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Chapter Author: Ralph C. Bryant, Patric H. Hendershott

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EMPIRICAL ANALYSIS OF CAPITAL FLOWS: SOME CONSEQUENCES OF ALTERNATIVE SPECIFICATIONS

RALPH C. BRYANT · Federal Reserve Board
PATRIC H. HENDERSHOTT · Purdue University

WE HAVE three specific objectives in this paper. First, we wish to incorporate lagged time responses into the theoretical framework set out in an earlier work [6]. Second, we report the results of subjecting our earlier empirical analysis to further tests. These include some testing of our earlier assumptions that (*a*) the lags in the portfolio adjustment of borrowing by Japanese foreign-exchange banks are short; (*b*) the desired borrowing relationship is homogeneous of degree one in bank net worth and Japanese imports; and (*c*) equations estimated for the entire 1959–67 period hold approximately for the 1964–68 period. Finally, we give some examples of how the estimation of equations attempting to explain capital movements are affected by alternative specifications, including some cases of theoretical misspecification. As in our earlier paper, we are concerned here primarily with the theoretical framework and the empirical methodology that are appropriate for the analysis of international capital movements. All of the empirical results reported in both of our papers refer to short-term borrowing from American banks by Japan, but we believe that the inferences we draw from our intensive study of this particular capital flow have wide applicability for the empirical analysis of all international capital flows. To begin with, we summarize our basic framework and preliminary results, as described in [6].

1 SUMMARY OF BASIC FRAMEWORK AND PRELIMINARY RESULTS

THE basic empirical relationship employed in our study of short-term borrowing from banks in the United States by Japanese banks expresses the Japanese banks' long-run desired level of borrowing, *Bus**,[1] as a function of the Japanese banks' net worth, *NW*, Japanese imports, *M*, and numerous borrowing and lending rates:

$$Bus^* = f(\overset{+}{NW}, \overset{+}{M}, \overset{-}{RBus}, \overset{+}{RLld}, \overset{+?}{RBj}, \overset{+?}{RBe\$}). \tag{1}$$

The specific interest rates employed in (1) are: the "own" rate on borrowing from the United States, *RBus* (calculated as a weighted average of the rate charged on U.S. bankers' acceptances and the rate charged by U.S. banks on short-term business loans); the return earned by Japanese banks on their lending, *RLld* (calculated as a weighted average of the rates charged on all loans and discounts); the cost to Japanese banks of borrowing funds in Japan, *RBj* (calculated as a weighted average of the discount rate of the Bank of Japan and the rate in the call-money market in Tokyo); and the cost to Japanese banks of borrowing in the Eurodollar market, *RBe$* (approximated in our study by the rate paid by London banks on ninety-day Eurodollar deposits). The theory of the demand and supply of financial instruments and the empirical approximations used to derive equation (1) are spelled out in some detail in our earlier work.[2] The intellectual lineage of functions of this sort goes back to the theory of portfolio choice as worked out by Markowitz [16] and Tobin [24], [25].

[1] Much the greatest portion of Japanese short-term borrowing from banks in the United States is carried out by Japanese banks. The exact proportion is not known but is probably well in excess of 90 per cent. For this reason we treat the time series as though it were entirely borrowing by Japanese banks, even though small amounts of borrowing by Japanese nonbanks may be included. In our earlier paper [6, pp. 27, 44, 53] we tested to see whether an episode of large official borrowing in 1961–63 (there referred to as LOJ) substituted for private borrowing that might otherwise have taken place. Since our test seemed to indicate that this official borrowing had no significant impact on the remainder of Japanese borrowing, the variable *Bus* in the present paper is defined as total short-term borrowing from American banks less these special official loans in 1961–63.

[2] Ideally, (1) should contain some additional variables—most notably some proxy variables for the risks which Japanese banks associate with the holding of the different liabilities and assets on their balance sheets. See [6, pp. 5–15, 34–46].

The signs of the partial derivatives of *Bus** with respect to its determinants are noted above the symbols in (1). Increases in the net worth of the Japanese banks should, by reducing the risk associated with portfolios of given size and composition, lead to an expansion of all forms of borrowing and lending. The positive relationship between Japanese imports and borrowing from American banks reflects the institutional fact that—perhaps because there is a lower risk associated with lending against trade documents as collateral—American banks prefer to lend to Japan in the form of bank acceptances based on import-trade bills. Thus, increases in Japanese imports lead, other things being equal, to greater borrowing from the United States. The economic rationale for this institutional relationship may have been particularly strong in the early 1960's, when Japanese banks were still establishing their overseas financial contacts and their credit-worthiness in international financial markets. Later in the decade, one might have expected the importance of this relationship between borrowing from American banks and Japanese trade bills to have declined. Some of our further experiments, reported in Part 3 below, do suggest such a diminution in importance.

Increases in the own interest rate, *RBus*, obviously make the Japanese banks less eager to borrow from the United States. Increases in the returns earned by Japanese banks on their loans and discounts should unambiguously induce increases in desired borrowings. The directional impact on *Bus** of increases in other borrowing rates is ambiguous. Increases in the costs of borrowing funds at home in Japan or in the Eurodollar market will induce familiar substitution effects (more borrowing from American banks relative to other sources), but they will also produce what might be termed an "income" effect (a reduction in total borrowing, some portion of which falls on borrowing from the United States). One cannot conclude a priori that the substitution effects will outweigh the income effect, but this seems to us the more likely outcome.

In most of our empirical work so far, we have modified equation (1) by assuming that the long-run desired function is homogeneous of degree one in net worth and imports:

$$Bus^* = g\left(RBus,\, RLld,\, RBj,\, RBe\$,\, \frac{M}{NW}\right) NW. \qquad (2)$$

We find this multiplicative form appealing because it makes the impact of increments in the scale variable on the desired quantity dependent on the levels of the interest rates, and the impact of changes in the interest rates dependent on the level of the scale variable.[3] Although we cannot directly justify the assumption of linear homogeneity in terms of the Markowitz-Tobin theory,[4] we regard the assumption as a practical modification of (1) that given the present state of our theoretical and empirical knowledge, is as plausible as any other specific modification we might choose to make. The functional form of (2) allows for easy estimation of an equation where $g(\quad)$ is approximated by a linear relationship. Such an approximation is:

$$Bus^* = \left(\theta_0 + \theta_1 RBus + \theta_2 RLld + \theta_3 RBj + \theta_4 RBe\$ + \theta_5 \frac{M}{NW}\right) NW, \quad (3)$$

where θ_1 is expected to be negative; θ_2 and θ_5 are expected to be positive; and θ_3, θ_4, and θ_0 could be either negative or positive. In Part 3 below, we report some results that seem to confirm the suitability of the assumption of linear homogeneity.

If actual Japanese borrowing, *Bus*, always coincided with long-run equilibrium borrowing, Bus^*, it would be appropriate to estimate (3) directly. For at least two reasons, however, a substitution of *Bus* for Bus^* in equation (3) may not be valid. First, desired borrowing in the short run, designated here as Bus^s, may differ from long-run equilibrium borrowing because of the presence of lagged responses (see Part 2 below). Second, and more important in the case of Japanese borrowing from American banks, Japanese and American capital restrictions have significantly influenced the actual level of borrowing attained. For much of the 1959–68 period, only a fraction of desired short-run borrowing was effectively demanded or supplied. Algebraically,

[3] These responses can be most easily seen by taking the first difference of equation (2):

$$\Delta Bus^* = g(\quad)\Delta NW + NW_{-1}\Delta g(\quad), \quad (2')$$

since $\Delta(AB) = A\Delta B + B_{-1}\Delta A$. In a growing (or declining) economic world—$\Delta NW \neq 0$—changes in interest rates bring about both "existing-stock" (through the second term), and "continuing-flow" (through the first term) impacts on capital flows. Given a "once-for-all" change in one or more interest rates, the existing-stock effect produces capital flows that are also once-for-all in nature (a reallocation of existing portfolios), while the continuing-flow impact persists indefinitely as long as $\Delta NW \neq 0$. See [6, pp. 11–13].

[4] See [6, pp. 10–11].

$$Bus = \alpha Bus^s;\ 0 < \alpha \leqslant 1, \tag{4}$$

where α equals unity when capital restrictions are absent or not binding.

The fraction α is itself related to three different phenomena. The first is the basic relaxation during 1959 and 1960 of restrictions which the Japanese government had imposed on external transactions throughout the 1950's, denoted by the variable BR. The second is a "learning process" (on the part of both Japanese banks and foreign lenders) and the growth of Japanese credit-worthiness in international financial markets during 1961–64, triggered by the earlier relaxation; this group of phenomena is designated by the variable CW. The third phenomenon is the introduction and varying effectiveness of the American Voluntary Foreign Credit Restraint program during 1965–68, summarized in the variable V.[5] Expressing the fraction α as a linear function of these three variables, we have

$$\alpha = 1.0 + \beta_1 BR + \beta_2 CW + \beta_3 V;\ \beta_i < 0 \text{ for all } i. \tag{5}$$

Assuming that $Bus^s = Bus^*$ (an assumption used in our earlier paper but relaxed in this one), we substitute the equations for α and for Bus^* into (4) to give an equation which relates actual borrowing to both the economic variables and the variables representing the effects of capital restrictions:

$$\frac{Bus}{NW} = (1.0 + \beta_1 BR + \beta_2 CW + \beta_3 V)$$
$$\times \left(\theta_0 + \theta_1 RBus + \theta_2 RLld + \theta_3 RBj + \theta_4 RBe\$ + \theta_5 \frac{M}{NW}\right). \tag{6}$$

Numerous variants of this equation have been estimated on quarterly data from the 1959–67 period. Four of them are given in Table 1. In the table we report nonlinear regression estimates together with their standard errors (below and in parentheses). Because estimates with unconstrained interest-rate coefficients seemed implausible [6, p. 52], the sum of the interest-rate coefficients was constrained to equal

[5] For a more detailed discussion of these three phenomena, see [6, pp. 13–15, 36–44, D1–D4]. The variable V is defined as zero (prior to 1965) or as the ratio of foreign claims of American banks at, or over, their individual ceilings to the aggregate ceiling for all American banks reporting in the Voluntary Foreign Credit Restraint program.

TABLE 1

Estimates of the

Equation	*BR*	*CW*	*V*	Constant	*RLld* minus *RBus*
(1.1)	−.349	−.437	−.199	.060	.426
	(.049)	(.036)	(.119)	(.399)	(.099)
(1.2)	−.410	−.402	−.231	.350	.183
	(.161)	(.146)	(.082)	(.844)	(.107)
(1.3)	−.370	−.415	−.158	.316	.336
	(.044)	(.032)	(.114)	(.333)	(.040)
(1.4)	−.397	−.393	−.202	.345	.184
	(.142)	(.142)	(.073)	(.719)	(.075)

NOTE: The dependent variable is *Bus/NW*. See text and [6] for exact specification of equation and definitions of variables. *Bus* and *M* are measured in

zero by the use of interest-rate differentials. We show for each equation the standard error of estimate (SEE), the Durbin-Watson test statistic for serial correlation in the residuals (*DW*), and the parameter ρ used in an autoregressive transformation that we employ when a low *DW* figure suggests high positive serial correlation in the residuals. In general, ρ is estimated by assuming that the serial correlation follows a first-order autoregressive scheme [6, p. 54], but occasionally it is specified a priori.

Equations (1.1) and (1.2) are estimates of equation (6) with and without the autoregressive transformation. The capital control variables, the principal interest-rate differential, and the imports-net-worth variable all seem to work reasonably well. Also as expected, given the offsetting nature of the income and substitution effects, the other two interest-rate coefficients are very small; these variables contribute nothing to the explanatory power of the equation. If these two variables are omitted from the equation, we obtain equations (1.3) and (1.4).[6]

[6] Omission of the rate spreads is the result of applying the theory of second best; it is not the preferred procedure. All variables implied by the theory ought to be retained in the equation unless they have coefficients that are clearly less plausible than zero. However, when we add variables to the equations in Section 3-A (below) to test for lagged responses, these insignificant rate spreads would likely generate substantial multicollinearity problems.

Basic Equation

RBj minus $RBus$	$RBe\$$ minus $RBus$	M/NW	ρ	SEE	DW
−.043	.182	.503	–	.134	.83
(.055)	(.174)	(.162)			
−.011	−.022	.650	.99	.105	2.05
(.053)	(.093)	(.146)			
		.485	–	.134	.76
		(.160)			
		.668	.98	.105	2.11
		(.140)			

billions of dollars, NW is in trillions of yen, and interest rates are expressed in per cent per annum. The period of estimation is 1959–67.

2 ESTIMATION OF LAGGED RESPONSES

THE importance of capital restrictions as a factor causing observed quantities of financial instruments to diverge from their long-run desired levels was discussed at some length in our earlier paper. A second source of discrepancy between observed and long-run desired quantities is the fact that economic units may fully adjust their holdings of financial instruments to changes in the determinants of long-run desired holdings only after some significant period of time has elapsed. We give here reasons for expecting lagged responses and describe methods of estimating them. (Some of these methods are employed in the estimation reported in Part 3.)

A. REASONS FOR LAGGED RESPONSES

It seems unlikely that an assumption of instantaneous adjustment would be appropriate in studies of most international capital flows. Recent empirical studies of domestic financial behavior, when they have been designed so as to allow for the presence of lagged responses, have generally reported the apparent existence of significantly long lags.

The reasons for lagged adjustment in financial behavior are at least two, one of which is dependent on the other. First, portfolio adjustment necessarily entails transactions costs; both time and money must be expended in obtaining the necessary information and in implementing the adjustment. Particularly large transactions costs are incurred when financial instruments with inferior secondary markets and/or penalty rates for prepayments or redemption before maturity are involved. Since a significant proportion of both pecuniary and nonpecuniary transactions costs tends not to vary with the size of transactions, there are economies in making less frequent adjustment.

Second, it is expected future yields and risks that are relevant to investors, but only observed yields are generally available for use as regressors. Given the existence of transactions costs, a change in the yield (or the risk associated with the yield) on short-term or fixed-valued financial assets must be of some permanence if adjustment of one's portfolio is to be at all profitable. While the permanency of changes in the yields on long-term or variable-valued assets is not of importance,[7] future expected values of these yields still are. For example, if the yield on a variable-valued asset rises, an investor will not move into that asset if he expects the yield to continue rising. To do so would invite future capital losses. Since investors may extrapolate recent interest-rate changes into the future and/or expect future rates to regress toward long-run "normal" rates, a weighted average of current and lagged values of yields may be a good proxy for the expected yield. This implies that current changes in yields and risks will lead to portfolio adjustments in future periods.

Transactions costs may be relatively less important for financial institutions whose costs for obtaining information and making adjustments are probably less than those of other economic units. Also, transactions costs stemming from inferior secondary markets and/or penalty rates for prepayments or earlier redemption can be avoided by short delays only, when short-term instruments are involved. Thus, we would

[7] Even if the yield movement is reversed, a capital gain is made (if the yield has risen) or a loss is avoided (if the yield has fallen).

expect responses to be most rapid for financial institutions rearranging their short-term portfolios, and least rapid for nonfinancial units rearranging their long-term portfolios.

B. METHODS OF ESTIMATION

A complete theory of the demand and supply of particular financial instruments would require an explicit utility-maximization treatment of the manner in which economic units attempt to adjust short-run desired quantities to long-run equilibrium levels.[8] We do not have such a theory and know of no study of financial behavior containing one. What we do here is merely outline some of the empirical procedures that can be used in estimating lagged responses.

It is useful to begin by considering how an equation would have to be specified (how the variables would have to be measured) if one wished to assume no lags, i.e., instantaneous adjustment. Clearly, the dependent variable at any point in time should be related to the independent variables at the same point in time. It is customary to use average values of the dependent variable during an interval in order to eliminate fluctuations that are extremely short-run. Thus, in our case, the assumption of instantaneous adjustment would be enforced (temporarily ignoring the existence of capital controls) by estimating

$$\overline{Bus} = f(\overline{NW}, \overline{M}, \overline{R}), \tag{7}$$

where average (mean) values are denoted by a bar over the variables and, for notational purposes, the interest rates are combined into a single-rate vector. If the observation period is a quarter, $\overline{Bus}$, $\overline{NW}$ and $\bar{R}$ are average values during the quarter and $\bar{M}$ is the cumulative flow of imports during that quarter. If some of the data are measured only as of the last day of the quarter, a simple average of beginning and end of quarter values could be employed.[9]

Consider now the case where lagged responses are expected to be

[8] If there were no relevant capital restrictions and no other discrepancies between actual and short-run desired quantities, the part of the theory concerned with lagged responses would explain the adjustment of actual quantities to long-run equilibrium levels.

[9] For a careful discussion of how continuous-time models should be approximated in discrete time for the purposes of empirical research, see H. Houthakker and L. Taylor [12, pp. 11–21].

not much, if at all, longer than the observation period itself. Here the best procedure to follow is probably to assume a uniform *intraperiod* lag distribution. This is achieved by measuring the dependent variable (a stock) at the end of the observation period and measuring all independent variables as averages during the period. (This, in fact, was the procedure we employed for the equations reported in [6]). Equation (8) reflects this measurement.[10]

$$Bus = g(\overline{NW}, \overline{M}, \overline{R}). \tag{8}$$

This case is not as rare as one might imagine. In fact, where the main research objectives do not include the rigorous investigation of lagged responses, it may be good strategy to select the observation period so as to correspond as closely as possible with the expected length of lag. One might employ monthly data, for example, if it were expected that adjustment would be completed in a month, or a month and a half. Quarterly data could be used if it were thought that one to one-and-a-half quarters were required for full adjustment, while annual data could be used if one to one-and-a-half years were required. Conversely, if one is particularly interested in studying lagged responses, obviously a short enough observation period must be chosen.

The general case (still ignoring capital controls) expresses *Bus* as a function of current and lagged values of all of its determinants:

$$Bus = h(\overline{NW}, \overline{NW}_{-1}, \ldots, \overline{NW}_{-r}, \overline{M}, \overline{M}_{-1}, \ldots, \overline{M}_{-s}, \overline{R}, \overline{R}_{-1}, \ldots, \overline{R}_{-t}). \tag{9}$$

If lagged responses are significantly long in relation to the observation period, direct estimation of (9) could prove to be impossible, owing to the limited degrees of freedom or the high collinearity among the regressors, or both. However, in the case where lags are thought to be short relative to the length of the observation period, a direct approach may be feasible. For those explanatory variables being tested for lags, one-period, and possibly two-period, lagged values can be explicitly entered in the regression equation. In Part 3, we report some illustrative equations where this procedure has been followed.

[10] If the dependent variable were a flow, each independent variable would be correctly measured as the average of its values in the current period and the previous period.

Perhaps the most popular method of estimating lagged responses is use of the Koyck-Nerlove "stock-adjustment" model.[11] In this model, a constant proportion of the gap between actual and long-run desired quantities is assumed to be closed each period:

$$\Delta Bus = \lambda(\overline{Bus^*} - \overline{Bus}), \tag{10}$$

where λ is the proportion closed and must lie between zero and unity. Since any gap between actual and desired quantities is being closed continuously through time, the relevant gap to use in a discrete-time model is the *average* gap prevailing during the period. If $\overline{Bus}$ is approximated as $\frac{1}{2}(Bus + Bus_{-1})$, then equation (10) can be algebraically rewritten as:[12]

$$\Delta Bus = \gamma(\overline{Bus^*} - Bus_{-1}) \tag{10$'$}$$

or

$$Bus = \gamma\overline{Bus^*} + (1 - \gamma)Bus_{-1}, \tag{10$''$}$$

where $\gamma = 2\lambda/(2 + \lambda)$.

If one wishes to use this method in a context such as ours, the existence of capital controls complicates matters, but it can be handled as follows. First, we express equation (10″) in terms of the value of Bus that would have existed in the absence of capital controls, Bus^s:

$$Bus^s = \gamma\overline{Bus^*} + (1 - \gamma)Bus^s_{-1}. \tag{11}$$

Then, we substitute for Bus^s and Bus^s_{-1} from equation (4) and multiply through by α. This yields

$$Bus = \alpha\left[\gamma\overline{Bus^*} + (1 - \gamma)\left(\frac{Bus}{\alpha}\right)_{-1}\right]. \tag{11$'$}$$

Equation (11′) differs from the usual form of the stock-adjustment

[11] For a discussion of the implied lag structures in several of the early empirical studies in terms of the stock-adjustment model, as well as of shortcomings of that model, see Hendershott's comment [10] on Stein [21].

[12] In many empirical studies, γ has been interpreted incorrectly as the "speed of adjustment." As the text illustrates, the estimate of $\lambda = 2\gamma/(2 - \gamma)$ should be interpreted as the speed of adjustment. For slow speeds of adjustment, the differences between γ and λ are small, but the difference becomes more important, the larger the estimate for γ. (Note that an estimate of γ greater than .67 is theoretically inadmissible.)

model because instantaneous adjustment to changes in α, the capital controls construct, is assumed.[13]

A disadvantage of the Koyck-Nerlove type of model is that it places relatively severe constraints on the lag distributions. Lag weights for each explanatory variable are assumed to decline geometrically, and the rate of decline is adjudged the same for all variables.[14] Another problem with the stock-adjustment model is that the coefficient on the lagged stock will be biased upward toward one (γ and therefore λ, will be biased downward) to the extent that autocorrelation of the residuals exists.[15] What is worse, even though the "desired" stock (or α) may be seriously misspecified, the stock-adjustment model can generate superficially plausible estimates. In the extreme case, the lagged values of a dependent variable—if it is a rather smooth series—will alone "explain" the dependent variable quite well, suggesting a "significant" slow speed of adjustment, even though there are no theoretically correct variables in the equation to which the dependent variable is adjusting.[16] Some examples of the use (and misuse) of the stock-adjustment model are given in the next two sections.

The technique of polynomial approximation first used by Almon [1] is a general method of estimating longer lag distributions that also conserves degrees of freedom.[17] In this method, the distributed lag weights are assumed to lie along a polynomial of given degree; both the

[13] In most cases (though conceivably not with all types of capital controls), it would seem preferable to assume that changes in capital controls bring about very rapid adjustments. Any a priori knowledge about lagged responses to capital controls, if these lags were thought to be important, might most appropriately be taken into account when the proxy variables for capital controls are themselves being constructed. (This is the procedure we followed in [6].)

[14] These constraints can be relaxed somewhat if the function for the "desired" stock is defined to include recent lagged values, as well as current values, of its theoretical determinants.

[15] The bias in the estimate of λ will be present if the "true" relationship is of the form

$$y_t = \gamma\beta x_t + (1-\gamma)y_{t-1} + u_t,$$

and the u_t are serially correlated.

[16] See Griliches' note [7] on serial-correlation bias in estimates of distributed lags, and also his recent survey article [8], for a discussion of these points. Procedures have been suggested for obtaining a consistent estimate of the coefficient on the lagged dependent variable when autocorrelation is present and major misspecification errors have been avoided. See, for example, the references cited in Griliches [8, pp. 40–42] and Wallis [26]. Still, no technique of estimation, however sophisticated, can give valid results if the specification of the desired stock is itself seriously incorrect.

[17] See also Tinsley [22], [23].

degree of the polynomial and the length of the lag are preselected by the researcher. Here we do not report any equations which make use of the technique of polynomial approximation, and therefore do not discuss it. Nonetheless, we believe that it is a method which ought to be given at least equal prominence with the Koyck-Nerlove Model.

3 FURTHER RESULTS

THE results of three types of test are presented below. First, we test for the existence of significant lagged responses in the borrowing behavior of Japanese foreign-exchange banks. Second, we report on a sample test of the linear homogeneity assumption. Third, we present three equations estimated on data from the 1964–68 period only.

A. LAGGED RESPONSES

The equations reported in our earlier work [6] and summarized in Table 1 assumed that the response to a disturbance was completed in the quarter following the disturbance. Table 2 contains eight equations testing for longer responses. In equations (2.1) and (2.2), current and lagged-one-period values of the imports–net-worth ratio, and the spread between the Japanese lending rate and the American borrowing rate, are employed as regressors. Equation (2.2) differs from (2.1) in that the autoregressive parameter is estimated. Both of these equations suggest that there is a lagged response to changes in interest rates, but not apparently to changes in the import–net-worth ratio. Another finding of interest is the small estimated impact of the VFCR variable; the current estimates are only about a third of those in equations (1.1)–(1.4).

Equations (2.3) and (2.4) are partial-adjustment equations using the form of equation (11′). Equation (2.4) differs from (2.3) in that the lagged, as well as current, value of the interest-rate spread is employed as a regressor. These equations are very similar. The only meaningful difference seems to be that the current interest-rate coefficient in (2.3) is divided between the current and lagged coefficients in (2.4). While

TABLE 2

Lagged Adjustment

Equation	BR	CW	V	Constant	$RLld$ minus $RBus$	$(RLld$ minus $RBus)_{-1}$	M/NW	$(M/NW)_{-1}$	γ	ρ	SEE	DW
(2.1)[a]	−.365	−.420	−.064	.004	.091	.267	.680	−.099	—	—	.121[a]	.66
	(.039)	(.029)	(.118)	(.376)	(.100)	(.096)	(.183)	(.191)			(.099)	
(2.2)[a]	−.333	−.348	−.083	.043	.150	.116	.735	−.059	—	.99	.101[a]	2.08
	(.187)	(.177)	(.105)	(1.185)	(.082)	(.091)	(.159)	(.135)			(.083)	
(2.3)	−.298	−.462	0	−1.010	.423		1.026		.322	—	.073	1.79
	(.110)	(.075)	(.064)	(.727)	(.058)		(.369)		(.071)			
(2.4)	−.300	−.464	0	−1.011	.277	.140	1.021		.365	—	.073	
	(.105)	(.067)	(.066)	(.654)	(.162)	(.148)	(.330)		(.087)			
(2.5)	−.246	−.442	−.032	−.205	.303	.169	1.376	−.810	.379	.20	.067	1.94
	(.139)	(.069)	(.063)	(.631)	(.163)	(.139)	(.425)	(.379)	(.100)			
(2.6)	−.302	−.456	−.031	−.341	.359	.121	1.531	−.902	.351	—	.064	1.68
	(.093)	(.060)	(.061)	(.586)	(.160)	(.137)	(.421)	(.379)	(.078)			
(2.7)	−.213	−.425	−.033	.004	.234	.215	1.152	−.675	.467	.5	.067	2.25
	(.177)	(.088)	(.061)	(.642)	(.128)	(.112)	(.341)	(.328)	(.110)			
(2.8)	−.187	−.380	−.035	.166	.185	.220	.990	−.555	.549	.7	.070	2.33
	(.218)	(.121)	(.062)	(.640)	(.110)	(.100)	(.293)	(.299)	(.130)			

NOTE: See Table 1 for measurement of variables. Period of estimation: 1959–67.

[a] Since the dependent variable in these equations is *Bus/NW*, while that in the remainder of the equations is *Bus*, the SEE's of these equations are not comparable to those of the others. The product of the SEE's of these equations and the mean value of *NW* is somewhat comparable to the SEE's of the other equations. Thus it is reported in parenthesis below the SEE.

the standard errors of the equations are roughly 10 per cent below that of equation (2.2), the equations exhibit a theoretically undesirable property: the response of borrowing to a change in imports is estimated to be more than one for one. The short-run response is almost exactly one for one, while the long-run response (the short-run response divided by the estimated speed of adjustment) is three times the change in imports. Since the rationale behind the variable is that import trade bills can be used as collateral, a greater than one-for-one relationship between borrowing and imports seems highly implausible.

Other somewhat surprising results are the zero coefficient on the VFCR variable (it was constrained a priori to be nonpositive) and the relatively low estimate of the speed of adjustment.[18] The latter implies that it takes eight quarters for 95 per cent of the adjustment to occur. As we noted above, the estimate is biased downward to the extent that, say, the exclusion of relevant explanatory variables from the equation tends to introduce autocorrelated residuals into that equation.

In an attempt to remove possible downward bias in the estimate of the speed of adjustment, we have estimated a partial-adjustment equation including the first-order autoregressive parameter. More specifically, the equation is of the form

$$Bus = \alpha\left[\hat{\gamma}Bus^* + (1-\hat{\gamma})\left(\frac{Bus}{\alpha}\right)_{-1}\right] + \hat{\rho}\left\{Bus_{-1} - \alpha_{-1}\left[\hat{\gamma}Bus^*_{-1} + (1-\hat{\gamma})\left(\frac{Bus}{\alpha}\right)_{-2}\right]\right\} + \epsilon, \quad (12)$$

where $\hat{\gamma}$, $\hat{\rho}$, and the parameters in α and Bus^* are all estimated, and where ϵ is a disturbance term assumed to have the customary desired properties.[19] To the extent that the estimate of the speed of adjustment increases, this inclusion is likely to reduce the estimate of the long-run

[18] The "true" speed of adjustment, λ (see page 217, note 12), is only .38 in equation (2.3).

[19] Equation (12) is derived by assuming that the error term in direct estimation of (11') is first-order positively autocorrelated. In other words, with u_t as the error term in (11'), we assume

$$u_t = \rho u_{t-1} + \epsilon_t$$

where $0 < \rho \leq 1$ and ϵ_t is a disturbance term that is not serially correlated. Lagging equation (11') one period, multiplying through by ρ, and subtracting the resulting equation from (11') yields equation (12), in which the disturbance term is ϵ_t.

response of borrowing to a change in imports. Another possible means of reducing the long-run value of this estimate is to employ the lagged imports–net-worth ratio as an explanatory variable. Since the lagged ratio assumed a negative coefficient in equations (2.1) and (2.2), it would appear that the response to changes in imports is faster (possibly even instantaneous) than the response to changes in interest rates.

Equation (2.5) suggests that these modifications, while having the desired influence, are not enough. The autoregressive parameter assumes a low value of 0.2, and its inclusion does not seem to have significantly raised the estimate of the speed at which banks respond to changes in interest rates (compare (2.5) with (2.4)). However, the combination of this parameter and the lagged-imports variable provides estimates that banks adjust very rapidly to changes in imports; the long-run partial of borrowing with respect to imports, 1.49, only slightly exceeds the short-run partial, 1.38.

The impact on these partials of including the autoregressive parameter is shown most clearly by comparing equation (2.5) with (2.6). The short-run partial implied by the latter exceeds that implied by the former by 10 per cent, and the difference in long-run partials exceeds 20 per cent. Even though including the autoregressive parameter worsens the equation in the sense that it raises the standard error of estimate (from .064 to .067), the more plausible (lower) estimates of the partials with respect to imports are enough to lead us to prefer equation (2.5) to (2.6). In fact, we are willing to trade off additional explanatory power in order to obtain more reasonable estimates of the import partials. Thus, a priori, we have estimated the basic equation constraining the autoregressive parameter to assume successively higher values.

Equations (2.7) and (2.8) are sample results. As was expected, increasing the value of ρ lowers the estimates of both the short- and long-run partials with respect to imports. In fact, equation (2.8) implies that American banks initially finance all of Japanese imports, but that perhaps a fifth of imports are eventually financed in some other manner; that is, imports act, at least partially, as an "impact" variable [11, pp. 48–49]. Also, as expected, the estimate of the time taken for adjustment diminishes. This estimate, in conjunction with the interest-rate coefficients, implies that 66 per cent of the eventual response to changes in interest rates occurs within two quarters, 85 per cent occurring within three. A much slower response in the adjustment of short-term portfolios by financial institutions would seem unlikely. At the same time, very rapid adjustment would be surprising, given the lack of a secondary market in bank loans.

To summarize, at the cost of a 10 per cent fall in the standard error of our equation (compare (2.6) and (2.8)) we have obtained theoretically acceptable estimates of the response of borrowing to changes in imports, and we have estimated what we consider to be a plausible time-response of borrowing to changes in interest rates.

Since equation (2.8) is our preferred equation, it might be useful to discuss some of its implications. The principal difference between it and the "best" equation reported in [6] (equation (1.2) in the present paper) is the largeness of the long-run interest-rate elasticities. By ignoring lags in our earlier paper, we forced the long-run elasticities to equal the short-run (one quarter) elasticities. In fact, the former (−1.9 for the United States rate, and 2.5 for the Japanese lending rate) seem to be about four times the latter.[20] The estimates from equation (2.8) imply that a 50 basis-point rise in the American rate would tend to reduce borrowing, at the end of 1967 values of *Bus* and α, by \$140 million, \$220 million, and \$100 million for the current and two future quarters, respectively. Lastly, the VFCR program seems to have mattered little. The estimates in equation (2.8) imply that the end of 1967 Japanese short-term borrowing was only \$25 million less than it would have been in the absence of the VFCR program. This estimate is substantially less than the \$160 million implied by equation (1.2).

B. THE LINEAR HOMOGENEITY ASSUMPTION

Linear homogeneity in the scale variable (or in all dollar magnitudes) has been assumed in recent work on international capital flows by Lee [15], Miller and Whitman [17], and ourselves [6].[21] In none of these studies has the assumption been explicitly tested, although it

[20] For a discussion of the mechanics of this and the other calculations in this paragraph, see [6, pp. 56–57].

[21] This has been a common assumption in empirical work on domestic financial behavior for some time.

was implicitly tested by Miller and Whitman.[22] We report here a test of our assumption that Japanese short-term borrowing from American banks is homogeneous of degree-one in Japanese imports and the net worth of Japanese foreign-exchange banks.

Given the use of a nonlinear regression program, testing the linear homogeneity assumption would appear to be a simple matter. Equation (2.1) was of the basic form

$$\frac{Bus}{NW} = \alpha \frac{Bus^*}{NW} NW^{\phi}, \tag{13}$$

where $\phi = 0$. To test the homogeneity assumption, we simply estimate ϕ. A value of ϕ insignificantly different from zero would support the homogeneity assumption. Equation (2.2) is of the same general form, but with the autoregressive parameter; thus, the homogeneity assumption could be tested in the same manner. Equation (2.8) was of the form

$$\begin{aligned} Bus = \alpha \left[\gamma \frac{Bus^*}{NW} NW^{\phi} + (1-\gamma)\left(\frac{Bus}{\alpha}\right)_{-1}\right] \\ + .7 \left\{Bus_{-1} - \alpha_{-1}\left[\gamma\left(\frac{Bus^*}{NW}\right) NW^{\phi}_{-1} + (1-\gamma)\left(\frac{Bus}{\alpha}\right)_{-2}\right]\right\}, \end{aligned} \tag{14}$$

where $\phi = 1$. This time an estimate of ϕ insignificantly different from unity would support the homogeneity assumption.

As simple as these tests appear, the large number of intricately

[22] Miller and Whitman's equation (18) is of the form

$$\log K = .8447 \log A_1 + .7245 \log W_{t-1} + \cdots,$$

where K is the stock of foreign portfolio assets held by American residents, A_1 is the scale variable (K plus various domestic long-term portfolio assets), and W is the ratio K/A_1. The short-run elasticity of K with respect to A_1 is .8447, not far below unity. To obtain the long-run elasticity, we express W_{t-1} in terms of K and A_1. Slight manipulation yields

$$\log K = .7245\ \Delta \log A_1 + .1202 \log A_1 + .7245 \log K_{-1} + \cdots.$$

The estimated long-run elasticity is $.1202/(1.0 - .7245) = .44$. Thus, it appears from this equation that A_1 acts partially as an impact variable; increases in A_1 temporarily raise K substantially above its new equilibrium level. (At first glance, this seems somewhat implausible.) Further, the estimated long-run elasticity appears to be substantially less than unity.

related parameters to be estimated seemed to raise insurmountable difficulties for our nonlinear regression program.[23] However, equation (13) has been approximated by iterating on the parameter ϕ.[24] When values of ϕ in the ± 1.0 range were selected, the equation with $\phi = 0.01$ yielded the lowest SEE. This equation is

$$\frac{Bus}{NW} = \underset{}{(1.0} - \underset{(.039)}{.366}BR - \underset{(.029)}{.418}CW - \underset{(.118)}{.069}V)$$

$$\Big[\underset{(.377)}{-.018} + \underset{(.100)}{.092}(RLld - RBus) + \underset{(.096)}{.270}(RLld - RBus)_{-1}$$

$$+ \underset{(.183)}{.688}\,\frac{M}{NW} - \underset{(.191)}{.101}\left(\frac{M}{NW}\right)_{-1}\Big] NW^{.01}. \qquad (13')$$

$$SEE = .121274. \ DW = 0.67.$$

The closeness of this estimate of ϕ to 0.00 and the minor reduction of the SEE (compared with .121280 in equation (2.1)) lends some support to our use of the linear homogeneity assumption in [6] and in this text.

C. 1964–68 ESTIMATES

We test the stability of the regression estimates over time by estimating equations on data from the 1964–68 period. These equations could conceivably provide a more accurate estimate of the basic behavioral relationship, and of the impact of the American Voluntary Foreign Credit Restraint program, since the data for this period were not affected by either the relaxation of Japanese restrictions or the responses which ensued.

Equations (3.1)–(3.3) in Table 3 (which are analogous to equations (2.1), (2.2), and (2.8)) are based on data from the 1964–68 period. They imply a 25–50 per cent smaller total borrowing response to changes in imports than do the equations based on the total 1959–67 period. This

[23] For reasons that are not yet clear to us, the program failed to iterate away from the initial guesses of the parameters.

[24] Even with ϕ specified a priori in equations of the form of (2.2) and (2.8), the program failed to iterate away from the initial guesses of the other parameters.

TABLE 3

Tests on

Equation	Constant	*V*	*RLld* minus *RBus*	(*RLld* minus *RBus*)$_{-1}$
(3.1)	.143	−.146	.155	.299
	(.464)	(.102)	(.107)	(.100)
(3.2)	.307	−.025	.148	.117
	(.556)	(.089)	(.077)	(.078)
(3.3)	.572	0	.329	.111
	(.947)		(.140)	(.116)

NOTE: *Bus/NW* is the dependent variable in equations (3.1) and (3.2); *Bus* is the dependent variable in (3.3). See Table 1 for measurement of var-

is consistent with our expectation that the bias of American lenders in favor of acceptances based on import trade bills would decline during the period as the Japanese established their credit-worthiness. Regarding interest-rate responses, the total responses implied by equations (3.1) and (3.3), respectively, are about 25 per cent greater than those implied by their 1959–67 counterparts. The coefficients in (3.2), however, are virtually identical to those in (2.2). The stock-adjustment parameter in (3.3) is noticeably (although probably not significantly) lower than in equation (2.8) (.44 compared with .55), and the time pattern of interest-rate responses is also different. These two changes offset each other to some extent, so that equation (3.3) implies an adjustment to interest-rate changes that is almost as rapid as that implied by (2.8).

These estimates reinforce our earlier finding that the impact of the VFCR program on American short-term bank lending to Japan has apparently been negligible. The estimate of the *V* coefficient in (3.1) is substantial, but that in (3.2) is negligible, and that in (3.3) has to be constrained from assuming a positive value. Since equation (3.1) is probably the least reliable of the three, we conclude that the impact of the program has been slight.

1964–68 Data

M/NW	$(M/NW)_{-1}$	γ	SEE	ρ	DW
.758	−.343	–	.111	–	1.04
(.221)	(.233)				
.509	−.039	–	.086	.87	1.60
(.180)	(.161)				
.804	−.617	.439	.075	.70	2.60
(.380)	(.411)	(.141)			

iables. The variables *BR* and *CW* are equal to zero throughout the period.

4 ALTERNATIVE SPECIFICATIONS

IN THE course of our research, extending over several years, we have more than once revised our notion of the proper theoretical and empirical specification of international capital demand and supply equations. To put it more bluntly, we have estimated a number of equations that, in retrospect, were poorly specified. Many of these equations were, of course, patterned after existing work in the literature. The fact that these earlier equations often yield at least superficially plausible results while improperly specified in important respects has aroused our curiosity.

We will now put forward some examples of alternative specifications for equations purporting to explain Japanese short-term borrowing from the United States. In every case we consider these alternative specifications inferior to those already discussed—inferior in the sense either that they are less acceptable on theoretical grounds, or that they seem less likely to provide valid empirical approximations to the underlying behavioral relationship. In several instances we deliberately employ obvious misspecifications. Our purpose in presenting these additional equations is to shed some light on a general question of interest to all researchers in this and other fields: How sensitive are one's conclusions to the equation specification employed?

A. TREATMENT OF CAPITAL CONTROLS

We have argued in [6] that there are probably few countries in which changes in governmental restrictions on capital flows have been negligible enough to be ignored. The importance of these controls in the case of the particular capital flow examined here is incontrovertible. It is not surprising, therefore, that a specification of the relationship determining Japanese borrowing from American banks which completely ignores these controls is incapable of providing meaningful estimates of the effects of the economic determinants. Equation (4.1) in Table 4 is an estimate of such an equation; it is identical to equation (1.3) except for the fact that α has been set equal to unity throughout the entire 1959–67 period. The coefficient on the import ratio is unexpectedly negative; the coefficient on the interest-rate spread, while positive, is less than its standard error; and the explanatory power of the equation is exceptionally low.

It is somewhat surprising, however, that simply adding a time trend[25] and a dummy variable to reflect the American VFCR program[26] yields results which are superficially plausible. Equation (4.2) includes a time trend[27] T_{68} and (4.3) includes a VFCR dummy variable as well. The latter variable is zero until the second quarter of 1965; in that and subsequent quarters, it is unity. The time trend greatly improves the equation. The standard error of estimate is nearly halved; the coefficient on the import ratio changes to the correct sign and is

[25] For examples of the use of a time trend in capital-flow equations, see [5], [18].

[26] For examples of the use of dummy variables to capture the impact of the United States Interest Equalization Tax and the VFCR program, see [5], [17], [19]. This method will not yield accurate estimates except in the unlikely event that the impact of the programs are constant over time.

[27] The trend variable T_{68} is equal to unity in the first quarter of 1959 and rises by 1 each quarter throughout the entire 1959–67 period.

TABLE 4

Estimates Ignoring Capital Controls

Equation	Constant	*RLld* minus *RBus*	*M/NW*	T_{68}	T_{64}	D_V	SEE	*DW*
(4.1)	2.448	.123	−.419				.616	.11
	(1.013)	(.139)	(.560)					
(4.2)	−3.236	.669	.576	.087			.323	.47
	(.812)	(.094)	(.313)	(.009)				
(4.3)	−2.397	.518	.440	.095		−.596	.297	.49
	(.814)	(.104)	(.292)	(.009)		(.230)		
(4.4)	−2.451	.365	.620		.116	−.057	.151	.83
	(.405)	(.051)	(.150)		(.005)	(.111)		

NOTE: *Bus/NW* is the dependent variable. See the text for the definition of variables and Table 1 for their measurement. Period of estimation: 1959–67.

nearly twice its standard error; and the coefficient on the interest-rate spread is seven times its standard error. While the inclusion of a time trend can presumably be rationalized as an attempt to mitigate the influence of excluded variables, such as Japanese capital controls, on the estimates, it is still surprising to us—and also disconcerting—that the trend variable seems to perform so well in a case where the excluded variables are clearly so important. Inclusion of the VFCR dummy also appears, superficially, to be a success. Its coefficient in equation (4.3) is negative, as expected, and it is twice its standard error. Moreover, this inclusion lowers the standard error of estimate of the equation by nearly 10 per cent.

In comparison with the equations in Tables 1 and 2, however, equation (4.3) does not stand up well. Its SEE is more than twice that of its closest analogue, equation (1.3). Further, the large coefficient on the VFCR dummy variable in equation (4.3) suggests that the program has had an implausibly large impact on bank lending to Japan. This equation implies that lending was approximately $1 billion less at the end of 1967 than it would otherwise have been. In contrast, equa-

TABLE 5

Equations with Misspecified

Equation	*RLId*	*RBus*	*RBe$*	Constant
(5.1)	−.151	−.108		1.641
	(.067)	(.036)		
(5.2)		−.205	.136	.291
		(.071)	(.065)	(.163)
(5.3)	−.003	−.098		.371
	(.057)	(.026)		(.501)
(5.4)	−.048	−.102		.621
	(.095)	(.035)		(.894)
(5.5)	−.057	−.107		.919
	(.058)	(.029)		(.495)

NOTE: The change in *Bus* is the dependent variable. See Table 1 for measurement of variables. Period of estimation: 1959–67.

tion (1.3) implies that the impact was only $160 million; and in numerous equations, the impact has been estimated to be negligible.

Equation (4.4) illustrates that a more plausible estimate of the impact of the VFCR program can be obtained even within the current simplistic framework. The equation differs from (4.3) only in that a truncated time-trend, rather than the "complete" trend, appears as a regressor. The truncated-trend variable rises by 1 each quarter through the first quarter of 1964 and thereafter remains at that level. The use of this variable might be rationalized along somewhat the same lines as we rationalized the *BR* and *CW* variables in our earlier research.[28] This switch alone slashes the SEE by half and yields an estimate of the impact of the VFCR program of only $90 million. The important point is that without the earlier equations reflecting the explicit consideration of Japanese capital controls, one might have been convinced that the VFCR program had a substantial impact on bank lending to Japan.

[28] See [6, pp. 39–44].

Desired Stock

ΔT_{64}	ΔNW	ΔM	M	SEE	DW
.027				.094	1.96
(.049)					
.047				.095	1.48
(.053)					
.043	2.209	.542		.069	2.05
(.041)	(.854)	(.104)			
.091			.102	.093	1.85
(.065)			(.069)		
−.004		.468		.076	1.56
(.040)		(.109)			

B. MISSPECIFICATION OF THE DESIRED STOCK

The most flagrant possible misspecification of the desired stock consists of relating capital *flows* to the *levels* of interest rates and nothing else. This framework combines the infamous "flow theory" of capital movements with an absence of theory on the scale of economic units. At a lower order of magnitude, errors of specification can be committed in the choice of interest rates and the scale variable.

Table 5 contains five misspecified equations, each having the change in American short-term bank lending to Japan as the dependent variable. In an attempt to isolate misspecifications of the desired stock from the problems of measuring the impact of capital controls, all equations contain the change in the truncated-trend variable.[29] The effect of including the truncated trend is to allow the constant term in the flow equations to be larger in the 1959-I–1964-I period than in the later period.

[29] It was established in the previous section that the truncated-trend approximately accounts for the impact during the 1959–64 period of Japanese controls on the level of bank lending. The closeness of the coefficient of the VFCR dummy variable to zero in equation (4.4), and the equations in Tables 2 and 3, suggest that omission of variables representing American controls will not seriously alter the results.

All of the equations in the table include the levels of interest rates, rather than the increments, as regressors. Given this basic misspecification,[30] we have not attempted to document in any systematic manner the effect of using the wrong set of rates as regressors; in all equations except one we have employed the two theoretically most important rates – the cost of borrowing in the United States, and the return from lending in Japan. In general, earlier researchers have only considered the yields on the "own" and substitute instruments. In our case this would mean including only borrowing rates (i.e., the cost to Japanese banks of borrowing in the American, Japanese, and Eurodollar capital markets). A more fundamental error is to include rates prevailing in international financial centers only.[31] This would require the exclusion of all Japanese interest rates from the set of regressors.

Turning to Table 5, equations (5.1) and (5.2) exclude all scale variables.[32] They differ only in that the former includes the Japanese lending rate, while the latter includes the Eurodollar borrowing rate. Quite deceptively, even though interest rates are incorrectly measured as levels rather than increments, the United States and Eurodollar rates are statistically significant, with the expected sign. The Japanese lending rate, however, is significant with the unexpected sign.

Equation (5.3) includes the net-worth scale variable and the import-distribution variable. Both variables enter significantly, substantially reducing the standard error of estimate. In addition, they eliminate the misspecified Japanese lending rate that was entering with the unexpected sign.

Another misspecification suggested by some of the earlier literature is the use of the level (rather than increment) of imports in an equation where the dependent variable is *changes* in borrowing.[33] If the level of Japanese borrowing is actually related to the level and increment in imports, as equation (2.8) suggested, the change in borrowing should be related to the first- *and* second-difference in imports, hardly the level. Equation (5.4) is the misspecified equation; (5.5) is

[30] Studies containing this misspecification include [2], [4], [9], [13], [14], [18], [19], [20], [21]. See also the Dobell-Wilson chapter in this volume.

[31] See [21] for an example of this error.

[32] Investigations ignoring scale variables altogether include [2], [4], [18], [20], [21].

[33] See [14] for an example. More generally, this misspecification would relate capital flows to the level of export or import "distribution" variables.

reported for purposes of comparison. As expected, the change in imports performs much better than the level of imports.

The equations listed in Table 5 are disturbing in an important respect. The standard errors of estimate of these equations compare favorably with those reported earlier. The relevant comparison is probably with the SEE of equation (1.4), appropriately adjusted. (Since the dependent variable in (1.4) is the ratio of borrowing to net worth, the SEE of that equation should be multiplied by .82, the mean value of net worth, to make it roughly comparable to those of the equations in Table 5. Such a multiplication yields a value of .086.) As can be seen, the SEE's of equations (5.3) and (5.5) are, in fact, below that of (1.4). Disconcertingly, the SEE of (5.3) is below even that of our preferred equation, (2.8). A corollary to the low SEE's is, of course, the fact that the misspecified interest-rate variables are often quite significant. A partial explanation for the relatively low SEE's in Table 5 is that all of the equations reported earlier employed interest-rate differentials. The use of differentials will raise the SEE when the individual rates would otherwise tend to assume coefficients that are much different in absolute magnitude; such is clearly the case with our data [6, Tabie 5].

While the equations in Table 5 explain the data quite well in a purely statistical sense, they all imply that the long-run elasticity of bank lending to Japan with respect to interest rates is infinite (plus or minus). Such an inference is, of course, inconsistent with observed diversified asset-and-liability portfolios.

C. SOME PARTIAL ADJUSTMENT EQUATIONS

In Part 2 we emphasized some problems associated with the partial-adjustment model. More specifically, we noted that "even though the determinants of the 'desired' stock (or α) may be seriously misspecified, the stock-adjustment model can generate superficially plausible estimates." In addition, the estimate of the speed of adjustment in such an equation is almost certain to be biased downward. In Table 6, we report four "typical" partial-adjustment equations. The principal misspecification of the equations is the complete absence of

TABLE 6

Misspecified Partial-

Equation	Constant	*RLld*	*RBus*	*M*
(6.1)	.766	−.069	−.090	.223
	(.755)	(.089)	(.036)	(.126)
(6.2)	.194		−.089	.255
	(.123)		(.031)	(.057)
(6.3)	1.102	−.113	−.089	.169
	(.792)	(.095)	(.035)	(.132)
(6.4)	.160		−.083	.268
	(.129)		(.032)	(.059)

NOTE: *Bus* is the dependent variable. See Table 1 for measurement of variables. Period of estimation: 1959–67.

any variables reflecting the impact of Japanese capital controls ($\alpha = 1$ throughout).

Equation (6.1) contains all the determinants of the desired stock —the principal borrowing and lending rates, net worth, and imports— and a dummy variable purporting to capture the impact of the American VFCR program. According to some standards, the equation looks "reasonably" satisfactory. All coefficients have the expected signs except the coefficient on the Japanese lending rate and the practically zero coefficient on net worth; the coefficients on both the borrowing rate and the dummy variable are significantly less than zero at the .05 level, while that on imports is significantly greater than zero. In addition, the standard error of estimate of the equation is respectable; for example, it is lower than that of the "best" equation published in our earlier paper (.086 on a comparable basis). Equation (6.2), which does not include the variables that were statistically insignificant in (6.1), looks even better. The regression coefficients are more significant, and the SEE is marginally lower.

On closer inspection, however, these equations exhibit a number of disturbing characteristics. First, with respect to imports, the equations imply that the long-run partial derivative substantially exceeds 2.0. This result seems economically implausible. Second, the equations

Adjustment Equations

NW	Dv	Dv_{-1}	Bus_{-1}	SEE	DW
−.012	−.138		.904	.082	2.11
(.303)	(.074)		(.045)		
	−.125		.892	.080	1.97
	(.058)		(.034)		
.152	−.068	−.133	.885	.081	2.15
(.326)	(.088)	(.104)	(.047)		
	−.065	−.080	.885	.080	1.90
	(.062)	(.090)	(.035)		

suggest that the VFCR program succeeded in reducing lending to Japan by more than $1¼ billion by the end of 1967. We noted earlier that an estimate as large as this is implausible. Third, the estimated speed of adjustment of these equations is less than three-tenths of the already slow speed of adjustment estimated in equation (2.3), a typical partial-adjustment equation including variables reflecting Japanese capital controls. This is, of course, exactly what was anticipated.

Equations (6.3) and (6.4) reflect an attempt, within the limiting constraints imposed by the employment of an on/off dummy variable, to obtain a more plausible estimate of the impact of the VFCR program. Not only is the estimate of the total impact of the program in equations (6.1) and (6.2) implausible, so also is the timing of the impact. The equations imply that 10 per cent of the impact was felt during the quarter the program was imposed, 9 per cent the next quarter, 8 per cent the following quarter, and so on. This obviously unacceptable result simply reflects the low estimate of the speed of adjustment, and the fact that the typical partial-adjustment equation forces the speed of adjustment in response to all variables to be identical. In general, one would probably expect the imposition or removal of any government restrictions to have a relatively rapid impact. To test for a more rapid impact of the VFCR program, we include the lagged value of the dummy variable in the equation. If the lagged value were to assume a coefficient with sign opposite to that of the current value, the effect would be to reduce the lag. However, since the lagged coefficient is also negative, equations (6.3) and (6.4) suggest an even slower response to the imposition of the controls than to changes in the underlying economic determinants.[34]

The fact that equations ignoring Japanese capital controls—probably the most important determinant of bank-lending to Japan during the period—look reasonably appealing, even at first glance, is disturbing. It is also worrisome that such equations might lead policymakers to overestimate greatly the impact of the VFCR program, and to underestimate the impact of interest-rate changes in the short run (due to the low estimate of the speed of response). If these results offer any general guide to problems encountered with other capital flows, and we suspect that they do, estimators and users of partial-adjustment equations would be well advised to proceed cautiously.

5 CONCLUDING NOTE

OUR objectives in this paper and in [6] have been primarily methodological. Working intensively with a single set of data, we have tried to pose, and to resolve as adequately as possible, many of the theoretical and econometric problems arising in the empirical analysis of all international capital flows.

As we noted in our earlier paper, our research has not led us to an optimistic assessment of the ease with which valid substantive conclusions can be reached in this field. The difficulties to which we have drawn attention are serious and cannot be easily overcome. Although we doubt that useful empirical knowledge of the behavior relationships determining capital flows can be acquired rapidly and at small cost, we

[34] An example of the successful measurement of a more rapid response to changes in government regulations than to changes in the economic determinants is given in [11]. Sixty per cent of the response of the commercial bank time-deposit rate to changes in the ceiling rate on time deposits is estimated as occurring in the first quarter, while only about 13 per cent of the adjustment to changes in economic determinants seems to occur within this time limit.

are not, on the other hand, so pessimistic as to want to discourage econometric research in this area. In our view, substantial research efforts are warranted, simply because many important problems of domestic and international financial policy cannot be dealt with wisely without a much better quantitative grasp of the determinants of capital flows.

It seems appropriate to conclude with a comment of a general nature about the priorities that we feel ought to be observed in future research. There is no doubt in our minds that the major investment of resources in future research should go, first, into the judicious selection of the particular capital flows to be studied; second, into the development of well thought out theoretical specifications appropriate in the particular circumstances; and, third, into the careful collection of high-quality data. The actual estimation of many additional equations for many types of capital flows should have a lower priority. Theory and techniques are both still in a relatively primitive state and reliable data are scarce. In this situation, the need is for intensive research, not for extensive application of existing (but meager) knowledge. In any empirical work, of course, one should always formulate a convincing theoretical framework, developing strong opinions on the type of results that would be theoretically acceptable before estimating any equations. This principle, honored more in the breach than in the observance in the existing literature (we cannot claim purity on this score ourselves), should be adhered to more strictly if it is to serve as a practical guideline for future empirical work on capital flows.

The pitfalls that lie in wait for the researcher unarmed with strong a priori theoretical views have been amply illustrated here. Many different specifications for a relationship—with widely varying implications—may, superficially, seem to work. Our experience with the data that we have been discussing certainly suggests that frequently one may be unable to discriminate clearly, in purely statistical terms, between alternative imperfect specifications. Fishing expeditions are virtually bound to be "successful" if a researcher is satisfied with merely finding some specification that will give a good statistical fit. If it does nothing else, this paper ought to underline the need for a much more robust definition of "success" than the customary tests of statistical significance. One well researched relationship based on prior development of a sound theoretical framework, and on careful matching of empirical counterparts to theoretical constructs, will usually prove more useful and interesting than scores of equations with good fits but weak theoretical underpinnings.

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COMMENTS

STANLEY W. BLACK

PRINCETON UNIVERSITY

Professors Leamer and Stern regard the portfolio-adjustment model of capital flows as a logical development of the older *activities* framework, which used the concepts of speculation, arbitrage, and trade flows of short-term capital. While they apparently regard the portfolio model as an improvement, they sound a useful warning of the limitations of the model. Not only does the portfolio model neglect changes in net worth, it also ignores liquidity considerations. However, changes in net foreign worth of a country are easily allowed for, since they are identical with net exports. Leamer and Stern's strictures against the portfolio model for nonfinancial corporations seem well taken.

The authors suggest that the *stock* of short-term capital be related to covered interest differentials, speculative variables, and *changes* in exports or imports as measures of changes in sales. This last suggestion disagrees with the practice advocated by Bryant and Hendershott, as well as by Branson: use of the *change* in exports or imports in capital-*flow* equations, or the *level* of exports or imports in *stock* equations. Furthermore, the use of changes in sales to explain the stock of a financial asset goes against the logic and history of the treatment of stocks and flows. Equilibrium stocks can logically be related to both

NOTE: The author has accepted a position at Vanderbilt University.

the level *and* change of an appropriate flow variable, but not to the change in flows alone.

I am particularly pleased that Leamer and Stern raise the simultaneous-equations problem in regard to exports and imports, a topic which I will discuss later. One point of contention is that the portfolio-adjustment equations *can* be modified to take account of short-run cash-flow constraints, as de Leeuw did for the monetary sector of the Brookings Model.

Concerning the treatment of speculative periods, I have two suggestions. Dropping out speculative periods throws away valuable information on demand functions. An alternative to the Stein approach to residuals is to estimate a function in a nonspeculative period and then calculate the residuals from that function *for the speculative period only*. The implied expectation can then be used to estimate other functions for the speculative period. I have done this with some success for the Canadian devaluation. Dummy variables can also be used with more imagination, especially if the data are for short time periods, such as weekly data. An exponential distribution of speculative impact suggests itself on several grounds, including Muth's "rational expectations." Thus, an unanticipated disturbance would have impact $1, \lambda, \lambda^2, \ldots,$ while an anticipated disturbance would have impact $\ldots \lambda^2, \lambda, 1, \lambda, \lambda^2, \ldots$.[1]

I found inappropriate Leamer and Stern's conclusion that disaggregation should be pursued only if it offers improved estimates of the *aggregate* relationship. It seems to me that economists should be willing to disaggregate as long as different *micro*behavior patterns turn up. Bryant and Hendershott give us a particularly good illustration of the value of disaggregation.

On multicollinearity, Leamer and Stern confuse me. This is not a case of dropping variables with t-ratios less than two. Rather, if A is dropped, B is significant; and conversely. Thus, as they say, neither coefficient is estimated accurately, but some linear function of the coefficients is. If A is left out, for example, the coefficient of B measures the effect of *both* A and B together. In that sense, A has not really been discarded.

[1] See my paper, "The Use of Rational Expectations in Models of Speculation, *Review of Economics and Statistics*, Vol. LIV, No. 2 (May, 1972).

Concerning the simultaneous-equations problem, Leamer and Stern remind us that exports and imports may be simultaneously determined with capital flows. However, as the paper progresses, they appear to be saying that specification of the model is a more important problem than simultaneity. I believe that *understanding* the simultaneity is one of our most pressing problems of specification. It is clear that the treatment of interest rates as exogenous is not correct in some cases. In recent works on Eurodollar liabilities of the United States, I have found it essential to regard the interest rate as jointly determined, and have estimated reduced-form equations for both the rate and the liabilities. In a paper presented at this conference, Miller and Whitman take a similar approach.

Once interest rates become endogenous, it can be seen that the issue is really the interaction of the real and monetary sectors of the international economy. Current practice regards the monetary sector as being affected by changes in the real sector, but as having no converse influence. For example, monetary tightening should reduce incomes and, therefore, imports. Financing costs may be crucial at the margin of some trade decisions, interest-rate "pessimism" to the contrary. We are not ready for three-stage least squares, but our models should begin to reflect a larger view of what is jointly determined.

Leamer and Stern's compilation of empirical estimates is interesting partly for its entertainment value. We clearly have our work cut out for us in reducing the uncertainty about interest-rate and trade impacts. Unfortunately, in some places they have inserted the wrong figures for Branson's work. The monthly estimates of interest-rate effects on short-term claims and liabilities given in the original version of Table 1 allow for an increase in the Eurodollar and Canadian interest differentials, respectively, vis-à-vis the United States. However, as Branson points out, interest differentials vis-à-vis the United States are unlikely to change because foreign rates follow the American ones. Branson's Table 1, equation (10), and Table 3, equation (11), in Chapter 4, indicate impacts of \$253 million and zero for a one point change in the American rate on short-term claims and liabilities, respectively. Branson's over-all figures for interest-rate impacts given in their footnote 30 are cited incorrectly by Leamer and Stern. The correct figures are \$3.1 billion for the monthly model and \$2.5 billion for the quarterly model, including errors and omissions in the latter. The discrepancies are disturbing.

	Monthly	Quarterly
Short-term claims	$ 253 [a]	$ 468
Long-term claims	800	315
Short-term liabilities	0	260
Long-term liabilities	2,000	693
Errors and omissions	n.a. [b]	794
	$3,053	$2,530

SOURCE: Branson, Reference [5], Bryant and Hendershott.
[a] All figures in millions.
[b] Not available.

Furthermore, can one believe that impacts of this magnitude would have no influence on the real-sector exports and imports? It is also disturbing that disaggregation increases impacts markedly. Can this be related to the increased substitution possibilities allowed by disaggregation? Professors Leamer and Stern are to be congratulated for raising and discussing cogently so many issues important for econometric work on capital flows.

Ralph Bryant and Patric Hendershott discuss some of the same problems as Leamer and Stern, but with the advantage of a well-tested model and a well-understood set of data with which to demonstrate many of their propositions. The result is an especially valuable "how to do" (or "how not to do") guide for empirical workers. I will not discuss their basic model except to say that it represents a remarkable blend of theory and knowledge of institutional detail.

The model, as given in Table 1, contains no lags. After a brief discussion of the Koyck-Nerlove lag structure, the authors present some estimates allowing this type of lag structure in Table 2. Several questions should be raised about the estimates in this table. The original model was of the ratio form

$$B/NW = \alpha f(R, M/NW).$$

In equations (2.3) to (2.7), Bryant and Hendershott drop, without explanation, the ratio form for the dependent variable. They have informed me that the original equation was multiplied through by net worth and estimated so as to retain homogeneity.

The authors note the existence of bias in the Koyck-Nerlove Model with autocorrelated residuals, and in fact reject the unconstrained results from this model (equation (2.5)). Thus, their preferred equation (2.7) has a coefficient for the lagged stock that is *higher* than the unconstrained version, which was biased downward. It would have been preferable had Bryant and Hendershott written out their statistical assumptions more explicitly. Their estimating equation (10) seems to imply rather special assumptions. Using β to represent their *Bus*, they seem to assume

$$\beta_t = \alpha_t \beta_t^s + u_t, \tag{1}$$

$$\beta_t^s = \gamma \beta_t^* + (1-\gamma)\beta_{t-1}^s, \tag{2}$$

or

$$\beta_t = \alpha_t[\gamma \beta_t^* + (1-\gamma)\beta_{t-1}^s] + u_t. \tag{3}$$

Since

$$\beta_{t-1}^s = (\beta_{t-1} - u_{t-1})/\alpha_{t-1}, \tag{4}$$

we have

$$\beta_t = \alpha_t \left[\gamma \beta_t^* + (1-\gamma)\frac{\beta_{t-1}}{\alpha_{t-1}}\right] - (1-\gamma)\frac{\alpha_t}{\alpha_{t-1}} u_{t-1} + u_t. \tag{5}$$

It is this last combination-disturbance term that is assumed to be first-order autocorrelated, which must imply that u_t satifies approximately

$$u_t - (\rho + \Delta)u_{t-1} + \rho\Delta u_{t-2} = \epsilon_t, \tag{6}$$

where $\Delta = (1-\gamma)\dfrac{\alpha_t}{\alpha_{t-1}}$ and ϵ_t is a random term. Thus, u_t is approximately second-order autocorrelated, instead of first-order.

Bryant and Hendershott give an informal test of the stability of their equation in Table 3. One wonders why they did not use the Chow test for stability, which, although formulated for a linear model, should hold approximately for a nonlinear one.

I find it difficult to do more than commend the rest of Bryant and Hendershott's paper as an illuminating catalogue and demonstration of pitfalls due to improper model formulation. Everyone should examine their results on capital controls, the flow theory, the scale variable, and the insensitivity of the Koyck-Nerlove Model to specification error. It is particularly instructive that they find it impossible to distinguish statistically between the stock and flow theories of effects of interest differentials. It should be noted that Miller and Whitman, in their contribution to this volume, find theoretical basis and empirical evidence for both stock *and* flow effects. Although I am included in the list of miscreants in footnote 30 (Bryant and Hendershott), I will point out that my work contains both stock and flow components.[2] The authors are to be thanked for pursuing the implications of their data set far beyond the call of duty, to our benefit and instruction.

ELINOR B. YUDIN

INTERNATIONAL BANK FOR RECONSTRUCTION AND DEVELOPMENT

To anticipate the most probable single conclusion of this conference: capital-account transactions are amorphous phenomena that do not easily lend themselves to formal economic analysis or measurement. Only through painstaking care and precision, both of thought and of observation, can understanding of these phenomena advance. Even then, the results obtained are often replete with ambiguity. Both papers on which I shall comment provide evidence supporting these contentions. Both focus on evaluating and reconsidering current analytical approaches to capital movements. That of Professors Leamer and Stern does so in rather general terms. They correctly question—but for reasons I find incorrect—the "activity" orientation of prevalent analytical models, advocating its replacement by a "transactor" orientation. They list quite completely, and defend rather generally, their selection of explanatory variables relevant to capital-flow analysis. Finally, they consider certain of the manifold problems that

NOTE: The views expressed are those of the author. They in no way reflect those of the International Bank for Reconstruction and Development.

[2] See equation (4.11). Reference [4], Bryant and Hendershott.

arise in the process of empirical estimation. Bryant and Hendershott, because they place their study in the concrete context of United States banks' short-term lending to Japan, necessarily confront problems of estimation more directly. But, like Leamer and Stern, their primary concern lies with the fundamental theoretical analysis of capital movements and the problems encountered in econometric application of a priori theory.

There is much in both papers that is well done and stimulating. Nonetheless, I have chosen to concentrate on two points and to comment briefly on a third. First, I focus on the activity-transactor choice in orientation that Leamer and Stern raise. My second point touches on issues that are more pervasive: I draw particularly on the Bryant-Hendershott paper to question just how much—if any—progress has been made in the econometric analysis of short-term capital movements. Finally, I attempt to verbalize my uneasy reaction to the concluding note of the Bryant-Hendershott presentation.

ACTIVITY VERSUS TRANSACTOR

Leamer and Stern argue that capital-account analysis will be improved by switching to an orientation based on "transactors" (that is, households, businesses, government, and so on) from the one based on "activities" (consumption, investment, and others) that they find dominant in current work. Such shifts in perspective may jog thinking out of old ruts, thereby providing new insights. Two instances in other areas of economic analysis where changes in perspective seem particularly fruitful come to mind. Barbara Bergmann, at the December meetings of the American Economic Association, advocated an analogous, but opposite shift—*to* an activity orientation—for analysis of social, environmental, and government problems.[1] Gilbert and Kravis, in 1954, recommended the merits (albeit with reservations) of the same shift for international comparisons of national accounts.[2]

[1] B. R. Bergmann, "The Urban Economy and the 'Urban Crisis.'" *American Economic Review*, Vol. LIX (September, 1969).

[2] M. Gilbert and I. B. Kravis, *An International Comparison of National Products and the Purchasing Power of Currencies*. Paris, Organization for European Economic Cooperation, 1954.

In the context of the Leamer-Stern paper, the shift in perspective does isolate certain subtleties of capital-account analysis, thereby providing new insight. However, I question whether this particular adjustment in perspective is needed.

First, if, as the authors suggest, the activity approach gives rise to analytical flaws because there is no one-to-one correspondence between activity and transactor, how does the shift they propose avoid the same problem? Transactors frequently embark on several activities simultaneously. Stern and Leamer do recognize this difficulty: "This means, therefore, that it will be extremely difficult to construct a single model of the capital account that will capture all the structural characteristics of the different transactors." More important, simultaneous activities of a single transactor are interdependent events. This interdependence provides the authors with their strongest argument in defense of their preference for the transactor approach.

Second, if one aim of disaggregation by transactor, not activity, is to impose greater homogeneity of behavior in the categories analyzed, would this be attained? Asset preferences (and attitudes toward risk) tend to vary markedly, even among nominally identical transactors. Viewed in this light, the transactor approach may prove too aggregated for predictive purposes. In addition, homogeneity is even more difficult to obtain in an international study. One must assume that each type of transactor behaves similarly with respect to the same activity, regardless of nationality. For example, one would have to assume that an American businessman and a Japanese businessman will make identical decisions when faced with a given option. But suppose that official restrictions (or traditional behavior) differed between the United States and Japan; then the decisions of the two men would be likely to differ. Since restrictions do differ, I see no a priori reason to expect like behavior patterns for like transactors in different countries.

Bryant and Hendershott contend that, ideally, one needs different approaches for different transactors *and* different instruments (activities). In examining capital flows between two countries, their first step in disaggregation is by country. By disaggregating in this manner, they treat each nationality as a homogeneously responding group. Strictly speaking, this treatment raises questions analogous to those raised by Leamer and Stern's criterion for disaggregation. Neither assures homogeneity of behavior within each group.

Third, and basic, has the switch to a transactor orientation already occurred? It seems to me that it has. The questions raised with respect to the behavior implications of the early capital-flow models by Kenen, Stein, et al., effectively set the pendulum in motion, swinging it completely to portfolio or stock-adjustment models. But a portfolio is a diversified collection of securities held by a single institution or investor. In consequence, *portfolio analysis already embodies a transactor orientation.*

These stock-adjustment models transpose the Tobin-Markowitz utility-maximization models, conceived originally in a national context, into an international one. In the later studies, however, several analyses do take account of Stern and Leamer's valid criticism that portfolio models do not allow for continuous-flow adjustments. Recent studies modify the Tobin-Markowitz framework, allowing both stock *and* continuous-flow adjustments. Among others, Willett and Forte,[3] Miller and Whitman,[4] and Bryant and Hendershott (in the paper now under scrutiny) have all attempted to deal with that criticism. Bryant and Hendershott comment:

> In a growing (or declining) economic world, changes in interest rates bring about both "existing stock" . . . and "continuing flow" . . . impacts on capital flows. Given a "once-for-all" change in one or more interest rates, the existing-stock effect produces capital flows that are also once-for-all in nature (a reallocation of existing portfolios), while the continuing-flow impact persists indefinitely as long as [the change in net worth is not zero].[5]

These attempts do suffer from some of the shortcomings that Leamer and Stern note. For example, to obtain this relationship in econometric form, Bryant and Hendershott invoke homogeneity assumptions that they are unable to defend rigorously within the Tobin-

[3] T. D. Willett and F. Forte, "Interest Rate Policy and External Balance." *Quarterly Journal of Economics*, Vol. LXXXIII (May, 1969).

[4] Norman C. Miller and Marina v. N. Whitman, "A Mean-Variance Analysis of U.S. Portfolio Investment." *Quarterly Journal of Economics*, Vol. LXXXIV (May, 1970). See, too, their article in this volume.

[5] See their footnote 3.

Markowitz framework. Nevertheless, there appears to be a centering of the pendulum in models that combine both stock-adjustment and continuous-flow relationships, an approach that seems curiously conformable with the very nature of the capital account.

Leamer and Stern do offer a most damaging criticism of the portfolio approach. But they do so in a brief assertion: "Some important economic transactors do not behave . . . according to portfolio-adjustment prescripts." This pronouncement clearly deserves elaboration or demonstration if their position is to prevail.

PROGRESS IN ECONOMETRIC ANALYSIS

Both of these papers demonstrate that analysis of capital accounts is advancing, however haltingly, as a result of recent theoretical developments. Bryant and Hendershott's work makes it abundantly clear that the empirical aspects of this subject confront large and (as yet) apparently immovable stumbling blocks. Their choice of explanatory variables is one of the first examples of this point. The discussion of the structural equations, particularly in their earlier paper,[6] acknowledges (as do Leamer and Stern) the influence of risk and expectation variables—the cost of forward cover, the risk of change in asset and liability prices—on exchange rates. Their estimating equations, by contrast, include no proxies for these variables. This exclusion represents a serious shortcoming of the study, limiting the possibilities of generalizing it to analyses of other capital flows.

The exclusion also stresses a general "state of the arts" problem: Given limited information, how can these variables be quantified? Nonetheless, experimentation with some proxies for them would seem a worthwhile endeavor.

Turning briefly to a few results which Bryant and Hendershott do present, several questions come to mind. They begin their paper with what they regard as a rather successful equation (1.4), modified from the earlier work in which they focused on estimating the impact of

[6] Published subsequently as: R. C. Bryant and P. H. Hendershott, *Financial Capital Flows in the Balance of Payments of the United States: An Exploratory Empirical Study.* Princeton Studies in International Finance No. 25. Princeton, 1970.

capital controls. In that equation, the key anomaly is the strength, measured by the t-ratio, shown by the United States Voluntary Credit Restraint program (V). The two additional variables which indicate the effectiveness of capital controls are the objective "basic relaxation" of Japanese restrictions between 1959 and 1960 (BR) and the more subjective "learning process" and growth of credit-worthiness of the Japanese in the American market (CW). As expected, both BR and CW effectively distort capital flows. In the equations experimenting with lags, however, the t-ratios for all variables change: that for V now suggests that the program had little impact; those for CW and, particularly, for BR decline. One suspects that if the autoregressive coefficient were set even nearer its level of the first experiments, these variables would evince no power at all. Since many people regard the Japanese case as exemplary in the use and effectiveness of capital controls, this result would be particularly disconcerting. Although the authors comment only on the change in V, it would seem that the "gain" in lags has been offset by the "loss" in capital controls. Which is better?

The "misspecified desired-stock equations" in Section 4.B and the partial-adjustment equations in Section 4.C emphasize this confusion. In Section 4.B, equations (5.1)–(5.5) employ levels, instead of increments, of interest rates, and imports as regressors, excluding all scale variables. Both these papers comment on the theoretical inadequacy of this specification. In Section 4.C, equations (6.1)–(6.4), the Japanese capital-control variables are absent. There, the strong import response and impact of the Voluntary Credit Restraint program, as well as the slow speed of adjustment, are implausible. Unfortunately, these findings are surrounded by low standard errors of estimate and regressors with appropriate signs and significant t-values. If Japanese controls are as important as the authors contend, one does wish even more that they had reconsidered their own earlier equations, particularly equation (2.8).

Other aspects of their comparison of their own work with alternative specifications based on the work of others are also intriguing. Equation (4.4), for example, yields results quite similar to those yielded by their own equation (1.3). Evidently this is because the equation (4.4) takes account of the authors' study of time patterns of specific controls. Yet the similarity between the two results, it would seem, argues better for careful thought and observation than for a particular form of equation.

It is indeed disconcerting that in the absence of Bryant-Hendershott's work, one might find these alternative specifications econometrically, if not theoretically, gratifying. Their experiments, like the discussion in the Leamer-Stern paper, should set off signals warning of the need for great caution in future work.

UNEASY REACTIONS

Bryant and Hendershott's concluding note is a forthright statement of their fundamental position. But, basically, it evokes an uneasy reaction, one that merits much more thought and discussion than is possible here.

They argue that a researcher must arm himself with "strong, a priori theoretical views." This is a valid and quite well-accepted point. But they have, at another time, come close to arguing that data are to be "forced" into the predetermined—and "correct"—theoretical mold.[7] If this is what they intend to suggest, I am uneasy.

First, there is an obvious danger: strong, a priori views can, all too easily, become armor complete with blinders instead of the first weapon of attack that the authors envisage. Second, and more fundamentally, data are malleable; they can be fitted into many molds. Economic theory is a logical construct; there is one mold for each set of initial assumptions. But there may be more than one set of assumptions—particularly where those assumptions relate to behavior. In consequence, more than one "correct" theoretical mold can be formulated. Third, and not unrelated to the first two, forcing data into *any* mold contradicts the very purpose of empirical analysis. Meaningful empirical testing requires the opposition, possibly implicit, of alternative hypotheses. The probability of accepting one hypothesis when an alternative is true—the *beta* error—plays a vital role in empirical analysis. To the extent that data are forced into a strong, a priori

[7] This position is not explicitly stated in their paper. In discussion during the conference Mr. Hendershott expressed this view. Admittedly, his views may have undergone modification since then, and Mr. Bryant may disagree—with either or both of us.

theoretical mold, the purpose of empirical testing is vitiated. The researcher who forces data into a mold says, in effect: "I'll be judge; I'll be jury."

I am far more comfortable with the authors' plea for a "more robust definition of 'success' than the customary tests of statistical significance." This is, perhaps, their key contribution. It focuses on a need well and clearly demonstrated in their own empirical work.