NBER WORKING PAPER SERIES

HOW INTEGRATED ARE WORLD CAPITAL MARKETS? SOME NEW TESTS

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Working Paper No. 2075

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
November 1986

Prepared for the UNU-WIDER conference in memory of Carlos Diaz Alejandro held in Helsinki, Finland, August 23-25, 1986. Michael W. Klein provided excellent research assistance; Robert E. Cumby, Alberto Giovannini, and Marc Nerlove participated in helpful discussions; Jacob A. Frenkel and Mohsin S. Khan, my conference discussants, made useful comments; and members of the Penn macro lunch group offered valuable suggestions. All opinions and any errors are, however, my own. Research support was provided by grants from the National Science Foundation and the Alfred P. Sloan Foundation. The research reported here is part of the NBER's research program in International Studies. Any opinions expressed are those of the author and not those of the National Bureau of Economic Research.

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ABSTRACT

This paper presents some new empirical evidence on the extent of world capital-market integration. The first set of tests carried out uses data from different countries to compare internationally expected marginal rates of substitution between consumption on different dates. If residents of different countries have access to a nominally risk-free bond denominated in dollars, say, their common expected marginal rate of substitution of future for present dollars should equal the gross nominal return on dollar bonds. Tests of the international equality of expected marginal substitution rates yield evidence consistent with a substantial degree of international capital-market integration after, but not before, 1973. These tests are naturally based on a particular model of intertemporal consumption choice, but direct estimation of the inter-country relationships implied by that model lends support to its assumptions. These last findings are relevant to the current debate in macroeconomics about the role of intertemporal substitution. The second set of tests conducted in this paper concerns correlations between countries' saving and investment rates. For a sample of ten countries, correlations between annual changes in saving and investment rates over the period 1948-1984 look quite similar to those those found in quarterly data. Surprisingly, however, the correlation coefficients are often lower before the mid-1960s than afterward. This finding throws further doubt on the interpretation of saving-investment correlation coefficients as structural parameters reflecting the reponse of domestic investment to shifts in national saving.

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Introduction

The vicissitudes of the international capital market are a recurring theme in the work of Carlos Diaz Alejandro. Simple microeconomic theory shows how internationally integrated financial markets can improve global resource allocation by channeling the world flow of saving toward its most productive uses. A major message of Diaz's work, however, is that a realistic analysis of the international capital market must contend with the influence of factors that sometimes are difficult to model formally: moral hazards, political pressures, and even shifts in the prevailing paradigms of economic science. Over more than a century and a half, all of these factors have helped produce a series of booms and busts in international financial intermediation.

The booming world capital market of the five decades ended by World War I provides a benchmark against which economists have often measured the adequacy of contemporary international capital flows. In that golden age, the market effected a continuing and substantial resource transfer from developed to developing countries in spite of occasional reverses.[1] The post-1945 world capital market appears to have been less vigorous on the whole. Only after the early 1970s did international lending expand to levels comparable with those of the pre-1914 period. And since the early 1980s, the net resource transfer to developing countries has stopped and a widespread default on foreign debts has been averted (so

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[1] Evidence on the absence of arbitrage opportunities between major financial centers also supports the view of a smoothly functioning world capital market in the decades before 1914. See, for example, Officer (1985).

far) only through the constant involvement of official financial agencies.

One important indicator of the contrast between the pre-1914 and post-1945 capital markets has been the average magnitude of countries' current-account imbalances in the latter period. The current account surplus, as the difference between a country's overall saving and its domestic investment, shows the amount of domestic savings being invested abroad--or, in the case of a deficit, the amount of foreign savings being borrowed to finance domestic investment. In 1965, countries classified by the World Bank as middle-income oil importers financed a mere 5 percent of their domestic investment by drawing on foreign savings. The figure rose to 7.6 percent in 1973 and to 15.4 percent by 1980, but dropped sharply after 1982.[2] Compare these figures with the one-third to one-half of Argentine investment that Diaz (1970, p. 31) reckoned was financed by foreign capital in the years 1880-1914! For developed countries in the postwar era, current accounts have tended to be even smaller (as a percentage of GNP) than for industrializing countries. The recent United States current-account deficit, which in 1985 amounted to nearly 18 percent of U.S. domestic investment, is an extreme outlier in this repect.

The fact that current accounts have on the whole been so small since 1945 is a major puzzle for economists hoping to apply open-economy theory to open-economy policy problems. Our predictions about specific policy measures, however, depend crucially on whether the limited net capital flows we observe reflect an efficient global resource allocation, given countries' preferences and intertemporal transformation

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[2] See World Bank (1985), table A.7.

opportunities, or arise instead from such barriers to capital-market integration as official controls and sovereign risk. A growing empirical literature has taken several routes in trying to assess the freedom with which capital flows across national boundaries.

In an earlier paper (1986), I surveyed two important approaches taken in the empirical literature on world capital-market integration. The first of these approaches attempts to compare the returns available on assets located in different countries. Because asset returns are inherently uncertain, the conclusions drawn from an international comparison of asset returns inevitably rest on an assumed model of the pricing of risk. To avoid taking a stand on the appropriate asset-pricing model, my earlier paper restricted its discussion to assets whose returns would be the same in all states of nature in a world of perfectly integrated capital markets. Evidence on the interest paid by onshore and offshore deposits denominated in the same currency seemed to me consistent with a high degree of international capital mobility.[3]

The second empirical approach I reviewed is based on a direct comparison of divergences between countries' saving and investment rates. This second approach, due to Feldstein and Horioka (1980) and Feldstein (1983), argues that the small size of average current accounts over long periods is indeed evidence that sizable barriers impede the free international movement of capital. I suggested that this interpretation of the data suffers from potentially serious identification problems, and presented quarterly time-series evidence with implications apparently different from those Feldstein and Horioka drew from their

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[3] Researchers who have attempted to model risk explicitly have reached differing conclusions. Two recent examples are the papers of Wheatley (1985) and Jorion and Schwartz (1986).

cross-sectional findings.

This paper develops additional evidence on the integration of world capital markets. The first set of tests I carry out is based on an international comparison of marginal rates of substitution between consumption on different dates. If residents of two countries have access to a nominally risk-free bond denominated in dollars, say, their common expected marginal rate of substitution of future for present dollars should equal the gross nominal return on dollar bonds. Tests of the international equality of expected intertemporal marginal substitution rates yield evidence consistent with a substantial degree of international capital-market integration after, but not before, 1973. These tests are naturally based on a particular model of intertemporal consumption choice, but direct estimation of the inter-country relationships implied by that model lends support to its assumptions. These last findings are relevant to the current debate in macroeconomics about the role of intertemporal substitution.

The second set of tests conducted here extends the work reported in my 1986 paper. For a sample of countries somewhat larger than the one I examined earlier, correlations between annual changes in saving and investment rates over the period 1948-1984 look quite similar to those those found in quarterly data. Surprisingly, however, the correlation coefficients are often lower before the mid-1960s than afterward. I argue that this finding throws further doubt on the interpretation of saving-investment correlation coefficients as structural parameters reflecting the reponse of domestic investment to shifts in national saving.

The paper is organized as follows. Section I examines the relation between expected intertemporal marginal substitution rates in the United

States, Germany, and Japan. Section II discusses some shortcomings of the data and methods used. As a partial check on the relevance of the conclusions drawn in section I, section III estimates the model underlying that section's tests. Section IV contains the new time-series estimates of saving-investment correlations for the postwar period.

I. A Test of World Capital-Market Integration

Recent work in finance and macroeconomics has drawn on consumption-based models of asset pricing developed by Breeden (1979), Lucas (1978), and others. These models extend to a stochastic setting Irving Fisher's (1930) celebrated account of intertemporal consumption choice under certainty. In the equilibria of the stochastic models, the joint distributions of asset returns and individual consumption satisfy a condition that generalizes Fisher's equality between marginal rates of intertemporal substitution in consumption and a relative intertemporal price.

Suppose that a typical consumer maximizes

$$(1) \quad E_t\{\sum_{\tau=t}^{\infty} \beta^{\tau-t} U(c_\tau)\}$$

subject to budget constraints, where $E_t\{.\}$ is a conditional expectation based on time-t information, $\beta < 1$ is a subjective discount factor, c_τ is consumption on date τ, and the period utility function $U(.)$ is strictly concave and differentiable. Then if R_{t+1} denotes the (possibly random) real time-(t+1) payoff on any asset relative to its real purchase price on date t, individual maximization forces the consumer's contingency plan for future consumption to obey the <u>expected</u> marginal equality

(2) $E_t\{R_{t+1} \times \beta U'(c_{t+1})/U'(c_t)\} = 1.$

This equation reduces to Fisher's marginal equality in the determnistic case.

In a world of integrated capital markets, equation (2) has strong implications about the <u>ex ante</u> relationship between consumption growth in different countries.[4] Consider two countries, a "home" country and a "foreign" country (which we make notationally distinct from the home country by using asterisks). Let P_t be the price level in the home country and i_t the nominal interest rate on a risk-free one-period bond (such as a U.S. Treasury bill). Then for this particular asset, equilibrium condition (2) takes the form

(3) $E_t\{(1 + i_t)(P_t/P_{t+1}) \times \beta U'(c_{t+1})/U'(c_t)\} = 1$

for a representative home consumer. A similar relationship naturally links the corresponding foreign variables. Foreigners consume a basket of commodities which may differ from the one consumed at home. Let the currency exchange rate X_t denote the home-currency price of foreign currency. Then the home-currency price of the characteristic foreign consumption bundle is $X_tP^*_t$, and for a foreign consumer, the <u>ex post</u> real return on the home-currency bond is

$$(1 + i_t)(X_tP^*_t/X_{t+1}P^*_{t+1}).$$

According to (2), therefore, foreign residents plan their consumption so that the following condition holds:

(4) $E_t\{(1 + i_t)(X_tP^*_t/X_{t+1}P^*_{t+1}) \times \beta^* U^{*\prime}(c^*_{t+1})/U^{*\prime}(c^*_t)\} = 1.$

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[4] Stulz (1981) presents a continuous-time analysis of open-economy asset pricing similar in spirit to the analysis carried out below.

Because the nominal interest rate i_t is part of the time-t information set, equations (3) and (4) together imply that

$$(5)\ 1/(1 + i_t) = E_t\{(P_t/P_{t+1}) \times \beta U'(c_{t+1})/U'(c_t)\}$$

$$= E_t\{(X_t P^*_t/X_{t+1}P^*_{t+1}) \times \beta^* U^{*\prime}(c^*_{t+1})/U^{*\prime}(c^*_t)\}.$$

Equation (5) states that if residents of the home and foreign countries can invest in the same nominally risk-free asset, then their expected marginal rates of substitution between current and future units of the home currency must be equal. Of course, if residents of both countries also have access to a nominally risk-free foreign-currency bond paying the interest rate i^*_t, then the home and foreign expected marginal rates of substitution between current and future units of the foreign currency must also be equal:

$$(6)\ 1/(1 + i^*_t) = E_t\{[(P_t/X_t)/(P_{t+1}/X_{t+1})] \times \beta U'(c_{t+1})/U'(c_t)\}$$

$$= E_t\{(P^*_t/P^*_{t+1}) \times \beta^* U^{*\prime}(c^*_{t+1})/U^{*\prime}(c^*_t)\}.$$

Under the rational expectations assumption, equations (5) and (6) provide the testable predictions about consumption, price-level, and exchange-rate movements that underlie the statistical tests carried out in this section and the next one.[5]

Before going on to assume the additional restrictions needed to infer testable implications from (5) and (6), I want to make two points

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[5] I am assuming that domestic and foreign agents have identical information sets. (Clearly, nominal interest rates at which both sets of residents can transact are common information.) The tests carried out below do not require this assumption provided they are based on common lagged information. Interest taxes are ignored. This omission should have little effect on the tests if tax rates are similar across countries.

about these relationships. First, if the interest rates in equations (5) and (6) are offered by assets issued in the same location (for example, if they are London Eurocurrency deposit rates), the model yields expressions for the forward foreign-exchange premium, which is related to the nominal interest-rate differential through covered interest parity. The intertemporal consumption allocation conditions have been used in this way by Hansen and Hodrick (1983), Mark (1985), Campbell and Clarida (1986), and Cumby (1986) in attempts to model forward premia. In my 1986 paper, I observed that tests which do not involve assets located in different political or regulatory jurisdictions are uninformative about capital mobility between countries. Nonetheless, the same basic theoretical framework can throw light on questions about international capital mobility if they are used to compare consumption paths in different countries. The marginal equalities in (5) and (6) do not require any particular location for the assets being considered, but they do require that residents of different countries be able to trade the same asset.

A second point about equations (5) and (6) is that they are not based on any assumption of purchasing power parity or perfect goods-market integration. The derivation of these equations requires only that measured exchange rates and price indexes reflect the true prices at which residents of the two countries can transform home or foreign money into the goods they usually consume.

To implement (5) and (6) empirically, however, two strong assumptions must now be made. First, it is assumed that consumers in each country are alike with respect to endowments and preferences, so that (5) and (6) may be tested using aggregate per capita consumption levels in the two countries. Second, it is assumed that preferences are identical in the two countries, such that the marginal utility of a consumption level c is given everywhere by

$$U'(c) = c^{-\alpha}, \quad \alpha > 0.$$

Thus, α, the reciprocal of the intertemporal elasticity of substitution, is the same in both countries, and $\beta = \beta*$. There is no justification for assumption one other than the absence of practical alternatives. The next section provides partial evidence that the data are consistent with assumption two.[6]

The assumptions just made, together with (5) and (6), lead to the equations

$$(7)\quad E_t\{(\frac{C_t}{C_{t+1}})^{\alpha}(\frac{P_t}{P_{t+1}}) \quad - \quad (\frac{C^*_t}{C^*_{t+1}})^{\alpha}(\frac{X_tP^*_t}{X_{t+1}P^*_{t+1}})\} = 0,$$

$$(8)\quad E_t\{(\frac{C_t}{C_{t+1}})^{\alpha}(\frac{P_t/X_t}{P_{t+1}/X_{t+1}}) \quad - \quad (\frac{C^*_t}{C^*_{t+1}})^{\alpha}(\frac{P^*_t}{P^*_{t+1}})\} = 0.$$

According to (7) and (8), international discrepancies between _ex post_ marginal rates of substitution are unpredictable on the basis of time-t information if everyone can trade the same nominally risk-free home- and foreign-currency bonds. Define the random variables η_{t+1} and η^*_{t+1} by

$$\eta_{t+1} = (\frac{C_t}{C_{t+1}})^{\alpha}(\frac{P_t}{P_{t+1}}) \quad - \quad (\frac{C^*_t}{C^*_{t+1}})^{\alpha}(\frac{X_tP^*_t}{X_{t+1}P^*_{t+1}}),$$

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[6] More precisely, the tests performed in the next section [which assume that (5) or (6) holds] do not reject the hypothesis that intertemporal substitution elasticities are the same in the U.S., Germany, and Japan.

$$\eta^*_{t+1} = \left(\frac{C_t}{C_{t+1}}\right)^\alpha \left(\frac{P_t/X_t}{P_{t+1}/X_{t+1}}\right) - \left(\frac{C^*_t}{C^*_{t+1}}\right)^\alpha \left(\frac{P^*_t}{P^*_{t+1}}\right).$$

Then (7) and (8) can be expressed compactly as

(9) $E_t\{\eta_{t+1}\} = 0,$

(10) $E_t\{\eta^*_{t+1}\} = 0.$

Both η_{t+1} and η^*_{t+1} would be observable <u>ex post</u> if the preference parameter α were known. In the tests conducted in this section, I examine conditions (7) and (8) over a wide grid of possible values for α.

In principle, conditions (9) and (10) can be falsified empirically if any information known at time t-1 or earlier is useful in forecasting values of η and η^* dated t or later. In practice, however, attention must be restricted to some subset of the information agents presumably use in forming their expectations. Because the factors that give rise to bond-market segmentation are likely to change only gradually over time, I follow the "efficient-markets" tradition of testing whether past discrepancies in marginal substitution rates help forecast future discrepancies. For different assumed values of α, I thus estimate regression equations of the form

$$\eta_t = \gamma_0 + \sum_{i=1}^{N} \gamma_i \eta_{t-i} + v_t,$$

$$\eta^*_t = \gamma^*_0 + \sum_{i=1}^{N} \gamma^*_i \eta^*_{t-i} + v^*_t,$$

where v_t and v^*_t are errors orthogonal to information dated t-1 or earlier. For each assumed value of α, a test of the hypothesis

$$H_0\colon \gamma_0 = \gamma_1 = \ldots = \gamma_N = 0$$

tests whether people in different countries equate ex ante marginal rates of substitution of present for future units of home currency through intertemporal trading at the same home-currency interest rate. Similarly, given α, a test of the hypothesis

$$H_0^*\colon \gamma_0^* = \gamma_1^* = \ldots = \gamma_N^* = 0$$

tests whether people in different countries equate ex ante marginal rates of substitution of present for future units of foreign currency through intertemporal trading at the same foreign-currency interest rate.

The data used were quarterly series drawn from the International Monetary Fund's International Financial Statistics data tape. The per capita consumption series were defined as nominal consumption divided by population and deflated by the consumer price index (CPI). Price levels are CPIs and exchange rates are quarterly averages. Over a grid of ten α values ranging from $\alpha = 0.5$ to $\alpha = 25.0$, these data were used to construct η and η^* series between the United States and Germany, and between the United States and Japan. Table 1 (United States-Germany) and Table 2 (United States-Japan) report significance levels for F-statistics under the null hypotheses H_0 and H_0^* over the entire sample period 1962:II to 1985:II. The lag length for the test was set at $N = 8$ quarters.[7]

The results in Table 1 are on the whole unfavorable to both null hypotheses. For all but the three highest values of α, both H_0 and H_0^*

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[7] The raw data run from 1960:I to 1985:II, but after first-differencing and then allowing for eight lags, only observations from 1962:II onward can be used in the regressions.

Table 1
Tests of H_0 and H_0^* between the United States and Germany

Sample: 1962:II - 1985:II

α	H_0 Significance	H_0^* Significance
0.5	.093	.091
0.75	.074	.075
1.0	.060	.062
1.5	.042	.046
2.0	.035	.039
3.0	.040	.045
5.0	.084	.089
7.0	.117	.121
12.0	.152	.155
25.0	.284	.291

Note: The distribution of the test statistic is F(9,84) under either null hypothesis. The significance level is the probability under the null hypothesis of drawing a realization of the test statistic at least as high as the calculated value.

Table 2
Tests of H_0 and H_0^* between the United States and Japan

Sample: 1962:II - 1985:II

	H_0	H_0^*
α	Significance	Significance
0.5	.028	.018
0.75	.038	.025
1.0	.045	.031
1.5	.040	.030
2.0	.024	.019
3.0	.008	.007
5.0	.005	.004
7.0	.012	.009
12.0	.110	.087
25.0	.890	.861

Note: The distribution of the test statistic is F(9,84) under either null hypothesis.

can be rejected at the 10 percent significance level or below. Since α values of 7 or greater are implausibly high, the tests seem to indicate that over the entire period since 1962:II, expected intertemporal marginal substitution rates for dollars and deutschemarks have no been the same in the United States and Germany.

The results for the U.S. and Japan show an even stronger rejection of the null hypotheses over the sample period as a whole. Except for the implausible cases α = 12 and 25, both H_0 and H_0^* are always rejected at the 5 percent level or below.

It is unlikely that the entire sample period studied in Tables 1 and 2 is structurally homogeneous. In particular, the international capital market has expended dramatically since the early 1970s, when a marked liberalization of industrial-country capital markets began.[8] One possible explanation of the rejections is that they reflect the influence of the earlier observations, which come from a period when international financial markets seem to have been less interdependent than they are today.

To check this possibility, I split the sample at 1973:I and conducted separate tests for the resulting subsamples. The results are reported in Tables 3 (United States-Germany) and 4 (United States-Japan). The striking feature of the results in Table 3 is that for all values of the inverse intertemporal substitution elasticity, the null hypotheses is always rejected at lower significance levels in the first subsample than in the second. This finding is consistent with the hypothesis that capital-market integration has increased since the early 1970s. In most cases, however, rejection of the null hypotheses in the

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8 The expansion in international financial intermediation is documented and analyzed by Bryant (1985).

Table 3

Subsample Tests of H_0 and H_0^* between the United States and Germany

	1962:II - 1972:IV		1973:I - 1985:II	
	H_0	H_0^*	H_0	H_0^*
α	Significance		Significance	
0.5	.074	.069	.562	.566
0.75	.110	.108	.520	.526
1.0	.135	.138	.479	.490
1.5	.213	.226	.414	.429
2.0	.235	.256	.369	.388
3.0	.190	.210	.347	.367
5.0	.149	.160	.497	.507
7.0	.146	.153	.674	.680
12.0	.161	.166	.842	.853
25.0	.228	.235	.899	.920

Note: Under either null hypothesis, the distribution of the test statistic is $F(9,34)$ for the first subsample and $F(9,41)$ for the second.

Table 4

Subsample Tests of H_0 and H_0^* between the United States and Japan

	1962:II - 1972:IV		1973:I - 1985:II	
	H_0	H_0^*	H_0	H_0^*
α	Significance		Significance	
0.5	.000	.000	.654	.605
0.75	.000	.000	.773	.723
1.0	.000	.000	.866	.824
1.5	.000	.000	.950	.928
2.0	.000	.000	.955	.940
3.0	.000	.000	.869	.843
5.0	.000	.000	.747	.705
7.0	.000	.000	.776	.748
12.0	.001	.001	.874	.869
25.0	.002	.002	.985	.983

Note: Under either null hypothesis, the distribution of the test statistic is $F(9,34)$ for the first subsample and $F(9,41)$ for the second.

first subsample is possible only at significance levels higher than 10 percent. This result suggests that the test may be weak, so conclusions about the second subsample cannot be drawn with confidence in the U.S.-Germany case.

The subsample tests comparing the United States and Japan tell a somewhat stronger story. Table 4 reports that for the period ending in 1972:IV, both null hypotheses are rejected at extremely low significance levels (which in most cases are essentially zero). Nonetheless, the signifcance levels of the test statistics are all extremely high for the period beginning in 1973:I. The results suggest that in the recent period, U.S. and Japanese consumption have behaved as if residents of the two countries had access to the same risk-free borrowing and lending opportunities in both dollars and yen. This was decidedly not the case before the early 1970s.

Another interpretation of the results comes from the fact that the ex post international differences between marginal rates of substitution become substantially more variable after 1973. On this interpretation, the higher test significance levels found in the second subsample reflect a drop in the test's power caused by additional noise in the data, not an increase in world capital-market integration. In principle, this ambiguity can be resolved in the future when more data are available.

II. Discussion

Some important caveats apply to the interpretation of the previous section's results:

1. The consumption series I have used include expenditure on

durable goods. Most recent studies of consumption use either expenditure on non-durables or expenditure on non-durables plus services. Both of these measures are only partial measures of consumption: implicit (or explicit) in the use of these measures is the arbitrary assumption that the excluded portion of consumption enters the utility function in a separable manner. As Mankiw, Rotemberg, and Summers (1985) argue, however, the separability assumption is implausible. Since some degree of misspecification seems likely no matter what consumption measure is chosen, results based on the consumption measure utilized above are of interest. Future research should examine the sensitivity of the results to alternative consumption proxies.

2. Available published consumption data are seasonally adjusted. The first-order Euler condition (2) from which the tests are derived, however, applies to seasonally unadjusted data. Miron (1985) has constructed seasonally unadjusted data for U.S. consumption and shown that the estimation of equations like (2) may be quite sensitive to the use of seasonal prefilters.[9] The tests in this paper, however, are based on data in the form of inter-country differences. This may reduce the bias due to deseasonalization, particularly if deseasonalization practices are similar across countries.

3. The theory underlying equation (2) assumes that consumption is uniform over the time period beginning on date t, with the consumption decision made at the beginning of t and all variables dated t in the consumer's time-t information set. In reality, the data used are quarterly averages, so measured consumption over the quarter starting on date t incorporates information that accrues between dates t and t+1.

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[9] Singleton (1986) gives a useful theoretical discussion of the effect of prefiltering in estimating Euler-equation models.

Hall (1985) has raised this point in connection with empirical studies of the intertemporal elasticity of substitution in the U.S.; since the issue is also important in the next section, I discuss it at greater length there.

4. If the conditional distributions of economic variables change over time, estimation in a finite sample may yield misleading inferences even if unconditional distributions are constant. This problem is essentially the "peso problem" discussed in the literature on exchange market efficiency. At the very least, shifting conditional distributions will induce conditional heteroscedasticity into estimation problems, and econometric technique should take this feature of the data into account. Although Cumby and I (1984) present evidence of conditional heteroscedasticity in data on exchange rates, interest rates, and prices, the estimates in the present paper assume the problem is unimportant. Clearly, future work will have to check on the validity of that assumption.

A more fundamental question is whether the model underlying the tests in this section has any claim to empirical validity. Because the tests are joint tests of certain propositions about capital mobility <u>and</u> a particular model of consumer behavior, test results have no implications about capital mobility if the model is wrong. It is therefore important to examine independent evidence on the adequacy of equations (1) and (2) as descriptions of economic behavior in the real world.

Much of the evidence on this question is discouraging. Studies of U.S. consumption by Hansen and Singleton (1982) and by Mankiw, Rotemberg, and Summers (1985) reject the model in many cases, often obtaining <u>negative</u> point estimates of the intertemporal elasticity coefficient α. Mark (1985) obtains estimates of α which, while positive, are in most cases imprecisely measured and implausibly high.

Some countervailing considerations suggest, however, that complete abandonment of the model given by (1) and (2) may be premature. In the study mentioned above, Miron (1985) finds that the model cannot be rejected for U.S. data if seasonally unadjusted data are used. As I suggested earlier, estimates such as those in the present paper, which are based on inter-country differences, may be less sensitive to problems of seasonality. In addition, tests of Euler conditions that use data from only a single country must find appropriate data series on rates of return. Some researchers, such as Summers (1984), suggest that this is a major difficulty.

Several studies point to liquidity constraints as a possible cause of deviations from (2) in the aggregate. Zeldes (1985), for example, analyzes data from the Michigan Panel Study on Income Dynamics and finds that the Euler condition is rejected for families with low ratios of liquid wealth to income, but not for the others. From that finding, and from direct estimates of the Lagrange multipliers associated with binding borrowing limits, he concludes that liquidity constraints may lie behind the rejections of (2) by U.S. aggregate data. International synchronization of monetary conditions could give rise to a high positive correlation between the fractions of households that are liquidity constrained in different countries. In this case, aggregate tests comparing consumption growth in different countries might be less sensitive than single-country tests to the presence of some liquidity-constrained households.

Another possible cause of the disappointing results reported by Hansen and Singleton (1982) and others is the existence of preference shocks or other random factors that are unobserved by the econometrician but prevent (2) from holding exactly. To the extent that disturbances are correlated across countries, tests based on cross-country comparisons of consumption behavior may again yield less biased results.

It seems fair to describe the evidence on the underlying Euler condition as mixed at best. In the next section, I therefore report my own attempt to estimate the model using inter-country differences of U.S., German, and Japanese data. The model imposes several strong restrictions on the data. Rejection of these restrictions would call into question the interpretation given to the results of section I. Conversely, results that are reasonably in accord with the model's predictions would suggest that the results of section I are relevant for evaluating world financial-market integration.

III. Cross-Country Tests of the Consumption Model

A test of the consumption model used in section I can be based on equations (7) and (8). To derive readily estimable equations, I follow Hansen and Hodrick (1983), Hansen and Singleton (1983), and Hall (1985) in assuming that per capita consumption levels, price levels, and the exchange rate are lognormally distributed in equilibrium, that is, that the natural logarithms of these variables are normally distributed. No attempt will be made to write down a general-equilibrium model that explicitly derives a lognormal distribution for these endogenous variables from the distributions of the exogenous variables.

Denoting by lower-case letters natural logarithms of the corresponding upper-case variables, I assume that the vector

$$y_t' = [c_t, c_t^*, p_t, p_t^*, x_t]$$

is generated by the autoregressive process

$$[I - A(L)]y_t = A_0 + \mu_t,$$

where I is the 5 × 5 identity matrix, A_0 is a 5 × 5 matrix of constants, and A(L) is a polynomial in positive powers of the lag operator L. A lognormal model results from assuming that the vector μ_t of disturbances is covariance stationary and normally distributed. Thus, the conditional mean of y_t may vary over time, but because μ_t is distributed independently of the information set $\theta_{t-1} = \{y_{t-1}, y_{t-2}, \ldots\}$, the covariance matrix of y_t conditional on θ_{t-1} is a time-independent constant matrix.

The restricted information set θ_t is a subset of the broader information available to agents in the economy. Let $E_t'\{.\}$ denote a conditional expectation with respect to the restricted information set, that is, $E_t'\{.\} = E\{.|\theta_t\}$. Then equations (7) and (8) continue to hold if $E_t\{.\}$ is replaced everywhere by $E_t'\{.\}$.

For the empirical exercise of this section, I drop the assumption that $\alpha = \alpha^*$ so that it can be tested against the data.[10] If the restricted expectations operator is applied to (7), the equation that results is therefore

$$(11)\quad E_t'\{\exp[-\alpha\Delta c_{t+1} - \Delta p_{t+1}]\} = E_t'\{\exp[-\alpha^*\Delta c^*_{t+1} - \Delta x_{t+1} - \Delta p^*_{t+1}]\},$$

where $\Delta = 1 - L$. Lognormality now implies that (11) can be written as

$$(12)\quad \begin{aligned} &\exp[E_t'\{-\alpha\Delta c_{t+1} - \Delta p_{t+1}\} + V_t'\{-\alpha\Delta c_{t+1} - \Delta p_{t+1}\}/2] \\ &= \exp[E_t'\{-\alpha^*\Delta c^*_{t+1} - \Delta x_{t+1} - \Delta p^*_{t+1}\} + V_t'\{-\alpha^*\Delta c^*_{t+1} - \Delta x_{t+1} - \Delta p^*_{t+1}\}/2], \end{aligned}$$

where $V_t'\{.\}$ is a variance conditioned on θ_t. As noted earlier, these

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[10] The assumption $\beta = \beta^*$ is also inessential at this point. Relaxing that assumption affects only the interpretation of the constant terms in the equations estimated below.

conditional variances are time-independent constants. Define the percentage change in the real exchange rate of the home currency as

$$\Delta q = \Delta x + \Delta p^* - \Delta p.$$

Then (12) implies

$$(13)\quad E_t'\{\Delta q_{t+1}\} = \sigma + \alpha E_t'\{\Delta c_{t+1}\} - \alpha^* E_t'\{\Delta c_{t+1}^*\},$$

where σ is a constant that depends on the time-independent conditional covariances in (12).[11]

The economic intuition behind (13) is standard. In a deterministic, continuous-time analogue of the present model, the marginal utility of consumption in each country grows at a proportional rate equal to the difference between the rate of domestic time preference and the domestic real interest rate. By interest parity, the international difference between home and foreign real interest rates is the percentage change in the real exchange rate, q. Thus, the difference between the derivatives $\alpha(dc/dt)$ and $\alpha^*(dc^*/dt)$ is dq/dt plus a constant reflecting any international time-preference difference. Equation (13) is the same condition in expectation, adjusted by a constant risk premium.

Equation (13) must be expressed in terms of observables before it can be estimated. Define the expectational errors

$$v_{t+1}^q = \Delta q_{t+1} - E_t'\{\Delta q_{t+1}\},$$

$$v_{t+1}^c = \Delta c_{t+1} - E_t'\{\Delta c_{t+1}\},$$

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[11] Of course, if (8) also holds, it can be used to derive an equation that differs from (13) only because of a different constant term, σ^*. The condition $\sigma = \sigma^*$ is, however, an equilibrium condition of the model if (7) and (8) both hold. This equality provides an additional restriction on the model which should be tested in future work.

$$v^{c*}_{t+1} = \Delta c^*_{t+1} - E_t\{\Delta c^*_{t+1}\}.$$

Substitution of these expressions into (13) leads to

$$(14)\quad \Delta q_{t+1} = \sigma + \alpha\Delta c_{t+1} - \alpha^*\Delta c^*_{t+1} + v_{t+1},$$

where $v_t = v^q_t - v^c_t + v^{c*}_t$. By construction, v_t is serially uncorrelated and uncorrelated with any variables in the information set θ_{t-1}. These properties of v_t imply that the parameters of (14) may be estimated by instrumental variables, with variables in θ_{t-1} serving as instruments.[12]

In a multi-country framework, there are also cross-equation restrictions that can be tested as an additional check on the model. Take the "starred" country in (14) to be the United States. Then for Germany and Japan, (14) implies the relationships

$$(15)\quad \Delta q^{DM/\$}_t = \sigma_1 + \alpha_G\Delta c^G_t - \alpha_{US}\Delta c^{US}_t + v^{GUS}_t,$$

$$(16)\quad \Delta q^{¥/\$}_t = \sigma_2 + \alpha_J\Delta c^J_t - \alpha_{US}\Delta c^{US}_t + v^{JUS}_t.$$

Equations (15) and (16) can be estimated jointly under the restriction that α_{US} be the same in both equations, and that restriction can be tested.

Notice that the disturbances v^{GUS}_t and v^{JUS}_t in (15) and (16) are likely to be highly correlated contemporaneously, if only because both include as an additive component the innovation in U.S. consumption. The two-equation system can therefore be estimated most efficiently by three-stage least squares, which takes the contemporaneous error covariance into account. The instrumental variables used in three-stage

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[12] Instrumental-variable methods are necessary because both Δc_t and Δc^*_t are correlated with v_t in general.

least squares estimation were a constant and the first through third lags of $\Delta q^{DM/\$}$, $\Delta q^{¥/\$}$, Δc^{US}, Δc^{G}, and Δc^{J}.

With three lags of the variables used as instruments, the remaining sample period is 1961:I-1985:II. Over that period, the estimated preference parameters are

$$\alpha_{US} = \underset{(0.778)}{2.669}, \quad \alpha_G = \underset{(0.761)}{-0.432}, \quad \alpha_J = \underset{(0.428)}{0.808},$$

where standard errors are given in parentheses. The model restriction that the coefficient of Δc^{US} be the same in both (15) and (16) is not rejected by the data: the significance level of the $\chi^2(1)$ test statistic is .355.

The results are somewhat favorable for the model, but not completely so. For the United States and Japan, the parameter estimates are of reasonable magnitude and quite significant. They are roughly consistent with the magnitudes found by Hansen and Singleton (1983), who also used a logarithmic specification but estimated Euler equations like (2) jointly with consumers' linear forecasting equations. In addition, the key cross-equation restriction implied by the model appears consistent with the data. The estimated intertemporal substitution parameter for Germany is negative, however, implying a convex utility function. Even though the German estimate is insignificant, its incorrect sign is troubling.

In light of the tests carried out in section I, it is of interest to test the restriction $\alpha_{US} = \alpha_G = \alpha_J$ that was assumed there. The restriction can be rejected at the 2.5 percent significance level.

A problem with the foregoing results arose already in section I: we have good reasons for believing that the structure of world capital markets changed dramatically after the early 1970s. This structural change may be behind the model's uneven empirical performance, so it is informative once again to split the sample and perform separate subsample tests.

Estimation over the subsample 1961:I - 1972:IV yields the estimates

$$\alpha_{US} = \underset{(0.501)}{0.897}, \quad \alpha_G = \underset{(0.301)}{0.175}, \quad \alpha_J = \underset{(0.242)}{0.067}$$

when the cross-equation restriction is imposed. The significance level of the test statistic for those restrictions is .721. The parameters are all correctly signed, but smaller and less significant than those found over the complete sample. These characteristics of the estimates is unsurprising in view of the low variability of real exchange rates over the first subsample period compared to the second. The restriction that all the α's are equal cannot be rejected for this sample; the point estimate for the common value of α is 0.244, and its standard error is 0.164.

When the model is estimated over 1973:I-1985:II the results are

$$\alpha_{US} = \underset{(1.015)}{2.254}, \quad \alpha_G = \underset{(1.306)}{0.804}, \quad \alpha_J = \underset{(0.611)}{1.086};$$

the cross-equation restriction cannot be rejected. (The significance level for the test statistic is .759.) These results are closer to the full-sample results, except that the German preference parameter is correctly signed. The parameter estimate is, however, insignificantly different from zero. The $\chi^2(3)$ test statistic for the hypothesis that all the α's are equal has a significance level of .693, so that hypothesis appears to fit the data. The estimate of α under this restriction is 1.240, with a standard error of 0.523. On the whole, the results from subsample two support the model, as well as the international equality of intertemporal substitution elasticities that was assumed in section I.

As noted in the last section, Hall (1985) has argued that the time aggregation problem inherent in existing consumption data may bias results such as those reported above. He suggests lagging instruments an additional period, and shows that the results of Hansen and Singleton (1983) are quite sensitive to the timing of the instrument set. To check whether the time-aggregation issue raised by Hall has an important impact on the results, I now discuss estimates in which the first lag of each instrument used is omitted. Thus, the estimates below are based on an instrument set containing only a constant and the second and third lags of $\Delta q^{DM/\$}$, $\Delta q^{¥/\$}$, Δc^{US}, Δc^{G}, and Δc^{J}. The results are summarized in Table 5.

The full-sample results are quite similar to those found using the original set of instrumental variables. Because of probable structural shifts, however, the subsample findings are of greater interest. For the 1961:I-1972:IV sample, the model appears to break down completely when the instruments are changed. All coefficients are incorrectly signed, quite insignificant, and small in absolute value. Once again, however. these results are to be expected in light of the relatively low capital-market integration and real exchange rate variability of the period.

The results for the second subsample, 1973:I-1985:II, are similar to those found with the original instrument set. The main differences are that the point estimate for Germany is once again negative while the point estimate for Japan is substantially higher. The cross-equation restriction easily fits the data, as does the restriction that the three α's are the same. The estimated common value of α is plausible, and the estimate is significant at the 5 percent level.

Table 5
Estimates of Preference Parameters for the United States, Germany, and Japan

<u>Sample: 1961:I-1985:II</u>

$$\alpha_{US} = \underset{(0.944)}{2.009}, \qquad \alpha_{G} = \underset{(1.111)}{-0.937}, \qquad \alpha_{J} = \underset{(0.575)}{0.964}$$

Test of cross-equation restriction: $\chi^2(1) = 1.179$, significance = .278

Test of $\alpha_{US} = \alpha_G = \alpha_J$: $\chi^2(3) = 5.363$, significance = .147

$$\alpha = \underset{(0.526)}{0.632}$$

<u>Sample: 1961:I-1972:IV</u>

$$\alpha_{US} = \underset{(0.757)}{-0.091}, \qquad \alpha_{G} = \underset{(0.492)}{-0.101}, \qquad \alpha_{J} = \underset{(0.313)}{-0.037}$$

Test of cross-equation restriction: $\chi^2(1) = 0.916$, significance = .338

Test of $\alpha_{US} = \alpha_G = \alpha_J$: $\chi^2(3) = 0.932$, significance = .818

$$\alpha = \underset{(0.238)}{-0.062}$$

<u>Sample: 1973:I-1985:II</u>

$$\alpha_{US} = \underset{(1.363)}{2.594}, \qquad \alpha_{G} = \underset{(1.507)}{-0.166}, \qquad \alpha_{J} = \underset{(0.896)}{1.949}$$

Test of cross-equation restriction: $\chi^2(1) = 0.021$, significance = .884

Test of $\alpha_{US} = \alpha_G = \alpha_J$: $\chi^2(3) = 1.885$, significance = .597

$$\alpha = \underset{(0.767)}{1.524}$$

Note: Standard errors appears in parentheses. The α estimate reported after the test of $\alpha_{US} = \alpha_G = \alpha_J = 0$ is the estimated common value of α under this hypothesis.

Taken as a whole, the results point to the persistently insignificant and frequently incorrectly-signed German preference parameter as the model's major empirical shortcoming. Another source of concern is evidence of some serial correlation in the equation residuals. Even though the procedure suggested by Hall (1985) does not make a dramatic difference for the parameter estimates, the timing problem Hall discusses may induce serial dependence in equation disturbances.[13] A more detailed specification analysis is therefore needed before firm conslusions can be drawn. Tentatively, however, it seems reasonable to view the results of this section as generally supporting the model used to construct the tests in section I.

A potential criticism of this view comes from the empirical literature on the determinants of forward foreign-exchange premia. As Hansen and Hodrick (1983) showed, the lognormal model implies a constant expected return to forward speculation. Their empirical tests rejected the resulting model of the forward premium. The evidence on conditional heteroscedasticity reported by Cumby and me (1984) also contradicts lognormality, as do Cumby's (1986) explicit estimates of forward-exchange risk premia, which vary significantly over time.[14] It is possible that the tests of this section are less sensitive to deviations from lognormality than tests using forward-market data. A closely re-

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[13] Hall's criticism also applies to the tests carried out in section I. When those tests were re-run using regressions on lags two through nine of the dependent variable (rather than regressions on lags one through eight), the results were qualitatively the same. Not surprisingly, though, significance levels tended to be higher.

[14] Fama (1984) and Hodrick and Srivastava (1986) report additional test results showing the variability of risk premia. Some indirect evidence comes from Hansen and Singleton (1983, pp. 262-264), who are able to reject a lognormal model in the closed-economy U.S. context.

lated conjecture is that this paper's tests are less sensitive to peso problems, since the tests involve only a single asset. In future work, it will be important to check these conjectures by applying distribution-free estimation procedures of the type employed by Hansen and Singleton (1982) and Mankiw, Rotemberg, and Summers (1985). Stronger tests can also be constructed by expanding the sample of countries.

IV. More on the Correlation between Saving Rates and Investment Rates

In my 1986 paper I reported time-series estimates, for several countries, of the correlation between quarter-to-quarter changes in saving and investment rates. The sample period ran from around 1960 to the early 1980s. Those results were compared with the cross-sectional findings reported by Feldstein and Horioka (1980) and Feldstein (1983). I argued strongly in the paper that serious identification problems make it difficult to interpret saving-investment correlations as unambiguous evidence about capital mobility, either in a time-series or cross-sectional context. Nonetheless, the pattern of time-series correlations I found in the quarterly data seemed to me inconsistent with the Feldstein-Horioka conclusion that capital is essentially immobile in some long-term sense.

In this section I extend my earlier work by presenting time-series estimates of correlations between annual changes in saving and investment rates. There are four reasons why tests based on annual data are of interest. First, use of annual data allows me to expand the sample of countries and the sample period of the test. Second, annual data may be more reliable than quarterly data, which are often based on interpolation and other approximate procedures. Third, annual data are not subject to seasonality. Fourth, short-term capital movements that are essentially self-reversing (such as trade credits) should be less important in annual than in quarterly data. Thus, calculations based on annual data may come closer to addressing the issues of "long-term" capital mobility that Feldstein and Horioka seem to have in mind.

The data I use are nominal yearly national account data from the International Financial Statistics data tape. Saving, S, is defined as gross national product (GNP) minus private plus government consumption. Investment, I, is gross fixed capital formation plus the change in stocks.[15] The correlations computed are those between $\Delta(S/GNP)$ and $\Delta(I/GNP)$, where Δ is now an annual first difference.

Table 6 reports the estimated correlation coefficients between year-to-year changes in the saving rate and the investment rate for ten countries. The sample period runs from around 1950 to 1984 in most cases, and because structural homogeneity is unlikely over such a long time span, I have split the sample period at 1967. The standard errors of these coefficients were calculated using the spectral estimator described in Obstfeld (1986).

Two major empirical regularities seemed to emerge from my earlier quarterly estimates. First, the estimated correlation coefficient r_{SI} between $\Delta(S/GNP)$ and $\Delta(I/GNP)$ seemed positively related to country size, and was statistically insignificant for some small countries and sample periods. Second, r_{SI} fell for all but one country between the 1960-1972 period and the period beginning in 1973. I noted that the first regularity was consistent with a high degree of world capital-market

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[15] Government consumption includes government investment in the U.S. data, while in the other countries government investment is included in I. When the alternative accounting convention was applied to the U.S., however, the estimation results were virtually the same.

Table 6

Saving-Investment Correlations Based on Annual Data

Country	Period	Correlation	Period	Correlation
Australia	1953-1966	-0.419 (0.272)	1967-1984	0.420 (0.246)
Austria	1949-1966	0.645 (0.288)	1967-1984	0.723 (0.279)
Canada	1949-1966	0.403 (0.233)	1967-1984	0.792 (0.299)
France	1951-1966	0.251 (0.248)	1967-1982	0.520 (0.257)
Germany	1951-1966	0.609 (0.288)	1967-1984	0.789 (0.294)

Table 6 (continued)

Saving-Investment Correlations Based on Annual Data

Country	Period	Coefficient	Period	Coefficient
Italy	1953-1966	0.401 (0.335)	1967-1984	0.746 (0.286)
Japan	1953-1966	0.912 (0.368)	1967-1984	0.773 (0.332)
Mexico	1953-1966	0.819 (0.323)	1967-1983	0.429 (0.269)
United Kingdom	1949-1966	0.513 (0.258)	1967-1984	0.512 (0.256)
United States	1951-1966	0.946 (0.327)	1967-1984	0.925 (0.316)

Note: Standard errors are in parentheses. The estimated coefficients are correlation coefficients between the change in the saving rate, $\Delta(S/GNP)$, and the change in the investment rate, $\Delta(I/GNP)$, over the sample periods indicated. Details about the estimation method are given in Obstfeld (1986).

integration because of the greater ability of larger countries to influence world interest rates. The second regularity seemed consistent with an increasing degree of capital mobility after 1973, a view that is also supported by the earlier results of the present paper.

Both of these stylized facts are to some extent overturned by the data in Table 6. For most countries, r_{SI} actually rises between the first and second periods in spite of the presumed increase in the international mobility of capital. Further, the association between country size and r_{SI} is much less striking. Austria, for example, which had a very low r_{SI} value in quarterly data, has a rather high one in Table 6. In contrast, the correlation coefficients for France (which was not in my earlier sample) are rather low.

The new estimates underline the pitfalls of drawing inferences about capital mobility from correlations such as those reported in the table. The change in current account patterns between the two subsamples probably has more to do with changing investment opportunities than with the extent of capital-market integration. It is plausible that emerging investment opportunities in Europe in the 1950s and early 1960s caused a pattern of investment increases financed by foreign (mostly American) savings. A relative scarcity of such opportunities from 1967 on would have tended to increase saving-investment correlation coefficients, in spite of increasing world financial integration. The reverse story certainly seems plausible for Mexico. The development of that country's oil resources is the probable cause of the sharp drop in its saving-investment correlation between the two subsample periods.

While it is difficult to place great weight on such explanations in the absence of complete structural models of saving and investment, the numbers do pose a challenge for those who argue that capital is essentially immobile. The capital immobility hypothesis is impossible to reconcile with many of the reported correlations, some of which do not differ significantly from zero at the 5 percent level and most of which are comfortably distant from the value of unity that would characterize a closed economy. The correlation coefficients furnish statistical facts about saving and investment which future structural models will have to explain.

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