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# Inventory Management Behavior of American and Japanese Firms*

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Firms view inventory/sales ratios as main decision variables. These ratios are also important in forecasting business activity. Understanding these ratios seems particularly relevant since many U.S. firms are beginning to implement the "just-in-time" production and inventory system that some see as the key to the typically lower inventory/sales ratios of Japanese firms. In this paper we present a partial adjustment model which can empirically characterize differences in the inventory behavior of U.S. and Japanese firms. Two key parameters estimated are the desired inventory/sales ratio and the speed of adjustment of the actual to the desired inventory/sales ratio. We find that Japanese firms in many industries have lower desired inventory/sales ratios and higher speeds of adjustment than U.S. firms. Simulation analyses are used to validate key assumptions of our model. *J. Japan. Int. Econ.*, September 1989, 3(3), pp. 270–291. Faculty of Business, University of Alberta, Edmonton, Alberta, Canada T6G 2R6. © 1989 Academic Press, Inc.

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## INTRODUCTION

Inventory/sales ratios play an important role in forecasters' models of business activity. It is frequently argued that executives view these ratios, or their reciprocals (inventory-turnover ratios), as important decision variables.[1] In fact, the business community's interest in these ratios is

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[1] Although the inventory turnover ratio is defined in many finance texts using the cost of goods sold, sales figures are often used in practice to calculate the ratio (see Schall and Haley, 1980, p. 398, and Brigham, 1979, p. 193).

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sufficiently great that graphs showing their movement over time often appear on the pages of leading business magazines.[2]

Understanding the nature of, and ultimately the reasons for, intercountry differences in inventory/sales ratios has become more important in the context of recent business interest in the United States and Canada in Japanese inventory management practices. In particular, many U.S. and Canadian firms are striving to implement the "just-in-time" production and inventory system (sometimes called the "Kanban" system)[3] that some see as the key to the typically lower inventory/sales ratios of Japanese firms. It is reported, for instance, that an important factor expected to contribute to the success of General Motors' Saturn project is that special features of Saturn's production facilities will make Japanese-style just-in-time inventory management more practical (see Fisher, 1984, p. 40).[4] Thus in this paper we focus on the movement over time of an annual measure of the inventory/sales ratio[5] for 1316 U.S. and Japanese firms in 15 industry groups. The period of analysis is 1964–1982 for the United States and 1961–1981 for Japan.

Unlike most published empirical studies of inventory behavior which are ultimately concerned with the impact of inventory investment on macro fluctuations, the object of this paper is to characterize differences

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[2] See, for example, the November 14, 1983 (p. 36) and March 3, 1986 (p. 31) issues of *Fortune*. These ratios are also an important macroeconomic indicator. (See, for example, Dornbusch and Fischer (1978, p. 200).)

[3] The just-in-time system practiced at Toyota is described in Monden (1981a,b). McClain and Thomas (1985) consider the costs associated with this system relative to other systems. Managerial and economic implications of the just-in-time system are discussed, for example, in Tsurumi (1984, p. 366), Abernathy *et al.* (1983, p. 75), Shimada (1985, p. 57), and Abegglen (1984, p. 176).

[4] See also *Time* (May 10, 1982, p. 57) and Brownstein (1984) for the impact of the just-in-time system on the inventory management of certain U.S. firms.

[5] Conceptually each of the raw materials, semifinished, and finished products could be subjected to inventory control to determine optimal ordering and holding quantities based on holding, shortage, and other costs. From the strategic management point of view, however, it is the aggregate inventory cost, rather than the individual material- or product-specific inventory cost, that is relevant to the firm. Such a firm-wide aggregate approach to inventory management is justified on several grounds. First, even though there may be thousands of different items to consider for inventory control, the demands for these items are highly correlated. Second, certain types of costs relevant to inventory control are also highly correlated. For example, holding costs for different items probably all depend on the cost of capital while the shortage costs all depend on the profitability of the final product (and hence of the firm). Last, it is common for a firm's manager to use the firm's total cost of inventory in formulating year-to-year budgets for the firm. In particular, it is standard practice to use the inventory turnover ratio, defined to be the ratio between the firm's total sales (or cost of goods sold) and inventory figures, as a management tool. (See, for example, Brigham (1979, p. 193) and Schall and Haley (1980, p. 398).)

in the inventory behavior of firms in two countries using a partial adjustment model as a descriptive tool. Partial adjustment models of inventory behavior have been found by other researchers to perform well in a forecasting sense. For example, the "Brookings Quarterly Economic Model of the United States" uses a partial adjustment model to explain investment in inventories (see Darling and Lovell, 1965). The partial adjustment model of inventory behavior used in this paper is similar in spirit to the Lintner (1956) model and the rational expectations model (Nakamura and Nakamura, 1985) of the dividend payout behavior of firms in that it allows us to estimate quantitatively certain decision parameters of firms' inventory behavior. Estimating inventory policy-related parameters empirically at the micro level using a partial adjustment model is new to the best of our knowledge.

In Section 1 we show how partial adjustment inventory behavior could result from cost minimization on the part of firms. The key parameters of the resulting model are the inventory/sales ratio that would be desired by the management of a firm in the absence of short-run transitionary costs and the speed of adjustment of the actual to the desired inventory/sales ratio. Our derivation of this model allows us to relate the determination of the values of its key parameters to the balance of underlying holding, shortage, and transitionary costs of inventory management. This in turn allows us to speculate on how we would expect the values of the desired inventory/sales ratio and speed of adjustment parameters to differ for U.S. versus Japanese firms in the same industry.

In Section 2 we first examine industry–country averages for the inventory/sales ratio. We next show industry–country-specific average estimation results for a suitably transformed version of our theoretical partial adjustment model that we have estimated using time series data for each of 1316 U.S. and Japanese firms. In Section 3 we report results from simulation analyses designed to reveal the extent of departures from certain key assumptions of our model.

## 1. A PARTIAL ADJUSTMENT INVENTORY MODEL

We will begin by considering a firm with a given stock of plant and equipment and given management practices. For this firm, the *holding costs* associated with any fixed level of inventory include the cost (interest or opportunity) of funds used to finance the inventory holdings, the rental or opportunity cost of warehouse space, inventory-related personnel costs, the actual or imputed rental cost of equipment used in inventory management, inventory-related utility and insurance costs, and so forth. The *shortage costs* associated with any fixed level of inventory include lost sales opportunities when the inventory level proves insufficient to meet customer demand and lost opportunities for speculative returns in factor and product markets. There are short-run *transitionary costs* associated with the movement to (as distinct from the maintenance of) some new inventory level. For instance, if inventories are to be expanded, time and other resources may have to be expended in search for additional storage facilities. Increases on short notice may also result in higher than normal financing, purchase, and shipping costs. On the other hand, if the inventory level is to be reduced it may be necessary to run advertisements to find a buyer or renter for unused warehouse space. If excess inventory is to be cleared quickly, it may even have to be sold off at below-normal sale prices.

We will assume that the total cost in period $t$ of inventory management for the given firm is

$$C_t = \mathrm{HSC}(I_t) + \mathrm{TC}(I_t - I_{t-1}),$$

where $\mathrm{HSC}(I_t)$ denotes the sum of the holding and shortage costs associated with inventory level $I_t$ and where $\mathrm{TC}(I_t - I_{t-1})$ denotes the short-run transitionary costs associated with moving from the inventory level $I_{t-1}$ in year $t - 1$ to the level $I_t$. Next we let $I_t^*$ denote the inventory level that would minimize $C_t$ for this firm if there were no short-run transitionary costs. That is, $I_t^*$ is the inventory level that would minimize $\mathrm{HSC}(I_t)$, the holding and shortage costs of the firm. We will sometimes refer to $I_t^*$ as the *desired inventory level*. Letting $\mathrm{HSC}(I_t^*)$ denote the holding and shortage costs associated with the desired inventory level and letting $\mathrm{HSC}(I_t - I_t^*)$ represent the additional holding and shortage costs due to the departure of $I_t$ from $I_t^*$, we can rewrite the total inventory-related costs of the firm as

$$C_t = \mathrm{HSC}(I_t^*) + \mathrm{HSC}(I_t - I_t^*) + \mathrm{TC}(I_t - I_{t-1}).$$

Note that from the definition of $I_t^*$ it follows that a one unit departure of $I_t^*$ from $I_t$ will increase holding costs more than it will decrease shortage costs if $I_t$ is greater than $I_t^*$, while it will increase shortage costs more than it will decrease holding costs if $I_t$ is less than $I_t^*$. Moreover it is assumed here that neither the holding cost nor the shortage cost associated with $I_t^*$ can be altered by the inventory-related short-run decision making of the firm. Suppose that the costs resulting from the deviation of $I_t$ from $I_t^*$ and from the transition from $I_{t-1}$ to $I_t$ can be approximated by $\mathrm{HSC}(I_t - I_t^*) = c_1(I_t - I_t^*)^2$ and by $\mathrm{TC}(I_t - I_{t-1}) = c_2(I_t - I_{t-1})^2$, respectively. Then the management of the firm can minimize $C_t$ by choosing the value of $I_t$ so as to minimize the discretionary component of $C_t$ given by

$$\mathrm{DC}_t = c_1(I_t - I_t^*)^2 + c_2(I_t - I_{t-1})^2, \tag{1}$$

subject to

$$I_t = I_{t-1} + x_t, \tag{2}$$

where $c_1$ is a cost factor associated with the deviation of $I_t$ from $I_t^*$, $c_2$ is a transitionary cost factor,[6] and $x_t$ is the adjustment in the inventory level from period $t - 1$ to $t$.

Substituting (2) into (1) and setting $(\delta \mathrm{DC}_t/\delta x_t) = 0$, we see that our cost minimizing firm will adjust the inventory level from period $t - 1$ to $t$ according to the partial adjustment relationship[7]

$$I_t - I_{t-1} = c(I_t^* - I_{t-1}), \tag{3}$$

where $c = c_1/(c_1 + c_2)$. Thus when $c_2 = 0$ (but $c_1 \neq 0$), then $c = 1$ and there will be full adjustment of $I_t$ to $I_t^*$; while when $c_1 = 0$ (but $c_2 \neq 0$), $c = 0$ and inventory in the current period will be maintained at the same level as in the previous period. Normally we would expect that $0 < c < 1$, with the *speed of adjustment parameter c* being closer to 1 (and hence the adjust-

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[6] The change in the inventory level from period $t - 1$ to $t$, $x_t = I_t - I_{t-1}$, is determined by aggregate effects of production and demand in period $t$. If production exceeds demand in $t$, then $I_t - I_{t-1} > 0$ and a transient holding cost for additional space requirements, additional capital needs, etc., will be incurred. On the other hand if demand exceeds production, then $I_t - I_{t-1} < 0$ and the probability of incurring a transient shortage cost due to, for example, the loss of customers' good will and price discounts that may have to be offered to get back lost customers increases. In the latter case the actual additional cost may be zero so long as $0 < I_t\ (<I_{t-1})$, i.e., production in $t$ and $I_{t-1}$ combined exceeds demand, but if $I_t$ becomes negative (i.e., demand exceeds availability), then the additional shortage cost is incurred. As in many studies (Holt *et al.*, 1960; and also Hax and Candea, 1984) on aggregate production/inventory planning, we assume here that this nonlinear cost of shortage and the cost associated with $I_t - I_{t-1} > 0$ are both convex in $(I_t - I_{t-1})$ and can be approximated by a symmetric quadratic function of $(I_t - I_{t-1})$. Holt *et al.* (1960) found that quadratic functions could approximate the actual costs involved in planning production and employment quite well. Our empirical results also suggest that this provides a reasonable approximation since estimates for the parameter $c = c_1/(c_1 + c_2)$ lie between 0 and 1 for all industries for the United States and Japan, as our model implies they should.

[7] The partial adjustment model was first used as an econometric specification by Nerlove (1956). Despite extensive use of this model in inventory studies since then, however, the fact that a simple cost minimization problem of the sort given by (1) and (2) leads to the partial adjustment specification given in (3) seems to have gone unnoticed in the published literature. This derivation of a partial adjustment model holds even if the firm's optimization problem (1) and (2) is replaced by a more general multiperiod stochastic minimization problem, so long as the certainty equivalent principle holds. As a referee pointed out, (3) can be derived directly by minimizing (1) with respect to $I_t$. The definitional relation (2) is added for purely expositional purposes.

ment of $I_t$ to $I_t^*$ being more complete) the larger $c_1$, the holding and shortage cost factor, is compared with $c_2$, the transitionary cost factor.

For a firm with a given stock of plant and equipment, given management practices, and in a given business environment, the holding costs associated with $I_t^*$ will depend primarily on the value of $I_t^*$, and the shortage costs associated with any given value of $I_t^*$ will depend primarily on the level of sales. Thus there will be some sort of a relationship between the (unobservable) desired inventory level and sales, that might be approximated by[8]

$$I_t^* = b_1 + \alpha \text{ SALES}_t, \tag{4}$$

where $\text{SALES}_t$ denotes the real value of sales in period $t$. If $b_1$ in (4) is close to zero, then $\alpha$ can be interpreted as the *desired inventory/sales ratio* in the absence of short-run transitionary costs. Combining (3) and (4), we get the partial adjustment model

$$I_t - I_{t-1} = cb_1 + c\alpha \text{ SALES}_t - cI_{t-1}. \tag{5}$$

Of course, there is no reason why the errors resulting from approximating $I_t^*$ as in (4) must have a mean of zero over any given time period. Also the magnitudes of these errors could clearly depend on firm size, resulting in a heteroscedastic error term. Thus the inventory adjustment equation actually estimated for each firm is

$$\frac{I_t - I_{t-1}}{\text{SALES}_{t-1}} = a_0 + a_1 \left(\frac{1}{\text{SALES}_{t-1}}\right) + a_2 \left(\frac{\text{SALES}_t}{\text{SALES}_{t-1}}\right) + a_3 \left(\frac{I_{t-1}}{\text{SALES}_{t-1}}\right) + \mu_t, \tag{6}$$

where $\mu_t$ is a random error term, $a_0$ is a parameter that will have a value of zero if the errors of approximation in (4) have a mean of zero, $a_1 = cb_1$, $a_2 = c\alpha$ and $a_3 = -c$. Obviously if the values of $\alpha$ and $c$ (and hence of $a_1 - a_3$) are not really constants for a firm over the observation period, then

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[8] The desired inventory level in general depends on expected future sales. The specification given by (4) implies that sales follow a random walk process. Since sales series are highly autocorrelated, (4) may be a reasonable specification and is often assumed. Our empirical results also indicate that there is a stable desired inventory/sales ratio $\alpha$ as specified in (4). Nevertheless, it is of interest to see if estimates for desired inventory/sales ratios would be different depending on the stochastic process assumed for sales, and new research is being planned to investigate this matter within a rational expectations framework. (See Nakamura and Nakamura (1985) for a similar problem in firms' dividend behavior.)

this variability in the true parameter values will be reflected in the residuals for the estimated inventory adjustment equation for the firm. (In this case, the estimated parameter values for Eq. (6) might more properly be viewed as some sort of firm-specific averages over the relevant time period.) Moreover we might expect that the values of $\alpha$ and $c$ for firms in the same industry, or even for firms in related industries, would be affected in a similar manner by major exogenous changes in cost conditions such as those resulting from the "oil shock" of 1974. This, in turn, would presumably lead to substantial departures in the years in which these cost changes occurred between the predicted and actual industry–country-average inventory/sales ratios for firms in the affected industry–country groups. In Section 3 we report the results of simulation analyses designed to check for departures over time of this sort.

The value for $\alpha$ for each firm is determined by the tradeoffs between holding and shortage costs, and the value of $c$ for each firm is determined by tradeoffs between the sum of holding and shortage costs versus transitionary costs. Thus systematic differences in the holding, shortage, and transitionary cost schedules faced by U.S. versus Japanese firms should lead to systematic intercountry differences in the estimated values of $\alpha$ and $c$ and in the observed inventory/sales ratios.

Equation (6) controls for the firm's inventory level in the previous time period and is suitable for measuring policy parameters using temporal changes in the inventory level. Partial adjustment models like (6) also produce reasonable parameter estimates even if the sample period contains exogenous shocks provided that the firm management does not alter its inventory policy. This is because these models rely primarily on firms' period-to-period (rather than static) responses to changing business conditions. On the other hand, if changing business conditions in, say financial, labor, or product markets also force firms to change their inventory policies, then Eq. (6) by itself cannot be used to estimate key policy parameters. Additional variables or equations to control for the changing business conditions, perhaps within a simultaneous equations framework, would have to be added to the specification of firms' inventory behavior. One difficulty in applying a simultaneous system to firms' inventory behavior at the micro level is that there are very few variables which can be used as good instruments. This is an important topic for future research.

One important intercountry cost difference is that land for commercial development is more scarce and hence land values are much higher in Japan than in the United States, resulting in higher rental or opportunity costs for warehouse space in Japan. At the same time, domestic delivery costs and delivery-related waiting times are innately lower for Japanese than for U.S. firms due to the geographic size disparity between the two countries. These differing natural comparative advantages of U.S. and

Japanese firms appear to have been enhanced over the years by endogenous factors. For instance, Japanese firms seem to have enhanced their natural shortage cost advantage through superior quality control (leading to fewer rejects and hence to lower shortage costs at any given inventory level) and through the location and design of plants which allow delivery costs to be kept to a minimum. Even Japanese labor relations practices seem to have evolved since World War II in a manner that enhances the natural shortage cost advantage of Japanese firms. Large firms in Japan offer lifetime employment to most of their regular employees, and most regular employees in Japanese commercial enterprises receive a significant fraction of their annual pay in the form of profit-sharing bonuses. Weitzman (1984, p. 74) observes: "Workers respond with corporate loyalty. . . ." Part of "corporate loyalty" is a pervasive willingness on the part of employees in Japan to work overtime, or delay or forego vacations, in order to meet heavier than normal production demands arising on short notice. This willingness makes it possible for firms to fill orders quickly without maintaining large inventories. Japanese firms facing relatively high holding costs but lower shortage cost schedules would be expected to have systematically lower values of $\alpha$ than U.S. firms in the same industry.

Suppose for simplicity that in (6) we have $a_0 = 0$ and $b_1 = 0$ where $a_1 = cb_1$. Then ignoring the random error in (6)

$$\frac{I_t}{\text{SALES}_t} = c\alpha + (1 - c)\left(\frac{I_{t-1}}{\text{SALES}_{t-1}}\right)\left(\frac{\text{SALES}_{t-1}}{\text{SALES}_t}\right). \tag{7}$$

Thus we see that a lower value of $\alpha$ will unambiguously contribute to a lower observed inventory/sales ratio in relationship to other firms. Whether the observed inventory/sales ratio for a particular firm lies above or below its own value of $\alpha$ will depend on the inventory/sales ratio for the firm in the previous year and the relationship between the firm's sales in the current and previous years, among other things. If sales have been growing (that is, $(\text{SALES}_{t-1}/\text{SALES}_t) < 1$) we would usually expect the annual inventory/sales ratio for a firm to lie below $\alpha$, with the discrepancy being greater the closer $c$ is to zero. Since sales were generally growing for most of the U.S. and Japanese firms for which we have data over the periods of observation for this study, we would expect our estimated values of $\alpha$ to lie above the observed inventory/sales ratios. We have no expectations concerning intercountry differences in $c$. Nevertheless, since we do expect the values of $\alpha$ to be lower for Japanese than for U.S. firms in the same industry, based on (7) we also expect the observed inventory/sales ratios to be lower for Japanese firms than for their U.S. counterparts.

## 2. EMPIRICAL RESULTS

Using annual observations over the period of 1963–1982 for U.S. firms and over the period of 1960–1981 for Japanese firms, we calculated average inventory/sales ratios for 15 industry groups for each country. These inventory/sales ratios are displayed in Table I together with estimated standard deviations. (See footnote *a* of Table I for the details of these computations.) In Table I we also show the number of firms in each industry–country group. The average inventory/sales ratios shown in Table I range from 0.05 to 0.27 for the United States and from 0.02 to 0.26 for Japan. The rankings of these average ratios by industry are also very similar for both countries. (The value of Spearman's rank-correlation coefficient is 0.89.) That is, the industries where the inventory/sales ratio

TABLE I
INDUSTRY–COUNTRY-SPECIFIC AVERAGE INVENTORY/SALES RATIOS

| | United States | | | Japan | | |
|---|---|---|---|---|---|---|
| Industry | Inventory/ Sales | SD | Number of firms | Inventory/ Sales | SD | Number of firms |
| Ratio lower for Japanese than for U.S. firms | | | | | | |
| Food | 0.19 | 0.05[a] | 51 | 0.14 | 0.04[a] | 47 |
| Textiles | 0.23 | 0.04 | 37 | 0.21 | 0.05 | 58 |
| Lumber | 0.14 | 0.03 | 59 | 0.12 | 0.04 | 32 |
| Chemicals | 0.18 | 0.03 | 79 | 0.14 | 0.03 | 75 |
| Rubber | 0.18 | 0.03 | 42 | 0.14 | 0.05 | 10 |
| Industrial machinery | 0.27 | 0.04 | 56 | 0.26 | 0.09 | 61 |
| Transportation machinery | 0.21 | 0.05 | 49 | 0.17 | 0.06 | 46 |
| Transportation | 0.05 | 0.05 | 33 | 0.02 | 0.02 | 22 |
| Utilities | 0.08 | 0.02 | 4 | 0.06 | 0.02 | 14 |
| Wholesale/retail | 0.15 | 0.03 | 76 | 0.07 | 0.03 | 39 |
| Ratio higher for Japanese than for U.S. firms | | | | | | |
| Petro refining | 0.12 | 0.06 | 29 | 0.14 | 0.05 | 8 |
| Cement | 0.15 | 0.04 | 14 | 0.18 | 0.05 | 30 |
| Steel/metal | 0.20 | 0.04 | 81 | 0.22 | 0.07 | 68 |
| Electrical machinery/ computing | 0.22 | 0.05 | 94 | 0.23 | 0.06 | 50 |
| Precision/others | 0.23 | 0.04 | 33 | 0.26 | 0.09 | 19 |

[a] The firm-specific average inventory/sales ratio and its standard deviation are first calculated for each firm using $T$ years of data, where $T = 20$ for the United States (1963–1982) and $T = 22$ (1960–1981) for Japan. These firm-specific averages and standard deviations are then averaged over the firms in each industry–country group to obtain the inventory/sales ratios shown in this table and the accompanying standard deviations.

tends to be high (low) for the United States are essentially the same industries where this ratio tends to be high (low) for Japan. Despite the U.S.–Japanese similarities in both the ranges and the rankings of the ratios shown in Table I, however, the Japanese averages are lower than the U.S. averages (as expected) for 10 of the 15 industry groupings.

Our inventory adjustment model, Eq. (6), was estimated using ordinary least-squares regression with annual data over the appropriate time period for each of the 1316 U.S. and Japanese firms included in this study. In Table II we show the average coefficient estimates and $R^2$ values from these regressions for the firms in each country in each of our 15 industry groups. (Average $t$ values are shown in parentheses.) For each industry–country group, the negative of the average coefficient estimate shown for the lagged inventory/sales ratio, and the negative of the average coefficient estimate for the growth variable ($\text{SALES}_t/\text{SALES}_{t-1}$) divided by the coefficient of the lagged inventory/sales ratio, can be interpreted, respectively, as the estimated values of $c$ and $\alpha$ for the average firm in the designated industry and country. The "average firm" estimates for $\alpha$ are shown in the next-to-the-last column of Table II, while the corresponding estimated values for $-c$ appear in column 2.

The average values for the coefficient estimates of $(a_1/10^4) = (cb_1/10^4)$ and $a_0$ (and for the corresponding $t$ statistics) are shown in columns 3 and 4 of Table II. We see that in general both $a_0$ and $b_1$ (from expression (4)) can be considered to be essentially equal to zero for the average firm in each industry–country group. Thus our average firm estimates of $\alpha$ can appropriately be interpreted as estimates of the desired inventory/sales ratio in the absence of short-run transitionary costs.

Our partial adjustment model implies that higher (lower) values of $\alpha$ should be associated with higher (lower) values of the actual inventory/sales ratio. Comparing the estimated values of $\alpha$ shown in Table II with the industry–country-specific average inventory/sales ratios shown in Table I, we find that this is true. (The value of the Spearman rank-correlation coefficient for the average inventory/sales ratios shown in Table I compared with the estimated values of $\alpha$ shown in Table II is 0.59 for the United States, 0.90 for the United States omitting the utilities industry for which we have observations for only four firms, and 0.84 for Japan.) Looking at the actual values of the estimates for $\alpha$ shown in Table II, we see they lie in a range of 0.088 to 0.894 for the United States, from 0.088 to 0.461 for the United States omitting the utilities industry, and from 0.050 to 0.425 for Japan. Despite these similarities, however, the estimated values for $\alpha$ are generally lower for Japan as expected. (The estimated values for $\alpha$ for Japan lie below the median value for the United States for all but 3 of our 15 industry groups.) From the summary information presented in Table III it can also be seen that the estimated $\alpha$ values shown in

TABLE II

INVENTORY BEHAVIOR OF U.S. AND JAPANESE FIRMS: COEFFICIENT ESTIMATES[a]

| Industry | Coefficient of $\text{Sales}_t/\text{Sales}_{t-1}$ ($c\alpha$) | Coefficient of $\text{Inventory}_{t-1}/\text{Sales}_{t-1}$ ($-c$) | Coefficient of $10^4/\text{Sales}_{t-1}$ | Constant | $R^2$ | $\alpha^b$ | Z statistics for $\hat{\alpha}_{\text{U.S.}} - \hat{\alpha}_{\text{J}}$ |
|---|---|---|---|---|---|---|---|
| Industry–country-specific average inventory/sales ratio lower for Japanese than for U.S. firms | | | | | | | |
| Food | | | | | | | 5.32 |
| United States | 0.149**[c] | −0.436* | 0.001 | −0.103 | 0.552 | 0.342 | |
| | (4.55) | (1.89) | (0.003) | (1.72) | | | |
| Japan | 0.117* | −0.666** | 0.033 | −0.039 | 0.494 | 0.176 | |
| | (2.08) | (3.00) | (0.79) | (0.65) | | | |
| Textiles | | | | | | | 3.18 |
| United States | 0.209** | −0.453* | 0.000 | −0.104* | 0.631 | 0.461 | |
| | (5.23) | (1.93) | (0.00) | (1.81) | | | |
| Japan | 0.175** | −0.551** | 0.022 | −0.078 | 0.645 | 0.318 | |
| | (3.36) | (3.01) | (0.38) | (1.46) | | | |
| Lumber | | | | | | | 8.88 |
| United States | 0.113** | −0.466* | 0.000 | −0.056 | 0.565 | 0.242 | |
| | (3.44) | (2.05) | (0.63) | (1.39) | | | |
| Japan | 0.087** | −0.595** | 0.045 | −0.037 | 0.582 | 0.050 | |
| | (3.14) | (2.95) | (0.85) | (1.19) | | | |
| Chemicals | | | | | | | 3.91 |
| United States | 0.164** | −0.481* | 0.000 | −0.085* | 0.577 | 0.342 | |
| | (3.97) | (2.01) | (0.11) | (0.07) | | | |
| Japan | 0.132** | −0.540** | 0.014 | −0.061 | 0.658 | 0.244 | |
| | (4.95) | (2.41) | (0.36) | (1.60) | | | |
| Rubber | | | | | | | 3.26 |
| United States | 0.160** | −0.511* | 0.001 | −0.083* | 0.635 | 0.313 | |
| | (4.14) | (2.19) | (0.33) | (1.62) | | | |
| Japan | 0.119** | −0.597** | 0.009 | −0.046 | 0.653 | 0.199 | |
| | (4.87) | (2.87) | (0.12) | (1.23) | | | |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| **Industrial machinery** | | | | | | | 0.37 |
| **United States** | 0.216** | −0.576* | 0.000 | −0.070* | 0.609 | 0.375 | |
| | (4.04) | (2.36) | (0.31) | (1.18) | | | |
| **Japan** | 0.173** | −0.477* | 0.008 | −0.065 | 0.560 | 0.363 | |
| | (3.90) | (2.52) | (0.736) | (1.14) | | | |
| **Transportation machinery** | | | | | | | 0.64 |
| **United States** | 0.183** | −0.507* | 0.001 | −0.081 | 0.564 | 0.361 | |
| | (4.03) | (2.17) | (0.08) | (1.28) | | | |
| **Japan** | 0.143** | −0.424** | −0.014 | −0.071 | 0.538 | 0.337 | |
| | (4.74) | (2.01) | (0.52) | (1.55) | | | |
| **Transportation** | | | | | | | −0.47 |
| **United States** | 0.043** | −0.488* | −0.000 | −0.022 | 0.487 | 0.088 | |
| | (2.99) | (2.12) | (0.04) | (0.68) | | | |
| **Japan** | 0.035** | −0.363 | −0.001 | −0.028 | 0.504 | 0.096 | |
| | (3.26) | (1.50) | (0.31) | (1.73) | | | |
| **Utilities** | | | | | | | 0.99 |
| **United States** | 0.126** | −0.141 | −0.001 | −0.112* | 0.392 | 0.894 | |
| | (2.84) | (0.61) | (0.24) | (1.93) | | | |
| **Japan** | 0.070** | −0.433* | −0.007 | −0.040* | 0.598 | 0.162 | |
| | (5.62) | (1.81) | (0.10) | (2.29) | | | |
| **Retail/wholesale** | | | | | | | 1.97 |
| **United States** | 0.128** | −0.550* | 0.000 | −0.060 | 0.640 | 0.233 | |
| | (4.69) | (2.30) | (0.32) | (1.47) | | | |
| **Japan** | 0.064** | −0.338 | −0.003 | −0.041* | 0.631 | 0.189 | |
| | (5.27) | (1.67) | (0.08) | (2.20) | | | |
| **Industry–country-specific average inventory/sales ratio higher for Japanese than for U.S. firms** | | | | | | | |
| **Petro refining** | | | | | | | −2.87 |
| **United States** | 0.084** | −0.458* | 0.003 | −0.044 | 0.653 | 0.173 | |
| | (5.88) | (2.18) | (1.43) | (1.64) | | | |
| **Japan** | 0.200** | −0.596* | −0.095 | −0.084 | 0.759 | 0.336 | |
| | (7.65) | (2.31) | (1.27) | (1.60) | | | |

TABLE II—*Continued*

| Industry | Coefficient of $\text{Sales}_t/\text{Sales}_{t-1}$ $(c\alpha)$ | Coefficient of $\text{Inventory}_{t-1}/\text{Sales}_{t-1}$ $(-c)$ | Coefficient of $10^4/\text{Sales}_{t-1}$ | Constant | $R^2$ | $\alpha^b$ | $Z$ statistics for $\hat{\alpha}_{\text{U.S.}} - \hat{\alpha}_J$ |
|---|---|---|---|---|---|---|---|
| Cement | | | | | | | −3.16 |
| United States | 0.099* | −0.559* | 0.001 | −0.046 | 0.494 | 0.177 | |
| | (2.51) | (2.07) | (1.18) | (1.30) | | | |
| Japan | 0.149** | −0.493* | −0.008 | −0.053 | 0.647 | 0.302 | |
| | (4.51) | (2.36) | (0.31) | (1.36) | | | |
| Steel/metal | | | | | | | −5.37 |
| United States | 0.132** | −0.569* | 0.002 | −0.060 | 0.551 | 0.232 | |
| | (3.40) | (2.32) | (1.22) | (1.20) | | | |
| Japan | 0.182** | −0.509* | −0.007 | −0.078 | 0.638 | 0.358 | |
| | (5.10) | (2.44) | (0.16) | (1.45) | | | |
| Electrical machinery/computing | | | | | | | −2.27 |
| United States | 0.180** | −0.520* | 0.001 | −0.062 | 0.593 | 0.346 | |
| | (3.56) | (2.27) | (0.24) | (0.97) | | | |
| Japan | 0.192** | −0.452* | −0.012 | −0.096 | 0.631 | 0.425 | |
| | (7.32) | (2.17) | (0.26) | (1.76) | | | |
| Precision/others | | | | | | | 1.19 |
| United States | 0.227** | −0.614* | 0.000 | −0.083 | 0.661 | 0.370 | |
| | (5.50) | (2.47) | (0.52) | (1.10) | | | |
| Japan | 0.190** | −0.598* | 0.004 | −0.077 | 0.651 | 0.318 | |
| | (5.25) | (2.53) | (0.58) | (1.11) | | | |

[a] The numbers reported in this table are industry means of regression coefficients and associated $t$ statistics (given in parentheses) from regression runs for each firm in each industry using 19 (for the United States) or 21 (for Japan) annual observations. The number of firms in each industry–country group is given in Table I.

[b] These estimates of $\alpha$ are derived by dividing the estimate for $c\alpha$ given under the column heading ($\text{Sales}/\text{Sales}_{-1}$) by the negative of the estimated $-c$ given under the column heading ($\text{Inventory}_{-1}/\text{Sales}_{-1}$).

[c] One and two asterisks indicate, respectively, two-tailed significance levels of at least 95 and 99% for the "average" firm.

TABLE III

U.S.—JAPANESE SUMMARY COMPARISONS

| Industry | Derived estimate of $\alpha$ for average Japanese firm lower than for average U.S. firm | Estimated value of $c$ for average Japanese firm higher than for average U.S. firm |
|---|---|---|
| | Industry–country-specific average inventory/sales ratio lower for Japanese than for U.S. firms | |
| Food | Yes | Yes |
| Textiles | Yes | Yes |
| Lumber | Yes | Yes |
| Chemicals | Yes | Yes |
| Rubber | Yes | Yes |
| Industrial machinery | Yes | No |
| Transportation machinery | Yes | No |
| Transportation | No | No |
| Utilities | Yes | Yes |
| Wholesale/retail | Yes | No |
| | Industry–country-specific average inventory/sales ratio higher for Japanese than for U.S. firms | |
| Petro refining | No | Yes |
| Cement | No | No |
| Steel/metal | No | No |
| Electrical machinery/ computing | No | No |
| Precision/others | Yes | No |

Table II are smaller for Japan than for the United States for 9 out of the 10 industry groups for which the observed average inventory/sales ratio is smaller for Japan, while the reverse is true for 4 out of the 5 industry groups for which the observed average inventory/sales ratio is larger for Japan.

It is also possible to test formally a null hypothesis $H_0$, $\alpha_{\text{U.S.}} = \alpha_J$, against $H_1$, $\alpha_{\text{U.S.}} > \alpha_J$, under certain circumstances, where $\alpha_{\text{U.S.}}$ and $\alpha_J$, respectively, denote population values of $\alpha$ for U.S. and Japanese firms. Under $H_0$, a test statistic is $Z = (\hat{\alpha}_{\text{U.S.}} - \hat{\alpha}_J)/\text{SD}(\hat{\alpha}_{\text{U.S.}} - \hat{\alpha}_J)$, where $\hat{\alpha}_{\text{U.S.}}$ and $\hat{\alpha}_J$ are, respectively, estimates of $\alpha$ for U.S. and Japanese firms and $\text{SD}(\hat{\alpha}_{\text{U.S.}} - \hat{\alpha}_J)$ is an estimated standard deviation of $(\hat{\alpha}_{\text{U.S.}} - \hat{\alpha}_J)$ for which we use Pearson's formula for ratios of random variables.[9] Values of $Z$-

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[9] Pearson's formula for ratio variables (Pearson, 1897; see also Mood *et al.* 1974, p. 181) is: for the two correlated random variables $X$ and $Y$, $V(X/Y) \simeq (EX/EY)^2\{(V(X)/(EX)^2 + (V(Y)/E(Y))^2 - 2\text{Cov}(X, Y)/EXEY\}$. Since $\hat{\alpha}$ is the ratio between the two estimated regression coefficients $c\hat{\alpha}$ and $\hat{c}$ its variance $V(\hat{\alpha})$ can be approximated by Pearson's

statistics are given in the last column of Table II. Assuming that $Z$ obeys approximately the standard normal distribution, $H_0$ is rejected for 6 out of 10 industries in the first panel in Table II setting the size of the test at 0.025. This is consistent with our earlier observations.

Because of generally growing sales for both U.S. and Japanese firms during the sample period, our prior expectation is that the estimated values of $\alpha$ will lie above the observed inventory/sales ratios. We find this to almost always be the case for both the United States and Japan, comparing observed inventory/sales ratios in Table I and estimated desired ratios in Table II.

Looking again now at Table II, we find that the estimated average firm values of $c$ range from 0.141 to 0.614 for the United States, from 0.576 to 0.614 for the United States excluding the utilities category, and from 0.338 to 0.666 for Japan. Thus the ranges for the two countries are similar. However, we cannot detect any common pattern over industries for the two countries. (The value of Spearman's rank-correlation coefficient for the United States compared with the Japanese values for $c$ is only 0.15). We do find from Table III though that for the 10 industries where the average inventory/sales ratio is lower for Japan than for the United States, the Japanese estimates for $c$ in Table II are often larger than the U.S. estimates (6 out of 10 times), while the reverse tends to be true (4 out of 5 times) for the remaining industries.

## 3. SIMULATION RESULTS

In estimating Eq. (6) for each firm using time series data we are implicitly assuming that the actual values of $\alpha$ and $c$ for each firm have remained approximately constant over the relevant observation period. Suppose that this is not so. Then the correspondence over time between the predicted and actual average inventory/sales ratios may be poor. (The predicted average ratios are obtained by using the firm-specific estimated versions of Eq. (6) to forecast the inventory levels in each year, and hence the annual inventory sales ratios, for the individual firms in each industry group, and industry–country annual averages are then computed.) This will be particularly the case if the movements in the firm-specific values of

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formula, where appropriate estimates of relevant moments are substituted in. Once the $V(\hat{\alpha}_{U.S.})$ and $V(\hat{\alpha}_J)$ are calculated, then $\mathrm{SD}(\hat{\alpha}_{U.S.} - \hat{\alpha}_J) = \{V(\hat{\alpha}_{U.S.} - \hat{\alpha}_J)\}^{1/2} = \{V(\hat{\alpha}_{U.S.}) + V(\hat{\alpha}_J)\}^{1/2}$ is derived, where $\mathrm{Cov}(\hat{\alpha}_{U.S.}, \hat{\alpha}_J) = 0$ since $\hat{\alpha}_{U.S.}$ and $\hat{\alpha}_J$ are estimated from two independent (Japanese and U.S.) data sets. (Another way to calculate the (large sample) standard error of $\hat{\alpha}$ which is a function of estimated parameters is found in Rao (1973, p. 323).) More research is certainly needed to develop a model in which $\alpha$ appears as a separate parameter to be estimated.

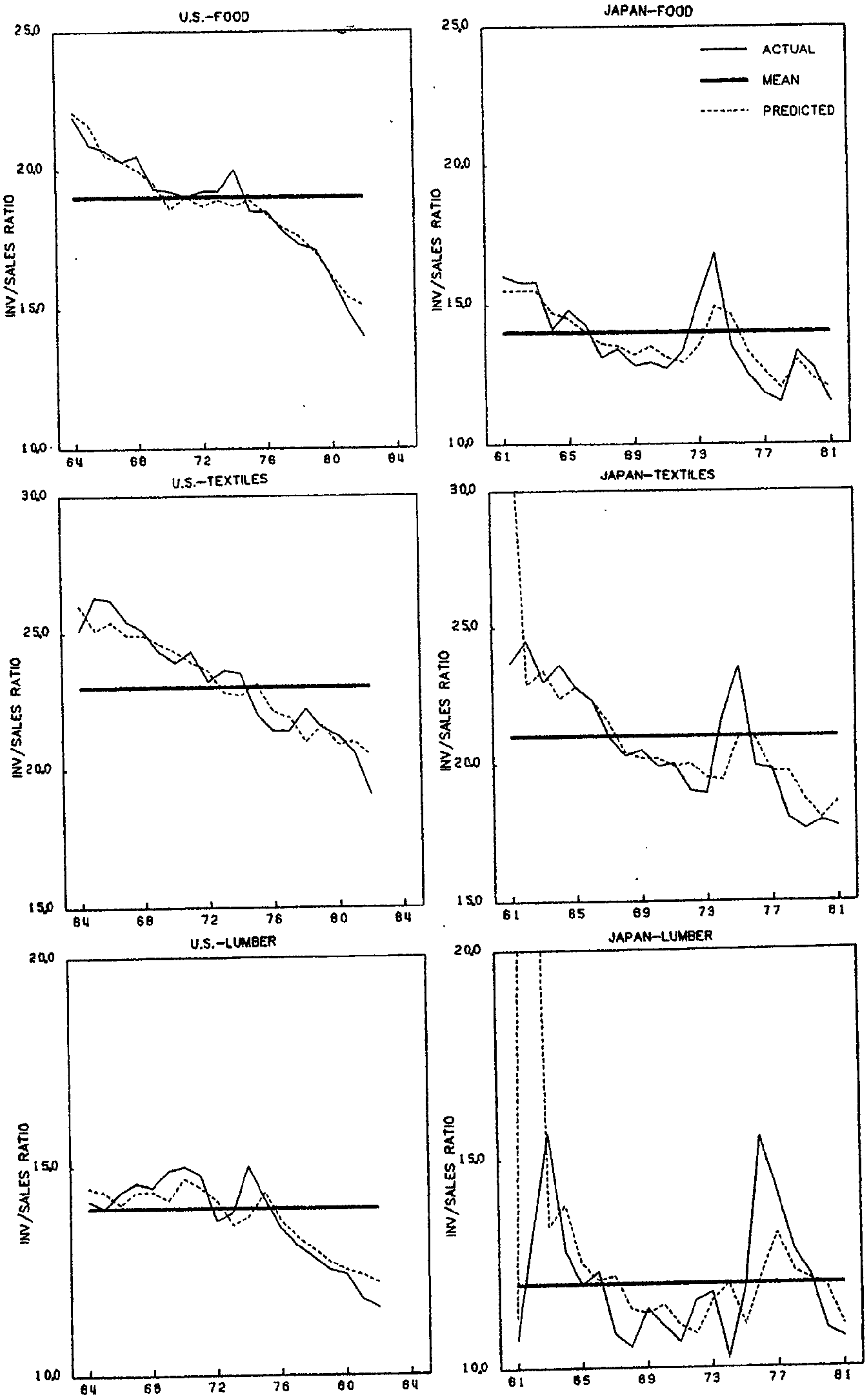

FIG. 1. Graphs for the U.S. and Japanese industries: Food, textiles, lumber, chemicals, and rubber. Each graph shows the inventory/sales ratio over the time period 1964–1984 for the United States and 1961–1981 for Japan. The solid lines correspond to the actual numbers; the bold lines correspond to the mean values; and the dashed lines correspond to the values predicted by the model.

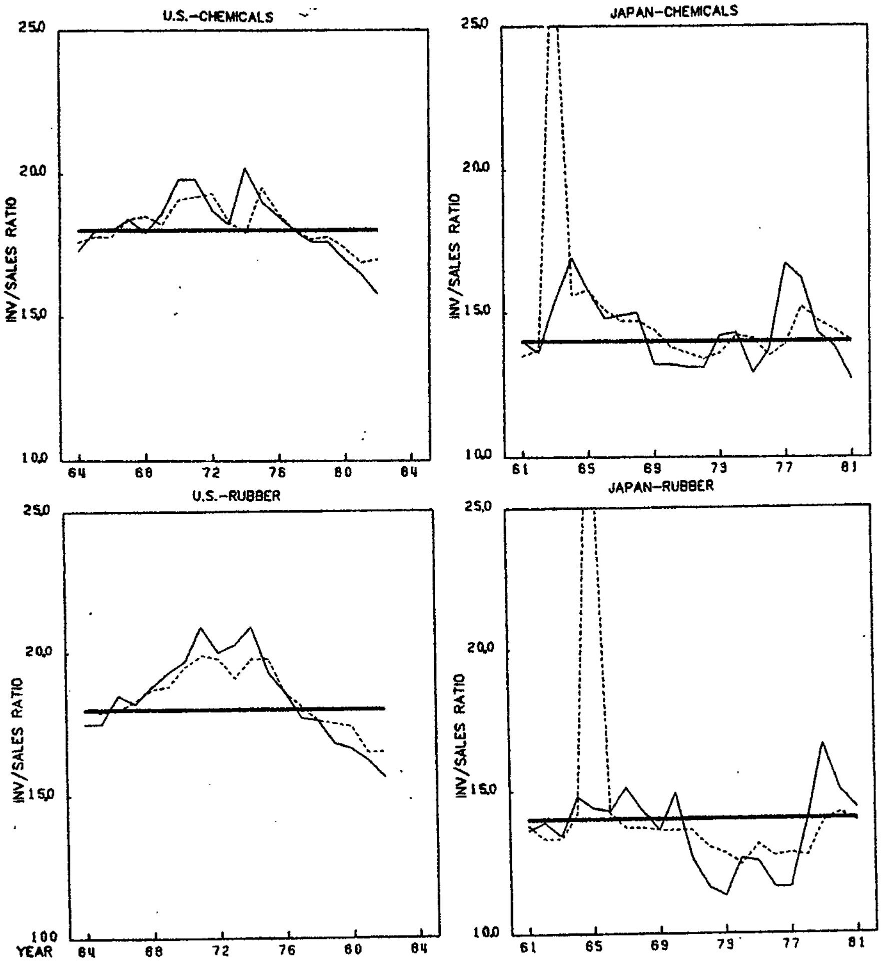

FIG. 1—*Continued*

$\alpha$ and $c$ tend to be correlated for the firms in particular industry groups. From Figs. 1–3 we see that in 1974, the year of the first oil shock, the actual average inventory/sales ratio does lie above the average predicted ratio for most of the 15 industries for the United States. There are also other substantial departures between the average actual and predicted inventory/sales ratios for both U.S. and Japanese firms. In general, however, we view the correspondence between the movements of the average actual and predicted ratios graphed in Figs. 1 through 3 as remarkably good, and feel that this correspondence provides limited support for our treatment of $\alpha$ and $c$ as fixed firm-specific parameters over the periods of 1964–1982 for U.S. firms and of 1962–1981 for Japanese firms. Certainly the prediction results based on Eq. (6) are far better than the results based

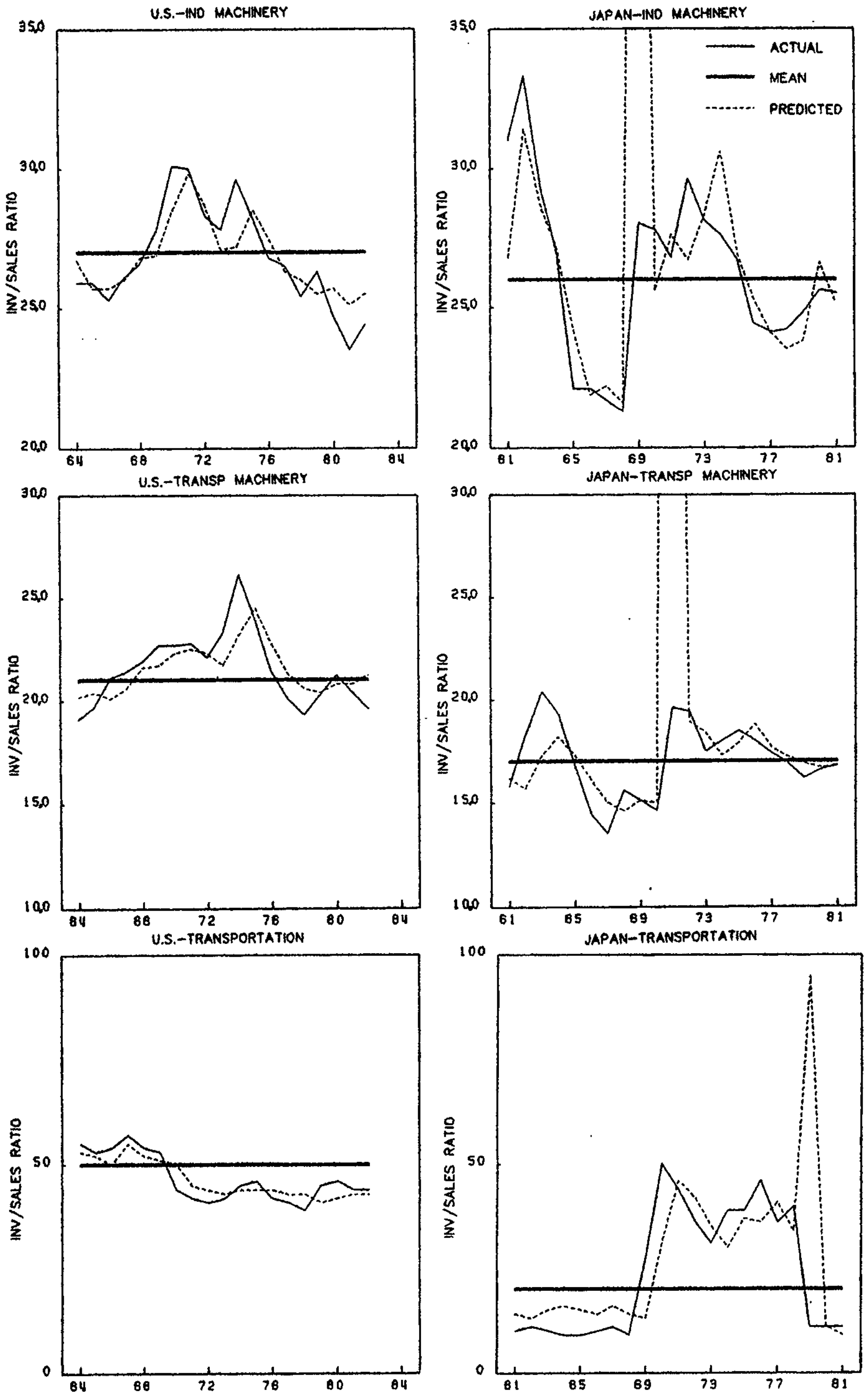

FIG. 2. Graphs for the U.S. and Japanese industries: Industrial machinery, transportation machinery, transportation, utilities, and wholesale/retail. The meaning of each graph is exactly the same as for Fig. 1.

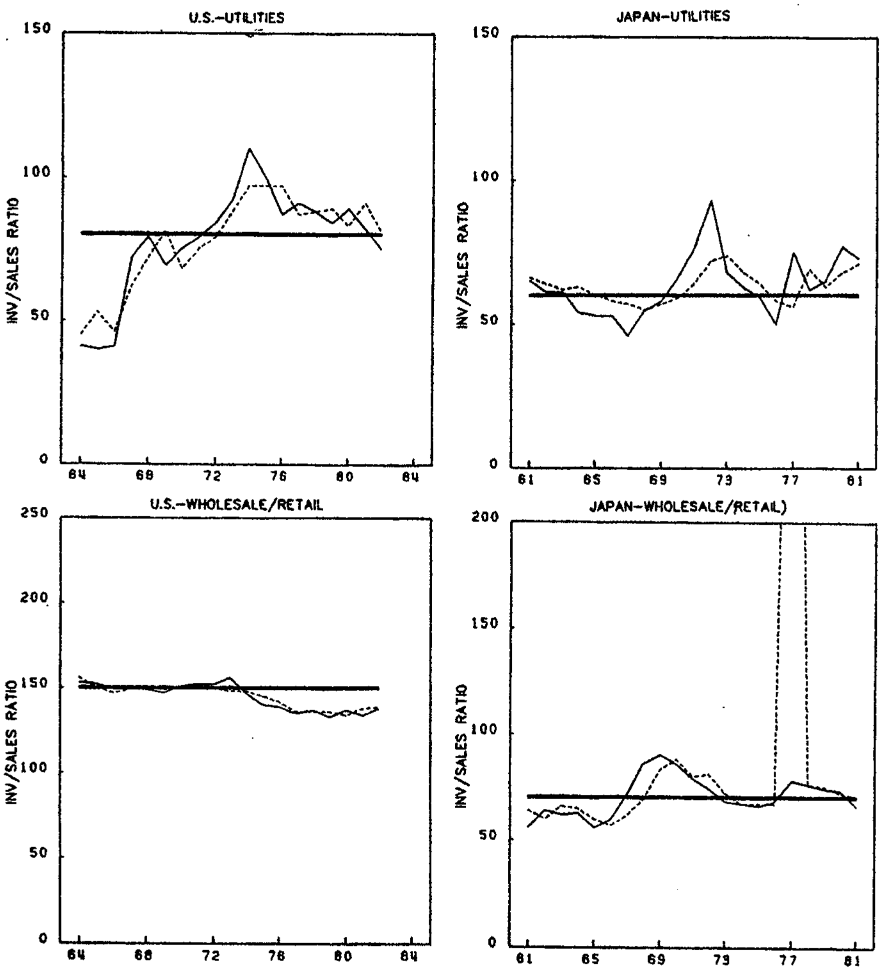

FIG. 2—*Continued*

on the assumption of a constant inventory/sales ratio for each firm (the horizontal line across each graph).

## APPENDIX: DATA AND VARIABLE DEFINITIONS

We used the Compustat tape (1963–1982) to create a data base for U.S. firms. We eliminated those firms which do not have observations for all relevant variables over the period of 1963–1982. All variables (measured in millions of dollars) were deflated by the consumer price index to 1960 dollars. A similar data base was created for Japanese firms (1960–1981) from the Japan Development Bank data base which includes information

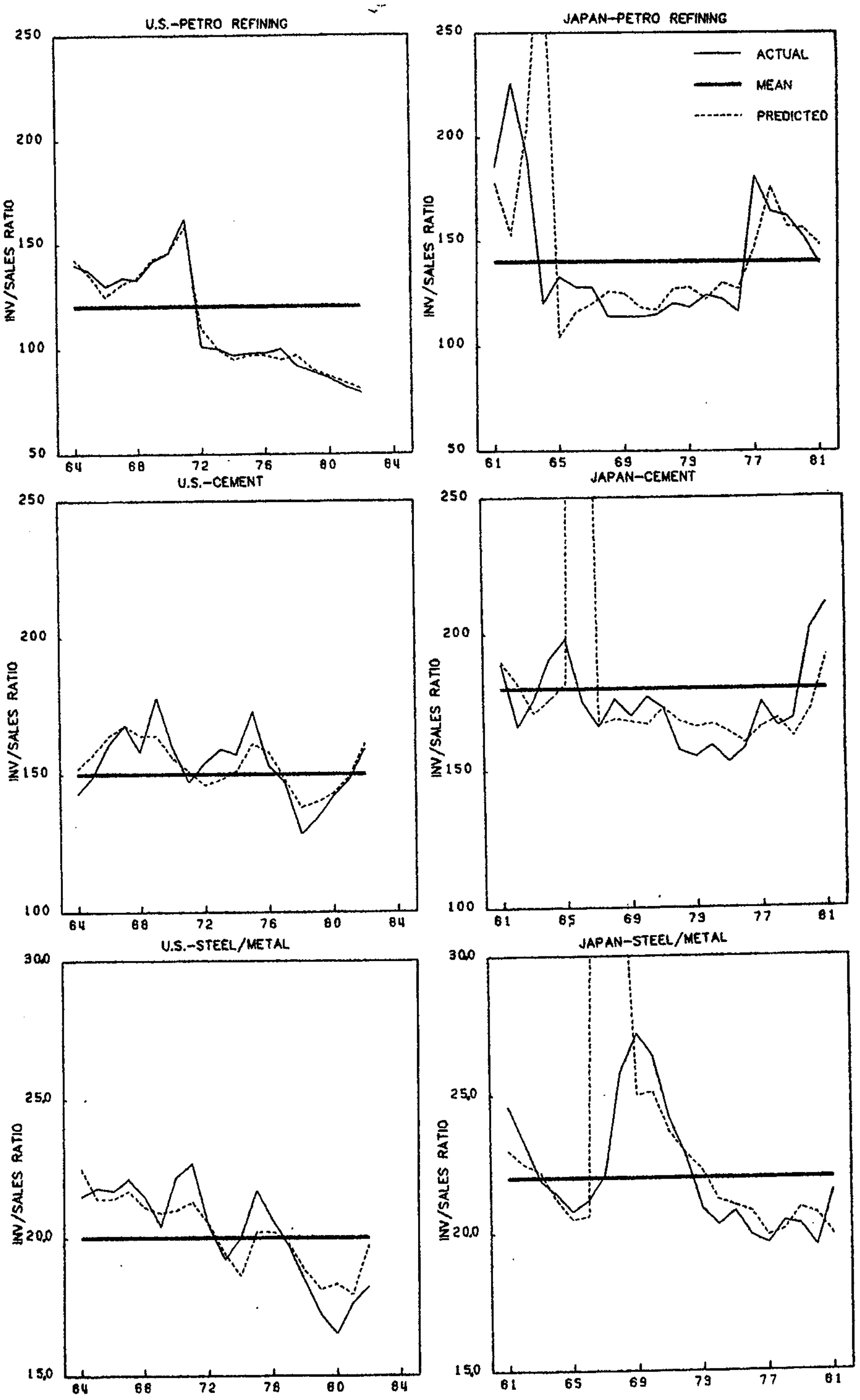

FIG. 3. Graphs for the U.S. and Japanese industries: Petro refining, cement, steel/metal, electrical machinery/computing, and precision/others. The meaning of each graph is exactly the same as for Figs. 1 and 2.

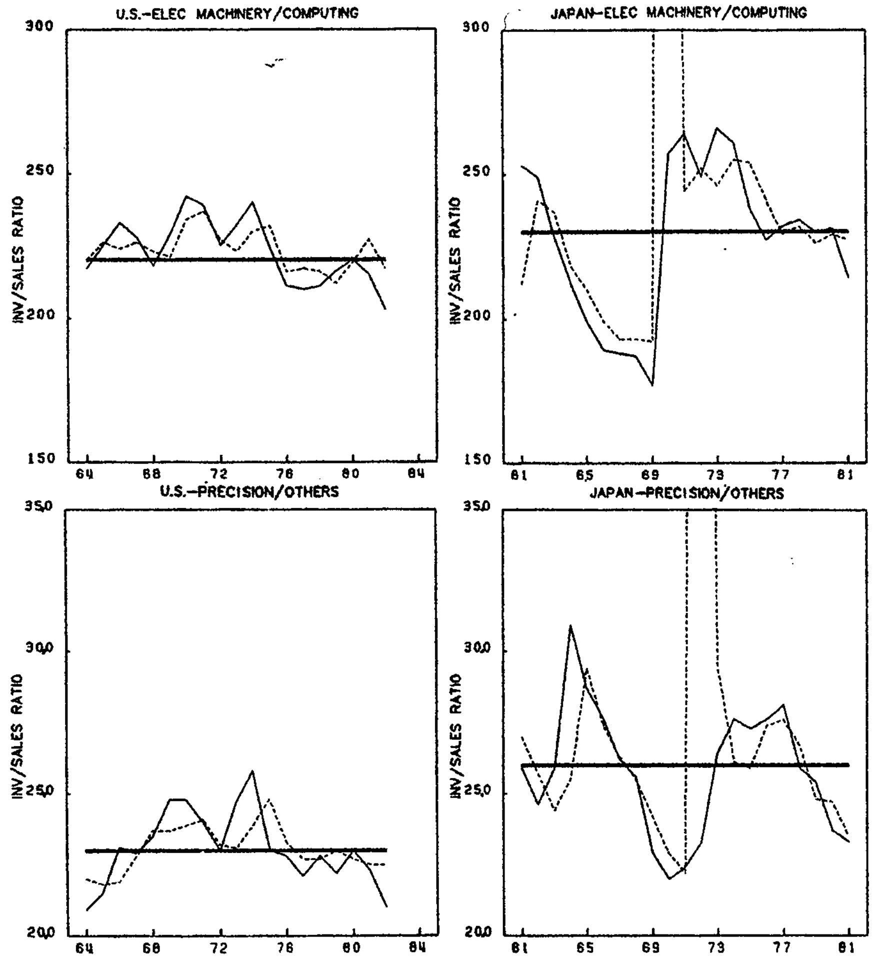

FIG. 3—*Continued*

for all firms listed on the Tokyo and Osaka Stock Exchanges. (All variables are measured in 1960 millions of yen.) Finally the industry titles used in Tables I and II are only suggestive. The details of the firms included in our industry groups are available on request from the authors.

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