to this question. Financial diversification makes it feasible for every country to reduce the variability of its consumption growth relative to that of world average consumption growth. Table 2 shows that only Japan and Germany have done so in a big way. World consumption growth became a third more variable after 1973, but Germany actually reduced the absolute variability of its consumption growth rate while Japan held the variability of its own about constant. Canada is again an outlier, with a massive increase in relative consumption-growth variability.

A potential alternative explanation for the results in Tables 1 and 2, one that does not rely on international diversification, comes from looking at changes over time in the behavior of national per capita *outputs* (that is, gross domestic products). The results indicate that Tables 1 and 2 are easily explained by a model in which consumption growth closely (and naively) tracks domestic output growth.

Table 3 shows the output correlations corresponding to the entries in Table 1. As in Table 1, the correlations have a tendency to increase over time. Furthermore, as noted by Backus, Kehoe, and Kydland (1992) and Stockman and Tesar (1990), the international output correlations in Table 3 tend to be higher than the consumption correlations in Table 1.

Table 4 does the calculations in Table 2 using output rather than consumption growth rates. For all the countries but the U.K., there is a decline in domestic relative to world output-growth variability after 1973. For Japan and Germany in particular, the declines in domestic relative to world output-growth variability match the corresponding declines in relative consumption-growth variability shown in Table 2.

Tables 2 and 4 reinforce the finding that even though world consumption growth is smoother than world output growth, the industrial countries' output-growth risks appear better "diversified" than their consumption-growth risks. This fact leads to a fundamental ambiguity, since the data do not obviously refute the view that apparent changes over time in international consumption correlations are entirely due to exogenous shifts in output correlations rather than improved risk sharing.

Table 5 presents some relevant additional information. The table shows correlation coefficients between domestic per capita consumption growth and the rest of the world's per capita output growth. Comparing this table with Table 1, one sees that for all countries save Italy and the United States, the post-1973 correlation of domestic consumption growth with rest-of-world output growth is below the corresponding correlation with rest-of-world consumption growth (though sometimes only barely so). This is weak evidence that for some countries more may be going on in the data than a simple proportionality of output and consumption.

The implications of these data for global financial markets depend critically on the importance of preference shocks and uninsured risks. The empirical model developed above provides a framework within which the various factors generating the changes in Tables 1 through 5 can potentially be identified. I therefore turn to tests of that model's predictions.

4. Empirical results

The model developed above has a number of empirical implications. This section reports tests of the model based on equation (19), which assumes complete markets, and on the less restrictive equation (20), which allows for some uninsurable risks.

National consumptions and world consumption

The first application of the model is to estimate directly equations of the form suggested by (19),

$$\Delta \log C_{it} = \delta + \alpha_{iW} \Delta \log C_{Wt} + \varepsilon_{it},$$

and test the hypothesis that $\alpha_{iW} = 1$. Table 6 reports the results of least-squares estimation over the 1951-1972 (panel A) and 1973-1988 (panel B) subsamples.[16]

For countries other than Canada and the United States, the coefficient on world consumption growth rises, usually sharply, in the second subperiod. (Canada's behavior is consistent with the results of Table 1; however, the big increase in the variability of Canada's consumption growth after 1973 lowers the precision of estimation.) The equation's $\bar{R}^2$ rises for the countries outside North America, also suggesting a greater coherence between domestic and world consumption growth.

For 1951-1972, the hypothesis $\alpha_{iW} = 1$ can be rejected at the 5% level only for Italy and the U.S., but this is in most cases the result of low-precision estimates $\hat{\alpha}_{iW}$, not of estimates near 1 (panel A). Thus, the hypothesis $\alpha_{iW} = 0$ also is not rejected at the 5 per cent level for three of the seven over 1951-1972. Over 1973-1988 (panel B), however, $\alpha_{iW} = 1$ is rejected only for the U.S. despite more precise estimates of α_{iW} for Germany and Japan. The hypothesis $\alpha_{iW} = 0$ is now rejected in five of seven cases.

The U.S. estimates pose a special problem in light of the country's size. Because U.S. consumption makes up a sizable fraction of world consumption, positive realizations of the U.S. preference shock θ_{US} are likely to have a large positive correlation with the world consumption measure used in Table 2; as a result, the least-squares slope estimate in Table 2 probably has an upward bias. In fact, the most likely explanation for the fall in that estimate between the first and second subperiods is the shrinking weight of U.S. consumption in world consumption.

To a lesser degree, this problem could plague the non-U.S. equations as well: in a world with investment, we may exaggerate the link between a given country's consumption and world consumption when we do not remove the country from the world consumption index. Let C^i_{Wt} be world per capita consumption outside country i. Table 7 shows the results of regressing

$\Delta \log C_{it}$ against $\Delta \log C^{i}_{Wt}$ for each of the G-7 countries.[17]

All slope coefficients and all but one of the $\bar{R}^2$s drop, but the substance of the results changes little. Comparing Table 7 with table 6, the results seem in general somewhat less compatible with global financial integration in panel A, but not more so in panel B. For the U.S. α_{iW} is no longer estimated to be significantly above 1 in either subperiod, but neither panel's estimate differs significantly from 0 (at the 5 percent level) either. Furthermore, regardless of period the rest-of-world consumption growth rate accounts for the same very small fraction of the variation in the U.S. rate. The post-1973 results for Italy are ambiguous, those for Canada even less decisive.

Notice that by changing the regressor in Table 7, an opposite bias may be introduced, one especially relevant for large countries like the U.S. Positive realizations of the preference shock θ_{US} are likely to lower rest-of-world consumption growth and thus to lead to *downward*-biased slope estimates. Without more information, it is impossible to know how large this bias is, or the extent to which it affects the other G-7 countries.

A concern raised by the data description above (Section 3) is that the results in Table 7 reflect nothing more than the typical high correlation between domestic consumption growth and domestic output growth, coupled with the typical high correlation between domestic output growth and world output growth. To address this concern, Table 8 reports the results of estimating

$$\Delta \log C_{it} = \delta + \alpha_{iW} \Delta \log C^{i}_{Wt} + \gamma_i \Delta \log GDP^{i}_{Wt} + \varepsilon_{it},$$

where GDP^{i}_{Wt} is world per capita output outside country i. The right-hand variables in this equation are quite collinear, so sharp conclusions are not expected. Nonetheless, the estimates suggest that for the G-7 countries other than Italy and the United States, it is world consumption growth rather than world output growth that was more closely related to domestic consumption growth after 1973. For France, Germany, and Japan, the reversal of this relationship between the two sample periods is noteworthy. The results in Table 8 are consistent with the simple correlations in Tables 1 and 5, and provide weak evidence that the patterns in the data are not driven entirely by changing output correlations as opposed to improved international risk sharing.

The role of oil-price shocks

For four of the G-7 countries, the rest of the world's consumption growth appears to play a statistically significant and economically important role in explaining domestic consumption growth after 1973. Table 8 notwithstanding, it still is possible that this finding is not due to international asset-market integration at all, but is the result of common shocks to the world macroeconomy that hit all industrialized economies simultaneously and with similar effects on consumption growth. Over the 1973-1988 sample period, a leading probable source of such common shocks is the real price of petroleum. The simple correlation coefficient over the period between the change in the log real price of oil and the change in the log of world real per capita consumption is -0.6.[18]

To explore this possiblity I add the change in the log real oil price between years t and $t-1$, ΔOIL_t, to the basic estimating equation:

$$\Delta \log C_{it} = \delta + \alpha_{iW}\Delta \log C^{i}_{Wt} + \gamma_i \Delta OIL_t + \varepsilon_{it}.$$

If the countries making up the world consumption index I use optimally insured each other against the idiosyncratic effects of oil-price shocks, then $\gamma_i = 0$ holds because oil prices affect an individual country's consumption only by affecting group consumption (recall the theorem in Section 1); otherwise $\gamma_i \neq 0$ in general.[19] If the results in Tables 6 and 7 are *entirely* due to the common effect of oil prices on group consumption, but idiosyncratic risks have not been shared within the group, then α_{iW} should become insignificant with ΔOIL added to the regression.

Table 9 reports the estimation results for 1973-1988. The oil variable enters significantly in the regressions for Italy, the U.K., and the U.S., suggesting that these countries did not fully trade to the rest of the 47-country world sample the idiosyncratic consumption risk due to oil-price changes. For all of the countries but Canada and the U.S., however, α_{iW} is now estimated to be fairly close to 1; it is significantly different from 0 at the 5 percent level for Germany, Italy, and Japan, and at the 10 percent level for France. The estimate $\hat{\alpha}_{iW}$ is not significantly different from 1 for any country.

Modeling imperfect allocation

A simple heuristic model of international asset-market inefficiency produces a more stringent test of the hypothesis that world financial-market integration increased during the period after 1973.

The quantity $TR_i \equiv C_i - (GDP_i - I_i - G_i)$ measures the net resource transfer from the rest of the world to country i due to foreign borrowing, interest/dividend earnings and capital gains on assets held abroad, and all other state-contingent payments on foreign wealth. Of course, $TR_i = 0$ when international capital markets are closed. I define the *domestic resource limit*, DRL_i, as

$$(21) \quad DRL_i \equiv GDP_i - I_i - G_i,$$

i.e., as the consumption level at $TR_i = 0$ given GDP_i, I_i, and G_i.

Let $C_i^\star$, $I_i^\star$, and $TR_i^\star$ be the hypothetical consumption, investment, and net resource transfer levels under free asset trade. To simplify I will suppose that in the short run GDP_i and G_i do not depend on the extent of trade, but that actual date-t investment is related to potential investment by

$$(22) \quad I_{it} = I_{it}^\star - \kappa(TR_{it}^\star - TR_{it}) \qquad (0 \le \kappa \le 1).$$

Assume next that actual transfers are given by

$$(23) \quad TR_{it} = \lambda TR_{it}^\star + \zeta_{it} \qquad (0 \le \lambda \le 1),$$

where ζ_{it} is an exogenous mean-zero disturbance. Combine the definition of TR_i, the assumption $TR_i^\star = C_i^\star - (GDP_i - I_i^\star - G_i)$, (21), (22), and (23). Apart from an error term, actual date-t consumption is a weighted average of $C_{it}^\star$ and $GDP_{it} - I_{it} - G_{it}$:

$$(24)\quad C_{it} = \frac{\lambda}{1 - \kappa(1-\lambda)} C^*_{it} + \frac{(1-\lambda)(1-\kappa)}{1 - \kappa(1-\lambda)} DRL_{it} + \frac{(1-\kappa)}{1 - \kappa(1-\lambda)} \zeta_{it}.$$

If $\lambda = 1$ consumption is at its efficient level (apart from the error term). If $\kappa = 1$ investment bears all the burden of any fall in net resource transfers, so consumption need not differ from C^*_{it}. If $\lambda = 0$ consumption equals DRL_i (apart from the fraction of ζ_i that does not go into home investment). Other cases, however, imply that both C^*_{it} and DRL_{it} will systematically affect consumption, with positive partial derivatives that sum to 1.

Now suppose that equation (19) characterizes the free-trade level of consumption and that (24) can be expressed in log-differences. The resulting equation is

$$(25)\quad \Delta\log C_{it} = \delta + \alpha_{iW}\Delta\log C_{Wt} + \gamma_i\Delta\log DRL_{it} + \nu_{it},$$

where ν_{it} is a linear combination of the preference shock ε_{it} from (19) and the net resource transfer shock ζ_{it} from (23). In estimating (25) we would expect to find that $\alpha_{iW} = 0$ and $\gamma_i = 1$ under a regime of limited global financial integration. Under high financial integration, however, we would expect that $\alpha_{iW} = 1$ and $\gamma_i = 0$.

The regression framework (25) is closely related to one developed by Feldstein and Horioka (1980) for estimation of the cross-sectional correlation between saving and investment. Intuitively, equation (25) gives an indication of whether domestic consumption growth is more closely correlated with global or with domestic factors. If domestic investment is constrained by domestic saving, then domestic consumption is constrained by the domestic resource limit and the hypothesis $\alpha_{iW} = 0$, $\gamma_i = 1$ should not be rejected. An advantage of the present framework is that it avoids the use of national income and product account data on national saving which (among other problems) fail adequately to measure the international asset-income flows central to the present inquiry.[20]

Table 10 presents estimates of equation (25). I used the variable ΔOIL in some of the regressions to control for associated uninsured risks.[21]

In panel A the hypothesis that $\alpha_{iW} = 1$ and $\gamma_i = 0$ is rejected at a very low significance level every time. Only for the U.S. is it possible to reject the hypothesis that $\alpha_{iW} = 0$ and $\gamma_i = 1$, but the reason is a coefficent on $\Delta \log C_W^i$ that is significantly *negative*. Only in that case, and in the case of Canada, is the latter coefficient estimated at far from 0. In contrast, all coefficients on $\Delta \log DRL_i$ (with France a marginal exception) are insignificantly different from 1. The picture that emerges for the years 1951-1972 is one of an industrialized world in which financial markets essentially provide no consumption insurance.

While panel B falls short of portraying the opposite extreme of full financial integration, its results are quite different from those of panel A (perhaps surprisingly so, in view of the Feldstein-Horioka findings). For four countries--France, Germany, Italy, and Japan--the hypothesis $\alpha_{iW} = 1$ and $\gamma_i = 0$ cannot now be rejected; for France, Germany, and Japan, the hypothesis $\alpha_{iW} = 0$ and $\gamma_i = 1$ is rejected decisively (and it fails at the 11 percent level for Italy). Germany stands out as showing most strongly the characteristics we would expect of an economy well integrated into world financial markets. Because Japan maintained capital controls until the start of the 1980s while France and Italy did so until past the middle of that decade, this result is plausible.

The United Kingdom's appearance of financial insularity may be due to its own controls on resident capital movements, which were dismantled only in 1979. If equation (25) is estimated for the U.K. over 1979-1988, the result (with the intercept suppressed) is:

$$\Delta\log C_{UK} = \underset{(1.03)}{1.45}\ \Delta\log C_{W}^{UK} + \underset{(0.45)}{0.40}\ \Delta\log DRL_{UK} - \underset{(0.03)}{0.02}\ \Delta OIL.$$

Neither hypothesis, $\alpha_{UK,W} = 0$ and $\gamma_{UK} = 1$, nor $\alpha_{UK,W} = 1$ and $\gamma_{UK} = 0$, can be rejected; but insofar as one can draw conslusions from only 10 observations, the results above seem more compatible with international financial integration of the U.K. than do those in table 10, panel B.

The U.S results in panel B may be due to strong negative correlation between world consumption growth and the residual in the U.S. equation. The results for Canada are a mystery, especially in view of other, independent evidence suggesting a high degree of openness for Canadian financial markets.[22]

A closer look at Germany and Japan

Having come this far, it is tempting to carry out further tests on the diversification of idiosyncratic macroeconomic shocks. For example, does consumption growth respond to idiosyncratic output risk or can such risk in large part be traded away? I will argue in this section, using Germany and Japan as examples, that severe endogeneity problems prevent such tests from giving unambiguous answers. The argument suggests that the econometric results of the last section are potentially consistent with contradictory structural interpretations.

Table 11 presents 1973-1988 regressions of German consumption growth on world consumption growth and key domestic macroeconomic variables: changes in output, total investment, and government consumption. Since output is in part a function of possibly unobservable effort, we would expect some income components to be uninsurable, as the micro-level studies of Mace (1991) and Cochrane (1991) confirm. Changes in investment profitability could widen any wedge between domestic and rest-of-world consumption growth if world savings cannot flow costlessly to their most productive uses. Finally, considerations of moral hazard make it implausible that government spending shocks are completely insurable abroad: such insurance would present governments with an irresistible incentive to overspend. If uninsured idiosyncratic consumption risks are uncorrelated with aggregate preference shocks, they should enter significantly into the Table 11 regressions.

Table 11 strongly supports the basic model of financial integration for Germany, but suggests that government-consumption shocks are not fully insurable abroad. Regression 3, for example, strongly rejects any hypothesis setting the coefficient on government consumption growth to zero. The results show that domestic and rest-of-world consumption move in proportion except for shocks to German government consumption, which actually raise domestic growth relative to world growth, contrary to the prediction of a neoclassical Ricardian model of purely wasteful government spending. Output and investment shocks, however, seem to play no role. It should be noted that the output and investment variables are highly correlated with rest-of-world consumption growth (the simple correlation coefficients are 0.84 and 0.77, respectively).

Is it possible that output shocks really do contribute to "excess" domestic consumption growth in regressions 1, 4, 5, and 7, but that their effect is masked by a correlation with the preference shocks in the equation disturbance? This seems implausible, as the correlation between preference shocks that raise home consumption and output would likely have to be negative to bias downward the coefficient of $\Delta \log GDP$.

The results for Japan in Table 12 present a quite different picture. Here domestic GDP growth is significant; moreover, its presence reduces the influence of rest-of-world consumption growth to zero. Over 1973-1988, Japanese output growth and rest-of-world consumption growth are highly correlated (the correlation coefficient is 0.72). One interpretation of our earlier results suggesting substantial financial integration for Japan is that world consumption growth was merely proxying the true factor driving Japan's consumption growth, namely, the country's own domestic output growth. On this view, the Japan regressions in Tables 6 through 10 are not strong evidence in support of a financial-market link between Japanese and foreign consumption growth.[23]

This interpretation of Table 12 relies, however, on an assumption that $\Delta\log GDP$ is uncorrelated with the unobservable Japanese preference shocks. A different possible interpretation is suggested, however, by the hypothesis that preference shocks that raise Japanese consumption growth also raise Japanese GDP growth (through Keynesian or other mechanisms).

To investigate the effects of such a correlation, write regression 1 in Table 12 (for example) as

$$\Delta\log C_{JN,t} = \alpha_{JN,W}\, \Delta\log C^{JN}_{W,t} + \gamma\, \Delta\log GDP_t + \varepsilon_{JN,t},$$

where $\varepsilon_{JN,t}$ is a pure relative-preference shock. Suppose that $\Delta\log GDP_t$ and $\varepsilon_{JN,t}$ have covariance $\sigma_{2\varepsilon}$, that ρ_{12} is the correlation coefficient between $\Delta\log C^{JN}_{W,t}$ and $\Delta\log GDP_t$, and that these two regressors have standard deviations σ_1 and σ_2, respectively. If $\sigma_{2\varepsilon} > 0$, the least-squares estimate $\hat{\gamma}$ of γ is upward biased, and at the same time $\hat{\alpha}_{JN,W}$ is a *downward*-biased estimate of $\alpha_{JN,W}$ if $\rho_{12} > 0$:

$$(26)\quad \text{plim}\ \hat{\gamma} = \gamma + \frac{\sigma_{2\varepsilon}}{\sigma_2^2(1-\rho_{12}^2)},$$

$$(27)\quad \text{plim}\ \hat{\alpha}_{JN,W} = \alpha_{JN,W} - \frac{\rho_{12}\sigma_{2\varepsilon}}{\sigma_1\sigma_2(1-\rho_{12}^2)}.$$

So when both $\sigma_{2\varepsilon}$ and ρ_{12} are positive, it is theoretically possible that $\alpha_{JN,W} = 1$ notwithstanding a large-sample regression like 1 in Table 12 with $\hat{\alpha}_{JN,W} = -0.01$.

Formulas (26) and (27) aid in judging the plausibility of so large an asymptotic bias. Imagine that the true value of $\alpha_{JN,W}$ is 1. Since $\sigma_1 = 0.011$, $\sigma_2 = 0.024$, and (as just noted) $\rho_{12} = 0.72$, (27) shows that the asymptotic value of $\hat{\alpha}_{JN,W}$ will be -0.01 when $\alpha_{JN,W} = 1$ only if $\sigma_{2\varepsilon} = 1.78 \times 10^{-4}$. To get a feel for the implications of this covariance number, note that it would result from assuming (for example) that the standard deviation of preference shocks is half that of output growth and that the correlation coefficient between output growth and preference shocks is 0.62. As one would expect, preference shocks must be large and highly correlated with output to create sufficient least-squares bias, but perhaps not unreasonably so. (For comparison, the standard deviation of annual Japanese consumption growth is 0.023 for 1973-1988).

Given $\sigma_{2\varepsilon} = 1.78 \times 10^{-4}$ and an asymptotic estimate $\hat{\gamma} = 0.82$, however, (26) shows that the true value of γ would be only 0.18. The results in Table 12 thus are not totally irreconcilable with a model in which output shocks are largely diversified abroad. Unfortunately, the validity of this reconciliation is hard to evaluate because it is based on the assumed properties of an unobservable preference shifter.

The discussion illustrates the identification problems bedeviling attempts to measure international financial integration and market completeness using aggregate data. The results of this section seem uniformly consistent with the proposition that the German economy is tightly meshed into international financial markets: its consumption moves in proportion to the rest of the world's, and appears not to rise more quickly when domestic output is high or more slowly when domestic investment is high. German government consumption does, however, have an idiosyncratic positive effect on German private consumption. Japan, which liberalized its financial markets more recently than did Germany, shows ambiguous evidence of financial openness. Only under much stronger identifying assumptions than those invoked in the German case can Table 12's results be made consistent with the hypothesis that Japan is as well integrated into world financial markets as Germany. If these assumptions are false, explanations other than increasing financial integration must be found for the post-1973 fall in Japan's relative consumption-growth variability (Table 2).

5. Concluding remarks

This paper has studied the relationship between domestic consumption growth and world consumption growth for the G-7 industrial countries. For most of these countries there appears to be a postwar trend of increasing coherence between domestic and world consumption growth, as predicted by models of international financial integration. But the correlation between those variables remains far from perfect--as one would expect, even in a world of unrestricted international asset trade, when asset markets are incomplete, national preferences are subject to shocks, and consumption is measured with error.

Another set of factors underlying empirical international consumption correlations has not been discussed in this paper: nontraded goods and services, including leisure. These factors' influences were impounded into the error terms of my econometric equations, but an attempt to measure and model them explicitly is an obvious next step that could alter the conclusions reached above.

It is worth emphasizing again that the empirical patterns reported in the paper could have been generated by developments other than increasing financial interdependence. In comparing 1951-1972 with 1973-1988 we see, for example, a marked rise in the correlations between British, German, and Japanese output growth and rest-of-world consumption growth. Conceivably these changes give the false appearance of greater financial integration, when all that has really happened is that the potential gains from asset trade have fallen exogenously. My conclusion that for Germany and perhaps other countries there is more to the story than this is based on identifying assumptions that certainly warrant further investigation. More could be learned as well by augmenting the limited data sample used here, and by studying additional types of disturbance, such as terms-of-trade and interest-rate shocks.

References

Atkeson, A., and T. Bayoumi (1992), 'Do private capital markets insure against regional risk? Evidence from the United States and Europe,' mimeo, University of Chicago and International Monetary Fund.

Backus, D.; P. Kehoe; and F. Kydland (1992), 'International real business cycles,' *Journal of Political Economy* 100 (August), 745-75.

Backus, D., and G. Smith (1992), 'Consumption and real exchange rates in dynamic exchange economies with nontraded goods,' mimeo, New York University and Queen's University.

Baxter, M., and M. Crucini (1993), 'Explaining saving/investment correlations,' *American Economic Review* 83 (June), 230-78.

Boothe, P.; K. Clinton; A. Côté; and D. Longworth (1985), *International asset substitutability: Theory and evidence for Canada*. Ottawa: Bank of Canada.

Cochrane, J. (1991), 'A simple test of consumption insurance,' *Journal of Political Economy* 99 (October), 957-76.

Cole, H., and M. Obstfeld (1991), 'Commodity trade and international risk sharing: How much do financial markets matter?' *Journal of Monetary Economics* 28 (August), 3-24.

Devereux, M.; A. Gregory; and G. Smith (1992), 'Realistic cross-country consumption correlations in a two-country, equilibrium, business cycle model,' *Journal of International Money and Finance* 11 (February), 3-16.

Duffie, D. (1992), *Dynamic asset pricing theory*. Princeton, NJ: Princeton University Press.

Feldstein, M., and C. Horioka (1980), 'Domestic saving and international capital flows,' *Economic Journal* 90 (June), 314-29.

French, K., and J. Poterba (1991), 'International diversification and international equity markets,' *American Economic Review* 81 (May), 222-26.

Golub, S. (1991), 'International diversification of social and private risk: The U.S. and Japan,' mimeo, Swarthmore College.

Granger, C., and P. Newbold (1974), 'Spurious regressions in econometrics,' *Journal of Econometrics* 2, 111-20.

Hall, R. (1986), 'The role of consumption in economic fluctuations,' in *The American business cycle: Continuity and change*, edited by R. Gordon. Chicago: University of Chicago Press.

Kollmann, R. (1992), 'Consumptions, real exchange rates and the structure of international asset markets,' mimeo, Université de Montréal.

Leme, P. (1984), 'Integration of international capital markets,' mimeo, University of Chicago.

Mace, B. (1991), 'Full insurance in the presence of aggregate uncertainty,' *Journal of Political Economy* 99 (October), 928-56.

Obstfeld, M. (1986), 'Capital mobility in the world economy: Theory and measurement,' *Carnegie-Rochester Conference Series on Public Policy* 24 (Spring), 55-104.

Obstfeld, M. (1989), 'How integrated are world capital markets?: Some new tests,' in *Debt, stabilization and development: Essays in memory of Carlos Diaz-Alejandro*, edited by G. Calvo, R. Findlay, P. Kouri, and J. Braga de Macedo. Oxford, U.K.: Basil Blackwell.

Radner, R. (1972), 'Existence of equilibrium of plans, prices, and price expectations in a sequence of markets,' *Econometrica* 40 (March), 289-303.

Scheinkman, J. (1984), 'General equilibrium models of economic fluctuations: A survey of theory,' mimeo, University of Chicago.

Stockman, A., and L. Tesar (1990), 'Tastes and technology in a two-country model of the business cycle: Explaining international comovements,' Working Paper 3566, National Bureau of Economic Research, December.

Stokey, N., and R. Lucas (1989), *Recursive methods in economic dynamics*. Cambridge, MA: Harvard University Press.

Summers, R., and A. Heston (1991), 'The Penn World Table (Mark 5): An expanded set of international comparisons, 1950-1988,' *Quarterly Journal of Economics* 106 (May), 327-68.

Tesar, L. and I. Werner (1992), 'Home bias and the globalization of securities markets,' Working Paper 4218, National Bureau of Economic Research, November.

Townsend, R. (1989), 'Risk and insurance in village India,' mimeo, University of Chicago.

van Wincoop, E. (1992), 'Regional risksharing,' mimeo, Boston University and IGIER, Milan, November.

Table 1: Correlation coefficients for per capita consumption growth rates, 1951-1972 and 1973-1988

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
Rest of World	0.43 / 0.10	0.26 / 0.50	−0.11 / 0.72	−0.02 / 0.27	0.06 / 0.62	0.29 / 0.59	0.26 / 0.31
Canada		0.07 / 0.00	0.04 / 0.38	0.04 / −0.12	0.03 / 0.03	−0.22 / −0.15	0.60 / 0.17
France			0.12 / 0.44	0.28 / 0.42	0.18 / 0.65	−0.21 / 0.21	0.05 / 0.30
Germany				−0.12 / 0.36	0.19 / 0.45	−0.13 / 0.39	−0.04 / 0.46
Italy					0.54 / 0.37	−0.15 / 0.19	−0.03 / −0.02
Japan						−0.23 / 0.68	0.05 / 0.46
U.K.							0.39 / 0.49

Table 2: Standard deviation of domestic consumption growth relative to standard deviation of world consumption growth

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
1951-1972	2.61	1.66	2.54	1.99	2.65	2.32	2.02
1973-1988	3.92	1.84	1.50	1.90	1.99	2.63	1.84

$$\text{Standard deviation of annual world consumption growth} = \begin{cases} 0.85\% & (1951\text{-}72) \\ 1.13\% & (1973\text{-}88) \end{cases}$$

Table 3: Correlation coefficients for per capita output growth rates, 1951-1972 and 1973-1988

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
Rest of World	0.42 / 0.30	0.41 / 0.56	0.31 / 0.87	0.35 / 0.61	0.43 / 0.71	0.49 / 0.66	0.19 / 0.67
Canada		0.49 / -0.11	0.09 / 0.29	-0.11 / -0.03	0.10 / 0.23	-0.21 / 0.14	0.63 / 0.37
France			0.07 / 0.63	0.10 / 0.82	0.52 / 0.52	-0.07 / 0.50	0.25 / 0.33
Germany				0.22 / 0.70	0.09 / 0.70	0.26 / 0.66	0.18 / 0.80
Italy					0.46 / 0.43	0.22 / 0.32	0.15 / 0.43
Japan						0.33 / 0.73	0.26 / 0.63
U.K.							0.29 / 0.66

Table 4: Standard deviation of domestic output growth relative to standard deviation of world output growth

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
1951-1972	2.20	1.04	2.11	1.75	2.01	1.13	1.84
1973-1988	1.94	0.99	1.30	1.71	1.35	1.59	1.69

$$\text{Standard deviation of annual world output growth} = \begin{cases} 1.31\% & (1951\text{–}72) \\ 1.75\% & (1973\text{–}88) \end{cases}$$

Table 5: Correlation coefficients between domestic consumption growth and rest-of-world output growth

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
1951-1972	0.13	0.59	0.11	0.17	0.43	-0.09	0.07
1973-1988	-0.07	0.34	0.58	0.49	0.49	0.51	0.36

Table 6: Regressions of national on world consumption growth rates

A. 1951-1972

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W$	**1.29** (0.34)	**0.55** (0.27)	0.06 (0.51)	0.13* (0.44)	0.45 (0.58)	**1.00** (0.47)	**1.77*** (0.22)
$\bar{R}^2$	0.20	−0.05	−0.05	−0.05	−0.02	0.14	0.75
Lags	1	3	1	0	0	0	0

B. 1973-1988

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W$	0.84 (1.02)	**0.63** (0.26)	**1.14** (0.37)	0.68 (0.47)	**1.45** (0.36)	**1.77** (0.49)	**1.53*** (0.20)
$\bar{R}^2$	−0.02	0.07	0.54	0.07	0.50	0.41	0.67
Lags	0	0	1	0	0	0	1

Note: Standard errors appear below estimates of the coefficient of world consumption growth. Boldface entries of this estimate are those differing from 0 at the 5 percent significance level or below. An asterisk (*) marks coefficients that differ from 1 at the 5 percent level or below. "Lags" shows the moving-average order assumed for the equation disturbance in calculating standard errors.

Table 7: Regressions of national on rest-of-world consumption growth rates

A. 1951-1972

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W^i$	**1.13** (0.41)	0.41 (0.32)	−0.27* (0.53)	−0.04* (0.43)	0.16 (0.59)	0.67 (0.50)	0.64 (0.54)
$\bar{R}^2$	0.15	0.02	−0.04	−0.05	−0.05	0.04	0.02
Lags	1	1	0	0	0	0	0

B. 1973-1988

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W^i$	0.38 (0.71)	**0.57** (0.26)	**1.08** (0.35)	0.50 (0.48)	**1.26** (0.43)	**1.60** (0.62)	0.63 (0.40)
$\bar{R}^2$	−0.06	0.20	0.48	0.01	0.34	0.30	0.02
Lags	1	0	1	0	0	1	2

Note: Standard errors appear below estimates of the coefficient of rest-of-world consumption growth. Boldface entries of this estimate are those differing from 0 at the 5 percent significance level or below. An asterisk (*) marks coefficients that differ from 1 at the 5 percent level or below. "Lags" shows the moving-average order assumed for the equation disturbance in calculating standard errors.

Table 8: Domestic consumption growth, world consumption growth, and world output growth

A. 1951-1972

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W^i$	1.61 (0.58)	−0.32* (0.38)	−0.77* (0.70)	−0.45* (0.57)	−0.90* (0.67)	1.64 (0.64)	1.47 (0.92)
$\Delta \log GDP_W^i$	−0.46 (0.27)	0.75 (0.24)	0.52 (0.47)	0.41 (0.38)	1.09 (0.44)	−0.86 (0.40)	−0.60 (0.54)
Lags	1	0	0	0	0	0	0
H_1	0.24	0.01	0.05	0.05	0.03	0.11	0.45

B. 1973-1988

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W^i$	2.51 (2.05)	0.94 (0.52)	1.22 (0.40)	−1.00* (0.81)	1.51 (0.85)	1.48 (0.96)	0.00 (0.95)
$\Delta \log GDP_W^i$	−1.59 (1.33)	−0.28 (0.34)	−0.11 (0.39)	1.17 (0.54)	−0.18 (0.53)	0.09 (0.50)	0.54 (0.64)
Lags	0	0	1	0	0	1	1
H_1	0.43	0.23	0.85	0.08	0.80	0.60	0.53

Note: Standard errors appear below coefficient estimates. Boldface entries of coefficient estimates are those differing from 0 at the 5 percent significance level or below. An asterisk (*) marks coefficients on $\Delta \log C_W^i$ that differ from 1 at the 5 percent level or below. "Lags" shows the moving-average order assumed for the equation disturbance in calculating standard errors. Marginal significance levels are reported for tests of the hypothesis $H_1: \alpha_{iW} = 1,\ \gamma_i = 0$.

Table 9: Effects of oil-price changes on consumption growth, 1973-1988

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W^i$	−0.07 (1.30)	0.64 (0.34)	**1.29** (0.50)	**1.40** (0.29)	**1.06** (0.54)	0.88 (0.52)	0.31 (0.36)
ΔOIL	−0.02 (0.04)	0.00 (0.01)	0.01 (0.01)	**0.04** (0.01)	−0.01 (0.02)	**−0.04** (0.01)	**−0.04** (0.01)
Lags	0	0	1	1	0	2	3
H_1	0.71	0.29	0.19	0.00	0.68	0.00	0.00

Note: Standard errors appear below estimates of regressor coefficients. Boldface entries of these estimates are those differing from 0 at the 5 percent significance level or below. An asterisk (*) marks coefficients on $\Delta \log C_W^i$ that differ from 1 at the 5 percent level or below. "Lags" shows the moving-average order assumed for the equation disturbance in calculating standard errors. The reported marginal significance levels are for the F or χ^2 test of the hypothesis H_1: $\alpha_{iW} = 1,\ \gamma_i = 0$.

Table 10: Domestic consumption growth, world consumption growth, and the domestic resource limit

A. 1951-1972

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W^i$	0.62 (0.31)	0.27* (0.17)	−0.18* (0.33)	−0.32* (0.28)	−0.15* (0.37)	0.20* (0.28)	**−0.64*** (0.25)
$\Delta \log DRL_i$	**0.81** (0.12)	**0.57*** (0.20)	**0.76** (0.13)	**0.82** (0.16)	**0.76** (0.13)	**0.95** (0.14)	**1.08** (0.10)
Lags	0	3	0	2	0	0	0
H_1	0.00	0.00	0.00	0.00	0.00	0.00	0.00
H_2	0.10	0.06	0.21	0.18	0.18	0.77	0.05

B. 1973-1988

	Canada	*France*	*Germany*	*Italy*	*Japan*	*U.K.*	*U.S.*
$\Delta \log C_W^i$	−0.25* (0.47)	0.57 (0.26)	**1.07** (0.32)	1.00 (0.53)	**1.18** (0.42)	−0.16 (0.51)	**−1.27*** (0.52)
$\Delta \log DRL_i$	**0.86** (0.12)	0.21* (0.22)	0.02* (0.20)	0.47* (0.27)	0.35* (0.26)	**1.10** (0.28)	**1.54** (0.43)
Lags	0	0	1	0	0	0	0
H_1	0.00	0.20	0.97	0.19	0.37	0.01	0.00
H_2	0.40	0.00	0.00	0.11	0.01	0.92	0.05

Note: Standard errors appear below coefficient estimates. Boldface entries of coefficient estimates are those differing from 0 at the 5 percent significance level or below. An asterisk (*) marks coefficients that differ from 1 at the 5 percent level or below. "Lags" shows the moving-average order assumed for the equation disturbance in calculating standard errors. Marginal significance levels are reported for tests of the hypotheses: H_1: $\alpha_{iW} = 1$, $\gamma_i = 0$ and H_2: $\alpha_{iW} = 0$, $\gamma_i = 1$.

Table 11: Regressions of German consumption growth on world consumption growth and various macroeconomic shocks, 1973-1988

	1	2	3	4	5	6	7
$\Delta\log C_W^{GY}$	**0.73** (0.32)	**0.99** (0.37)	0.95 (0.23)	0.72 (0.55)	**0.93** (0.45)	0.77 (0.37)	**1.04** (0.44)
$\Delta\log GDP$	0.20 (0.25)	—	—	0.41 (0.46)	0.01 (0.23)	—	−0.54 (0.48)
$\Delta\log I$	—	0.02 (0.05)	—	−0.06 (0.10)	—	0.03 (0.05)	0.13 (0.11)
$\Delta\log G$	—	—	**0.49** (0.17)	—	**0.49** (0.18)	**0.50** (0.17)	**0.68** (0.23)
$\bar{R}^2$	0.46	0.44	0.66	0.43	0.63	0.65	0.65
Lags	1	1	0	0	0	0	0
H_1	0.66	0.94	0.03	0.81	0.09	0.08	0.10
H_2	0.42	0.75	0.01	0.65	0.05	0.04	0.06

Note: Standard errors appear below coefficient estimates. Boldface entries of coefficient estimates are those differing from 0 at the 5 percent significance level or below. An asterisk (*) marks first-row coefficients differing from 1 at the 5 percent level or below. "Lags" shows the moving-average order assumed for the equation disturbance in calculating standard errors. Marginal significance levels are reported for tests of two hypotheses. H_1 is the hypothesis that the coefficient of rest-of-world consumption growth is 1, those of all other variables 0; H_2 is the hypothesis that the coefficients of variables other than rest-of-world consumption growth are all 0.

Table 12: Regressions of Japanese consumption growth on world consumption growth and various macroeconomic shocks, 1973-1988

	1	2	3	4	5	6	7
$\Delta \log C_W^{JN}$	−0.01* (0.43)	0.36 (0.53)	1.21 (0.48)	−0.04* (0.46)	0.02* (0.44)	0.13 (0.59)	0.04 (0.50)
$\Delta \log GDP$	**0.82** (0.20)	—	—	**0.77** (0.29)	**0.85** (0.21)	—	**0.88** (0.36)
$\Delta \log I$	—	**0.23** (0.09)	—	0.03 (0.11)	—	**0.25** (0.10)	−0.01 (0.14)
$\Delta \log G$	—	—	0.10 (0.29)	—	−0.12 (0.20)	0.23 (0.25)	−0.14 (0.26)
$\bar{R}^2$	0.69	0.50	0.29	0.67	0.67	0.50	0.64
Lags	0	0	0	0	0	0	0
H_1	0.00	0.08	0.80	0.01	0.01	0.12	0.03
H_2	0.00	0.03	0.76	0.01	0.01	0.07	0.02

Note: Standard errors appear below coefficient estimates. Boldface entries of coefficient estimates are those differing from 0 at the 5 percent significance level or below. An asterisk (*) marks first-row coefficients that differ from 1 at the 5 percent level or below. "Lags" shows the moving-average order assumed for the equation disturbance in calculating standard errors. Marginal significance levels are reported for tests of two hypotheses. H_1 is the hypothesis that the coefficient of rest-of-world consumption growth is 1, those of all other variables 0; H_2 is the hypothesis that the coefficients of variables other than rest-of-world consumption growth are all 0.

Footnotes

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[1]van Wincoop (1992) examines the degree of risk sharing evident in Japanese regional consumption data.

[2]Stockman and Tesar (1990) stress the potential importance of country-specific preference shocks in matching a real business cycle model to industrial-country data.

[3]See Radner (1972) for a similar approach to modeling possibly incomplete markets.

[4]For applications of measure-theory concepts in economics and finance, see Stokey and Lucas (1989) and Duffie (1992).

[5]Because the complete-markets case yields Pareto-optimal allocations, this efficiency condition could alternatively be derived by considering the choices a benevolent social planner would make [see, e.g., Cole and Obstfeld (1991)].

[6]If indexed bonds were widely traded, which they are not, a rejection of the implications of (7) could be construed as evidence of imperfect capital mobility, i.e., of impediments to free asset trade. My 1989 paper tested the implications of free trade in nominal bonds between the United States and Japan and Germany. I found evidence of substantial trade impediments before the early 1970s, but not afterward. To conserve space, I do not carry out analogous tests in the present paper.

[7]My preferred point estimate for ρ was 1.52.

[8]See the studies listed in the introductory section. Simple correlations of log consumption differences for this paper's sample are presented at the end of this section.

[9]In their setups, these fixed effects arise from planner utility weights in a social welfare function.

[10]Kollmann (1992) reports similar findings for different consumption data sets.

[11]Equation (19) could have been derived from (6).

[12]An alternative estimation approach would be use instrumental variables correlated with the growth in world consumption. Lagged variables are plausible candidates for instruments, but the inherent near-unpredictability of consumption changes makes it difficult to find lagged variables that are closely correlated with $\Delta \log C_{Wt}$. Tim Cogley has suggested, in analogy with Hall (1986), that a contemporaneous variable such as world military expenditures might provide a suitable instrument for $\Delta \log C_{Wt}$. I plan to pursue this suggestion in future work.

[13] These are variables 3 through 6 from appendix A.1 of Summers and Heston (1991). I also used population (variable 1).

[14] The countries included are Kenya, Morocco, South Africa, Canada, Costa Rica, the Dominican Republic, El Salvador, Guatemala, Honduras, Mexico, Trinidad and Tobago, the United States, Argentina, Bolivia, Chile, Colombia, Ecuador, Paraguay, Peru, Uruguay, India, Japan, Pakistan, the Philippines, Thailand, Austria, Belgium, Cyprus, Denmark, Finland, France, West Germany, Greece, Iceland, Ireland, Italy, Luxembourg, the Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, Turkey, the United Kingdom, Australia, and New Zealand.

[15] Thus, in Table 1's first row I report the correlation of $\Delta\log C_{it}$, not with $\Delta\log C_{Wt}$, but with $\Delta\log C^{i}_{Wt}$, where $C^{i}_{Wt} \equiv (1-n_{it})^{-1}(C_{Wt} - n_{it}C_{it})$.

[16] The estimates were done in RATS. When there was strong evidence of serial correlation, standard errors were corrected using the "lags" option in LINREG, with a damping factor of 0.8. In the tables, "Lags" indicates the order moving-average process assumed for the equation disturbance. Because the time-series sample under study here is so small, the autocorrelation corrections suffer from a small-sample bias that seems to understate standard errors. I have therefore tried to be conservative in using the correction and in drawing inferences from corrected estimates.

[17]The theoretically expected slope coefficient is still 1 because (12) is replaced by:

$$\log C_{it} = \log C^{i}_{Wt} + \log C_{i0} + (\log\beta_{i})(t/\rho) + \left\{\theta_{it}/\rho - \log[\tfrac{1}{1-n_{it}} \textstyle\sum_{j\neq i}\beta_{j}^{t/\rho}\exp(\theta_{jt}/\rho)n_{jt}C_{j0}]\right\}.$$

[18]The price of oil is an index of the U.S. dollar prices of Saudi Arabian crude petroleum exports, as reported in the International Monetary Fund's *International Financial Statistics*. These dollar prices are deflated by the U.S. GNP deflator reported in the *Economic Report of the President*.

[19]Even though there are no major oil exporters in my 47-country index, the countries do not all depend on oil imports to the same extent. Thus, countries face differential levels of oil-price risk and can benefit from reallocating that risk through trade.

[20] Let F be net factor payments from abroad, CA the current-account balance, and S national saving. One version of the national-income identity is $I = CA + (GDP + F - C - G) = CA + S$; Feldstein and Horioka (1980) in effect regress I on S to determine whether CA has an impact on domestic investment independent of S. Another way to write the national-income identity--one that highlights the dependence of consumption and investment on *all* net resources from abroad--is as $C = (F - CA) + (Y - I - G) = TR + (Y - I - G)$; when $\alpha_{iW} \approx 0$ because of low capital-market integration, F should be negligible too and regression (25) should lead to the same result ($\gamma_i \approx 1$) as a time-series version of the Feldstein-Horioka regression. Notice that data on F, which are notoriously inaccurate, are *not* required for (25), as they are for accurately measuring S. For a discussion of the biases lack of accurate data on F could cause, see my 1986 paper.

[21] The log oil-price change was entered into the regressions for Italy and the U.K. in panel B (the only cases in which oil entered significantly). Coefficient estimates for oil are not reported.

[22] See Boothe, Clinton, Côté, and Longworth (1985).

[23] In the Table 10 regressions for Japan involving the domestic resource limit $DRL = GDP - I - G$, it is possible that the strong correlation between output and investment growth (the 1973-1988 correlation coefficient is 0.85) reduced the composite variable to insignificance. Notice in Table 12 (regressions 2 and 6) that investment is significant when it, instead of output, is entered into the regression.